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An Analysis of the Relationship between Inflation and Gold Prices: Evidence from Pakistan

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Abstract

In this study, we formulate a new inflation equation to capture the potential effects of gold and stock prices on inflation in Pakistan. We aim to assess the inflation-hedging properties of gold compared to other assets such as real estate, stock exchange securities, and foreign currency holdings. Applying time-series econometric techniques (cointegration and vector error correction models) to data for 1960–2010, we find that gold is a potential determinant of inflation in Pakistan. On the other hand, it also provides a complete hedge against unexpected inflation. Real estate assets are more than a complete hedge against expected inflation, although stock exchange securities outperform gold and real estate as a hedge against unexpected inflation. Foreign currency proves to be an insignificant hedge against inflation. Given the dual nature of the relationship between gold and inflation, it is increasingly important for the government to monitor and regulate the gold market in Pakistan. Moreover, stock market investment should be encouraged by the government given that asset price inflation does not pose a critical problem for Pakistan as yet.

Keywords: Gold prices, inflation hedging, assets, time series econometrics technique, Pakistan.

JEL classification: E31, E37, E4.

1. Introduction

Inflation is a worldwide macroeconomic problem owing to its adverse implications for economic expansion and income redistribution. Achieving a moderate level of inflation is, therefore, one of the main objectives of both developed and developing economies. In this context, a large body of theoretical literature has evolved over time identifying the

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determinants of inflation. There is also extensive research on this issue in Pakistan (see, for example, Rehman, 1994; Akhtar, 1990; Choudhri & Khan, 2002; Hyder & Shah, 2004; Khalid, 2005; Akbari, 2005; Kemal, 2006; Khan & Schimmelpfenning, 2006; Haque & Qayyum, 2006; Khan & Rashid, 2010). However, the dynamic and globally integrated nature of economic has altered previous relationships and led to the emergence of new phenomena affecting inflation. This means that we may require a revised inflation equation.

Gold prices have attracted considerable attention for their potential effect on inflation. Like other assets that are found to predict inflation behavior because their returns embed inflation expectations, gold prices can also serve as a leading indicator of inflation. This argument has been put forward by many researchers based on the failure of some financial assets to predict the behavior of inflation over a longer period of time (see, for instance, Stock & Watson, 1999; Cecchetti, Chu, & Steindel, 2000; Boivin & Ng, 2006; Banerjee & Marcellino, 2006).

There is very little literature on the inflation-predicting properties of gold and other financial assets in Pakistan. Gold prices have not been incorporated into inflation dynamics, nor has the relationship between inflation and asset prices been closely explored. This is because the idea is a relatively new one in the first case, while studies on the latter expect the two variables to be related in the opposite direction. However, with the heavy use of gold in Pakistan[1] and the integration of its financial markets with international markets, it is reasonable to assume that both variables may contain some information about inflation.

Along with the evolution of the determinants of inflation, another key issue concerns hedging against inflation. Various assets are considered to provide a "safe haven" against inflation, e.g., stock exchange securities, real estate, and foreign currency. Gold is also expected to be an inflation-hedging asset, given its long-run association with inflation. However, the literature on the hedging properties of gold and assets in Pakistan is limited. Only one OLS-based study (Nishat & Mustafa, 2008) has been carried out, which rejects the significance of gold and stocks in providing a hedge against inflation. Accordingly, the objectives of this study are:

[1] According to the World Gold Council, Pakistan is the tenth largest consumer of gold, which is used for industrial, ornamental, and investment purposes and also kept in physical bars for saving purposes.

- To examine the relationship between gold and stock prices and inflation in Pakistan
- To examine the inflation-hedging properties of gold and other assets such as real estate, stock exchange securities, and foreign currency.

Using time-series data for the period 1960–2010, we apply cointegration and vector error correction techniques to assess the long-run relationship between inflation and the assets mentioned above. The same techniques are applied to examine the inflation-hedging properties of these assets.

Section 2 provides a literature review, Section 3 describes the methodology used, Section 4 gives a detailed discussion of the results obtained, and Section 5 concludes the analysis and presents some policy recommendations.

2. Literature Review

This section reviews the theoretical and empirical literature.

2.1. Theoretical Viewpoint

The theoretical relationship between asset prices (gold and other stocks) and inflation can be examined through a variety of channels. Marx's theory of money was important in explaining this relationship before the adoption of a fiat money system. The theory postulates that upheavals in the intrinsic value of a money commodity (gold) can cause the general price level to rise. Subsequently, many theoretical arguments were put forward to explain the asset prices-to-inflation transmission mechanism while positing the demand side of the economy as a catalyst. However, numerous arguments have also negated this relationship over the years.

Tobin's q theory proposes that an increase in asset prices makes the installment of new capital more profitable, thus giving impetus to investment growth, which then translates into demand-pull inflation (Mishkin, 1990). The wealth effect of this asset price increase not only leads to an increase in private consumption but also in borrowing capacity and, hence, inflation (Kent & Lowe, 1997). On the other hand, the classical dichotomy between these variables recognizes that asset prices are innately impulsive and highly vulnerable to changes in investor sentiments, quite independent of any change in the basic economic variables. Consequently, extracting correct and timely information from any observed movement in asset prices is extremely difficult (Bernanke & Gertler, 2000).

The literature on inflation hedging establishes that different assets have unique characteristics that provide safe havens against inflation. For instance, Fisher's (1930) hypothesis concerns the role of stock market securities in hedging against inflation through a one-to-one relationship between expected nominal stock returns and the interest rate.[2] Similarly, real estate as an investment and durable consumer good also has inflation-hedging properties. When treated as investment, its market worth represents the present value of corresponding rents on the property. Owners of real estate normally prefer to conserve real rent cash flows, and thus rental properties must provide a hedge as landlords raise the rent to deal with inflation (Hodges, Kneafsey, McFarren, & Whitney, 2011).

The hedging properties of foreign currency can be explained in terms of dollarization. This takes place when the residents of a country use a foreign currency—along with or as a replacement for the domestic currency—as a store of value, medium of exchange, and/or unit of account in the face of uncertainty in the domestic economy. This provides a hedge against inflation. Similarly, gold, which has always been a symbol of wealth, is preferred in times of economic uncertainty arising due to inflation; it thus provides protection against inflation.

2.2. Empirical Viewpoint

The literature on the relationship between gold prices and inflation is very limited because the phenomenon is relatively new. Mahdavi and Zhou (1997) study the role of gold prices and commodity prices in predicting inflation and find some evidence of cointegration between commodity prices and the consumer price index (CPI). However, the relationship between the CPI and gold prices is not found to be significant.

[2] This relationship is given in the form of the Fisher equation, which states that the real interest rate equals the nominal interest rate minus expected inflation. Assuming the real interest rate remains constant, the nominal interest rate (returns on stocks) must change points one-to-one with every change in expected inflation.

Ghosh, Levin, Macmillan, and Wright (2004) test the existence of a stable long-run relationship between the monthly price of gold and inflation in the US from 1945 to 2006 and from 1973 to 2006. Given that both the price of gold and the CPI have been subject to structural changes over time, instead of assuming these structural changes to be exogenous, they apply a unit root test that allows them to estimate the timing of significant breaks. The inclusion of endogenously determined structural breaks provides evidence of a cointegrated relationship between gold and inflation both for the postwar period and since the 1970s.

Ranson (2005) argues that gold is a better indicator of inflation than oil and a sound alternative for investors seeking an asset to hedge against inflation. Dempster and Artigas (2009) also find evidence that supports the existence of a relationship between gold prices and inflation from 1974 to 2008, and show that peaks in the gold price and inflation tend to occur simultaneously.

The empirical evidence confirms that, other than gold, the price of other financial assets also determines inflation. Feldstein (1983) examines the role of expected and unexpected future inflation in stock prices and observes a positive relationship between the former and a negative one between the latter. Kent and Lowe (1997) find that a rapid decline in asset prices, when the intermediary role of the financial sector is undermined, creates a negative output gap and causes deflation in the economy.

Although market reports on inflation hedging often regard gold to be a safe haven, very few academic studies have looked at this issue. Mahdavi and Zhou (1997) study the performance of gold prices and commodity prices in predicting inflation, and find that the role of gold as a hedge against inflation seems to have diminished over time. Ranson (2005) holds that gold, along with silver and platinum, are excellent inflation-immunizing assets. However, investors using gold to hedge against inflation should take into account two important caveats. First, due to one period lead correlation between gold prices and t-bond prices, the investment in gold must be done a year before the expected level of inflation is realized. Second, gold should not be kept in the portfolio during deflationary pressure as the price of gold is leveraged on either side.

Baur and Lucey (2006) acknowledge the role of gold in providing a hedge against losses experienced in the bond and stock markets. Dempster and Artigas (2009) establish that gold has a role to play both as a tactical inflation hedge and as a long-term strategic asset. If the world economy were to experience a revival in inflation, then gold, like other traditional inflation hedges, would likely outperform conventional financial assets. Moreover, it could enhance the investor's risk-controlled returns even in a low- to medium-inflation environment.

The empirical evidence maintains that a range of assets have inflation-protecting characteristics. Real estate is regarded as a strong inflation hedge on the conceptual basis that it actually performs better than other inflation-sensitive assets. Its prices can also change in response to several factors, especially exogenous supply or demand shocks, endogenous cycles within the real estate market, and assigned rents (Case & Wachter, 2011). When real estate markets are in balance, it tends to offer a hedge against expected inflation, but there is little relationship between returns and unexpected inflation (Wurtzebach, Mueller, & Machi, 1991; Hartzell, 2001).

Adrangi, Chatrath, and Christie-David (2000) find a negative relationship between stock returns and inflation rates in industrialized economies. The Johansen and Juselius cointegration tests validate the presence of a long-run equilibrium between the general price level, stock prices, and real economic activity. In addition, stock prices and the general price level demonstrate a strong long-run equilibrium with real economic activity and each other. These findings lend support to Fama's proxy hypothesis[3] for the long run.

Mahmood and Dinniah (2009) investigate the relationship between stock prices and three macroeconomic variables—inflation, exchange rates, and output—for six countries in the Asian-Pacific region. They find that these variables have a negligible effect on the performance of stock market returns, except for Thailand and Hong Kong. Most earlier studies point out that stock prices, real economic activity, and consumer price levels have strong long-run equilibrium relationships (Adrangi, Chatrath, & Sanvicente, 2000). Hondroyiannis and Papapetrou (2001) establish, however, that there is no long-run relationship among these economic variables. There is evidence though, that suggests a negative relationship

[3] Fama's proxy hypothesis states a negative relationship between stock returns and inflation.

between stock returns and both expected and unexpected inflation particularly for the US (Lintner, 1975; Fama & Schwert, 1977; Fama, 1981; Geske & Roll, 1983; Ajayi & Mougoue, 1996; Yu, 1997).

The literature provides a valuable insight into the role of gold as both a commodity and an asset, and shows why it is important to explore its relationship with inflation and compare its hedging properties with those of other assets for Pakistan.

3. Methodology

This section presents a model to determine the relationship between gold prices and inflation, and then describes the methodology used to examine the inflation-hedging properties of different assets.

3.1. Gold Prices and Inflation

Numerous studies have examined the determinants of inflation in Pakistan. Khalid (2005), Khan and Schimmelpfennig (2006), Hossain (1990), and others consider inflation to be a monetary phenomenon determined by money supply, the exchange rate, and inflationary expectations. Other studies have identified the fiscal deficit and total bank borrowing by the government sector, imported inflation, openness, adaptive inflation, and expectations as the determinants of inflation (see Agha & Khan, 2006; Haque & Qayyum, 2006; Khan, Bukhari, & Ahmed, 2007). Bilquees (1988) points to structural factors as the cause of inflation in Pakistan, while Mahdavi and Zhou (1997) find that gold and asset prices are leading indicators of the inflation rate.

Tkacz (2007) has developed a model that posits the rate of return on gold and the exchange rate as the major determinants of inflation. The basic framework we use to examine the relationship between gold prices, stock prices, and inflation is adapted from this model, which is constructed as follows:

$$\pi_t = R_t - r_t \qquad (1)$$

π is the inflation rate, which is calculated as follows:

$$\pi_t = [LogP_t - LogP_{t-1}]$$

where P_t is the CPI, R_t refers to the nominal annualized rate of return on gold, and r_t is the real annualized rate of return.

As the price of gold determined in world markets is usually expressed in US dollars, it is multiplied by the prevailing exchange rate to compute its rate of return in local currency units. Henceforth, we denote the domestic rate of return on gold by R_D while R_t is the US dollar rate of return on gold. *G* denotes the price of gold (per ounce) in US dollars and *E* is the domestic currency against the US dollar exchange rate. The international (US) and domestic annualized rates of return on gold are, respectively, given by:

$$R_t = [LogG_t - LogG_{t-1}] \tag{2}$$

$$R_t^D = [Log(G_t * E_t) - Log(G_{t-1} * E_{t-1})] \tag{3}$$

$$R^D = R_t + \dot{E}_t \tag{4}$$

where EY_t is the annualized percentage change in the exchange rate. The following equation shows whether the return on gold contains information on realized inflation:

$$\pi_t = \alpha + \beta R_{t-1} + \epsilon_t \tag{5}$$

$$\pi_t^D = \alpha + \beta R_{t-1}^D + \epsilon_t \tag{6}$$

$$\pi_t = \alpha_0 + \beta_1 R_t + \beta_2 ER_t + \mu_t \tag{7}$$

where R_t is the return on gold and ER_t is the rate of change in the exchange rate.

Next, we modify the above model to achieve the objectives of this study. Instead of taking the return on gold, which is calculated as the difference between gold prices in two successive periods, we incorporate gold prices as a direct determinant of inflation. For the purpose of comparison between gold and stocks to explain inflation, stock prices are also incorporated in the equation. The final form of the model is:

$$inf_t = \alpha_0 + \alpha_1 G_t + \alpha_2 GI_t + \alpha_3 ER_t + \epsilon_t \tag{8}$$

The variables are given in natural logarithmic form where *G* is the gold price, *GI* is the general index of stock prices, and *ER* is the exchange rate.

It is worth mentioning that oil prices are included in the equation as an exogenous variable. The large price hike witnessed in the 1970s was triggered by the increase in oil prices. Similarly, the decline in inflation in the 1980s and 1990s followed a decrease in oil prices. Therefore, there is enough historical proof of the relationship between oil prices and inflation to support the inclusion of oil prices in our model. Oil-importing countries such as Pakistan are highly sensitive to fluctuations in oil prices, given that oil is a major input for its industries. However, it has no influence over the supply of oil, given that the Middle East oil cartels are able to affect prices. Additionally, its demand for oil cannot influence oil prices (unlike China, where an increase in the consumption of oil is likely to lead to an oil price rise). Therefore, oil is taken as an exogenous variable.

3.2. Theoretical Justification of Variables

We incorporate stock prices measured by the general share price index as a potential determinant of inflation in Pakistan. Financial asset prices are also useful leading indicators of inflation as their rates of return incorporate inflation expectations. Moreover, through their effect on aggregate demand, they are expected to have a positive relationship with inflation. We also expect a positive relationship between the exchange rate and inflation based on the exchange rate pass-through, i.e., the percentage change in import prices in local currency as a result of a one-percentage point change in the exchange rate between the exporting and importing countries (Jaffri, 2010). Given this strong theoretical relationship, the exchange rate becomes a pertinent variable in the inflation equation.

Gold serves as a good store of value that is expected to affect the inflation rate positively. We include gold prices in our analysis because they contain information about the expected rate of inflation and are determined in complete information financial markets where agents are assumed to form rational expectations. This is not the case with other commodities (Mahdavi & Zhou, 1997). The price of gold is a leading indicator of inflation because gold is usually retained as a store of value owing to its durability and attractiveness (Ghosh et al., 2004).

3.3. *Assets as an Inflation Hedge*

Capie, Mills, and Wood (2004), Levin and Wright (2006), Khan et al. (2007), and Worthington and Pahlavani (2007) find that gold has inflation-hedging properties, mainly for the US economy. Nishat and Mustafa's (2008) study explores which assets serve as an inflation hedge in Pakistan; we adapt their methodology for this part of the research.

First, total inflation is decomposed into expected inflation and unexpected inflation. This step is essential for disaggregating different assets on the basis of their explanatory power for expected and unexpected inflation. Expected inflation is the return on risk-free assets while unexpected inflation is taken as the difference between total observed inflation and expected inflation. Mathematically, this is represented by

$$inf_t = \beta_0 + \beta_1 RD_t + \sigma_t \tag{9}$$

where $\inf_t$ is inflation and RD_t is the return on risk-free assets.

Next, we specify the inflation-hedging properties of different assets by regressing the return on specific assets on both expected and unexpected inflation. This is represented by

$$R_t = \gamma_0 + \gamma_1 I_E + \gamma_2 I_U + \epsilon_t \tag{10}$$

where R_t is the total return on assets, I_E is expected inflation, and I_U is unexpected inflation.

The total return on assets is calculated as follows:

$$R_t = Rate\ of\ return + capital\ gain\ (or\ loss)$$

where the capital gain or loss is measured by the rate of change of the price of a particular asset.

Whether an asset serves as protection against inflation depends on the sign and magnitude of γ_1 and γ_2. The following situations can arise:

1. γ_1 and γ_2 are both equal to 1. In this case, the asset provides a complete hedge against both expected and unexpected inflation.

2. γ_1 and γ_2 are both greater than 1. In this case, the asset is an over-hedge against both expected and unexpected inflation.
3. γ_1 and γ_2 are both less than 1. In this case, the asset is a partial hedge against both expected and unexpected inflation.
4. γ_1 is 1 and γ_2 is less than 1. In this case, the asset is a partial hedge against unexpected inflation but a complete hedge against expected inflation.
5. γ_1 is less than 1 and γ_2 is 1. In this case, the asset is a partial hedge against expected inflation but a complete hedge against unexpected inflation.

We study the inflation-hedging properties of four assets. On the basis of equation (10), the following four equations are formulated for the final regression:

$$Tg_t = p_0 + p_1 Ie_t + p_2 Iu_t + \epsilon_t \tag{11}$$

$$Tr_t = \sigma_0 + \sigma_1 Ie_t + \sigma_2 Iu_t + \epsilon_t \tag{12}$$

$$Ts_t = \tau_0 + \tau_1 Ie_t + \tau_2 Iu_t + \epsilon_t \tag{13}$$

$$Fc_t = \gamma_0 + \gamma_1 Ie_t + \gamma_2 Iu_t + \epsilon_t \tag{14}$$

where *Tg* is the total return on gold, *Tr* is the total return on real estate, *Ts* is the total return on stock exchange securities, and *Fc* is foreign currency.

Oil prices and the exchange rate are taken as exogenous variables in all four models. Both variables can dampen (increase) the inflation-hedging properties of different assets. For instance, high oil prices may hamper the functioning of the stock market and the country's economic activity as a whole.

The same theory applies to gold prices. Oil prices can interact with gold prices in the following ways. A high oil price can drag down share prices and change the inflation-hedging abilities of stock returns. In response to the capital loss on stocks, investors will look for alternative assets, such as gold. Thus, the oil price can indirectly affect the price of gold and reduce its ability to protect against inflation. The exchange rate can also affect prices and returns on different assets by changing the relative strength of a particular currency. However, since there is no evidence that these variables are affected by the inflation rate in Pakistan,[4] we take oil prices and the exchange rate as exogenous variables.

3.4. Theoretical Justification of Variables

If gold is a perfect internal hedge, its nominal price and domestic inflation will rise at the same time. For it to be an external hedge, the magnitude and time of price change has to be perfectly aligned with the change in the exchange rate but in the opposite direction. This implies that one can protect against exchange rate fluctuations by investing in gold. Many empirical studies suggest that direct and indirect gold investment serves as an effective inflationary hedge.

Real estate was considered a separate asset class by institutional investors in the 1990s. Gordon's (1962) growth model suggests that real estate can also protect investors against inflation as both its value and income generated adjust to inflation. We expect a short-run positive and long-run negative relationship between inflation and real estate.

Theoretically, stock returns appear to act as a good hedge against inflation both in the long and short run. The relationship between stock returns and inflation suggests that investment in equity markets can act as a good hedge against inflation if firms' revenues and earnings grow over time. There is ample evidence in favor of a positive relation between inflation and stock returns, supporting a generalized Fisher hypothesis. Moreover, the black economy raises stock market returns, which is significant in the long and short run. This can be explained as the effort of black money holders to 'clean' their assets by channeling them through the stock market.

The importance of foreign currency holdings as a hedge against inflation emerged after the final breakdown of the Bretton Woods system in 1971 when the exchange rates of the major currencies varied persistently against each other. Investors diversified their portfolios by holding

[4] Inflation in Pakistan does not translate into higher world oil prices (Section 3.1). The same argument applies to the effect of inflation on the exchange rate, as the value of the dollar is determined in international markets.

different currencies to seek profit and to protect themselves against the risk arising from currency fluctuations. These currencies are known as "safe havens" and their inflation-hedging properties are worth examining.

4. Estimation and Discussion of Results

This section examines the relationship between gold prices and inflation and the inflation-hedging properties of different assets.

4.1. Data Sources

We have used annual data for 1960 to 2010 retrieved from the *Handbook of Statistics on Pakistan Economy* (2010) published by the State Bank of Pakistan. The CPI uses 2000/01 as the base year, the exchange rate is given in Pakistani rupees per US dollar (annual average), the weighted average rate of return on precious metals is used as a proxy for the rate of return on gold, and stock exchange securities and real estate are given in percent per annum. Gold prices are given in rupees per 10 grams and the general index of stock prices is used to represent stock prices.

4.2. Descriptive Statistics

Gold prices are taken in growth rate form while the exchange rate, inflation rate, and returns on different assets are taken as absolute values. For each decade, we calculate the average of each variable to determine its average behavior. The standard deviation for each decade is also calculated to capture the extent of fluctuation in each variable. Table 1 presents the behavior of all the variables used.

Table 1: Descriptive analysis of variables

Variable	1960–69	1970–79	1980–89	1990–99	2000–09
Gold prices					
Average	15.18	11.23	38.98	7.97	15.54
Standard deviation	17.18	34.27	62.30	7.17	24.60
Inflation					
Average	1.58	5.55	10.56	5.88	12.17
Standard deviation	0.97	2.18	0.23	0.22	12.14
Exchange rate					
Average	4.76	7.33	15.50	34.02	61.09
Standard deviation	0.00	3.63	7.91	17.30	5.22
Return on gold					
Average	7.88	8.67	8.51	10.74	12.51
Standard deviation	1.59	0.13	0.84	0.85	34.16
Return on securities					
Average	6.62	10.44	10.50	10.75	12.51
Standard deviation	1.97	3.06	1.97	0.85	34.16
Return on real estate					
Average	6.25	9.395	10.67	12.00	13.20
Standard deviation	0.66	1.97	0.09	2.53	32.49

Source: State Bank of Pakistan (2010).

Gold prices have shown an increasing trend over the last five decades. The average value increased continuously and peaked in the 1980s, which can be explained by the Iran crisis and the invasion of Afghanistan by the Soviet Union. In 1982, China allowed its citizens to own gold, which led to record gold prices after the previous 28 years. The volatility experienced during this period was 38.98—the highest point in this time. This also led to higher returns (the standard deviation is 62.30), implying higher risk. Since 2001, the price of gold has risen steadily. This surge has an obvious association with the expansion of US national debt and the weakening of the US dollar relative to other currencies. In 2005, gold reached PRs 8,216 for the first time and by 2008, the rate was more than PRs 15,000 per 10 grams. The financial crisis also fueled the demand for gold and exchange-traded funds.

Recently, gold prices have set a new record of PRs 29,587 per 10 grams, the reasons for which are rooted in the US uncertainty about a sustainable economic recovery, increasing inflation, possible corporate insolvency and default of corporate bonds, growing national debt, low interest rates, and the expansion of the money supply. The decrease in gold production by 10 percent since 2001 and high demand for jewelry as well as higher demand for gold and by institutional investors are other factors that have pushed up the value of gold ("History of gold," 2011). The average growth in price was 15.54 percent with a much higher volatility (24.60) compared to the previous decade (7.17). This also shows that the return on gold was much higher during this decade (Bhose, 2012).

Inflation has followed a mixed trend for the last five decades, rising during 1960–90, but remaining low in the 1990s. The reasons for this include the slow pace of economic growth, output delays, expansionary monetary policies, heavy duties and taxes, a depreciating rupee, and frequent adjustments in the administered prices of gas, electricity, petroleum oil and lubricants, and gas products. An improved supply position, strict budgetary measures, and depressed international market prices led to a meltdown in the inflation rate in the late 1990s (Rehman, 2010).

Inflation showed much greater volatility during the 2000s than in the previous decades. The inflation rate fell to 3.1 percent in 2002/03 (from 5.7 percent in 1998/99)—the lowest in the last three decades. This low level of inflation was an outcome of strict fiscal discipline, output recovery, a reduction in duties and taxes, the low monetization of the budget deficit, and the appreciation of the exchange rate. 2003/04 witnessed a boost in the inflation rate to 9.3 percent due to an increase in the support price of wheat, wheat shortages, and a rise in international prices, specifically oil prices. The inflation rate then dropped to 7.9 percent at the end of 2005/06.

Subsequently, despite a tight monetary policy and the resolution of supply blockages, nonfood-nonenergy inflation (core inflation) remained high at 7.5 percent in 2007/08 on account of rising house rents and other factors. Various domestic and international factors contributed to the astonishing surge in domestic price levels in Pakistan to an average of 12.17 percent—the highest point in this period.

Even with an overall slowdown in world economic growth, crude oil prices have continued to grow. Oil prices increased from USD 55 per barrel in January 2007 to over USD 130 per barrel in May 2008—a surge of more than 145 percent. The dollar depreciated against major currencies throughout the year along with rising oil prices. This correlation between oil prices and the dollar caused Pakistan to face huge import bills, driving up inflation (Haque & Qayyum, 2006).

4.3. Estimation

In this section, we estimate a model revealing the relationship between gold prices and inflation, and then examine empirically the inflation-hedging properties of different assets.

4.3.1. Gold Prices and Inflation

The long-run relationship between gold prices and inflation for the period 1960–2010 is determined based on Johansen (1998) and Johansen and Juselius (1990). These studies use cointegration to determine the long-run relationship between two or more variables that are individually nonstationary but have a stationary linear combination. Two prerequisites must be met in order to obtain reliable results: (i) cointegration requires all the variables to be integrated of the same order, and (ii) the appropriate lag length must be determined.

Time-series data is usually nonstationary and thus subject to spurious regression. To carry out a cointegration analysis, the first step is to check the stationary properties of the data. We conduct an augmented Dickey-Fuller unit root test to check the order of integration for each variable (see Table 2). The results show that all the series are integrated of order one, I(1), and we can proceed with our cointegration analysis.

Table 2: Results of ADF test

Variable	Level t-statistic	1st difference t-statistic	Order of integration
inf_t	2.41	-3.63*	I(1)
G_t	3.07	-3.57*	I(1)
GI_t	0.69	-3.17*	I(1)
ER_t	2.94	-6.13*	I(1)

Note: * = significant at 1%.

The second step is to determine the optimal lag length based either on the Akaike information criterion or the Schwarz information criterion (SIC). Table 3 gives the results for the SIC. The appropriate lag length selected for this model is 1.

Table 3: Lag length selection

Lag	SIC
0	3.62
1	-4.45*
2	-4.27
3	-3.80

Note: * = minimum value of SIC.

Johansen's cointegration technique proposes two tests—the trace test (λ_{trace}) and maximum eigenvalue test (λ_{max})—which are used to determine the existence and number of cointegrating vectors. In the case of the trace test, the null hypothesis is that the number of cointegrating vectors is less than or equal to *r* where *r* = 0, 1, 2, 3…, etc. In each case, the null hypothesis is tested against the general hypothesis (full rank *r* = *n*). In the case of the maximum eigenvalue test, the null hypothesis of the existence of *r* cointegrating vectors is tested against the alternative of *r* + 1 cointegrating vectors (Mukhtar & Ilyas, 2009). The multivariate cointegration test is expressed as:

$$Z_t = K_1 Z_{t-1} + K_2 Z_{t-2} + \cdots + K_{k-1} Z_{t-k} + u + v_t$$

where $Z_t(\inf_t, G_t, GI_t, ER_t)$ is a 4 x 1 vector of variables that are integrated of order one, i.e., I(1), μ is a vector of the constant term, and υ_t is a vector of normally and independently distributed error terms (see Mukhtar, 2010). The results in Table 4 indicate the existence of one cointegrated vector among the selected time series.

Table 4: Cointegration test (Johansen's maximum likelihood method) for equation (8)

Equation (8): $\inf_t = \alpha_0 + \alpha_1 G_t + \alpha_2 GI_t + \alpha_3 ER_t + \varepsilon_t$

Null hypothesis	Alternative hypothesis			Critical values	P-values*
λ trace rank tests		Eigenvalues	λ trace rank value	(0.05 %)	
H0: $r = 0$	H1: $r = 0$	0.527	63.88	54.07*	0.005
H0: $r = 1$	H1: $r = 1$	0.416	34.66	35.19	0.060
H0: $r = 2$	H1: $r = 2$	0.193	13.66	20.26	0.310
H0: $r = 3$	H0: $r = 3$	0.126	5.26	9.14	0.250
λ max. rank tests			λ max. eigenvalue		
H0: $r = 0$	H1: $r > 0$	0.527	29.22	28.58*	0.040
H0: $r \leq 1$	H1: $r > 1$	0.416	21.00	22.29	0.070
H0: $r \leq 2$	H1: $r > 2$	0.193	8.39	15.89	0.500
H0: $r \leq 3$	H1: $r > 3$	0.126	5.26	9.16	0.250

Note: * = MacKinnon-Haug-Michelis p-values. Trace test indicates one cointegrated equation at 0.05 level. Max. eigenvalue test indicates one cointegrated equation at 0.05 level.

A vector error correction model (VECM) is a general framework used to depict the dynamic relationship among variables that are stationary in their differences, i.e., I(1). We use the VECM because the time series are not stationary at level but are stationary in differences, and the variables are cointegrated, which means that they are likely to be engaged in a long-run relationship.

Adding error correction features to a multifactor model can tell us how much the error in the variables is corrected each year. The equation generated by the VECM takes the following form:

$$\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + \Gamma_2 \Delta Z_{t-2} + \ldots + \Gamma_{k-1} \Delta Z_{t-k-1} + \Pi Z_{t-1} + u + \upsilon_t$$

where $\Gamma_i = (I - A_1 - A_2 \ldots - A_i)$, i=1,2,3 ... k–1, and $\Pi = -(I - A_1 - A_2 - A_3 \ldots A_k)$. The coefficient matrix Π provides information on the long-run relationships among the variables in the data. Π can be factored into $\alpha\beta$ where α gives the speed of adjustment to the equilibrium coefficients while β is the long-run matrix of coefficients. The presence of r cointegrating vectors between the elements of Z implies that Π is of the rank r ($0 < r < 3$) (Mukhtar, 2010). The results of the VECM are presented in Table 5.

Table 5: Results of VECM for equation (8)

Variable	Coefficient	t-statistic
C	0.91*	4.56
G_t	0.33*	6.70
GI_t	0.02	0.31
ER_t	0.73*	9.80
ECT	-0.20*	-4.51

Note: * = significant at 1%.

The results reveal a positive and significant relationship between gold prices and inflation, where a one-percent increase in gold prices causes a 0.33 percent increase in inflation. These results are consistent with those of Levin and Wright (2006) and Ghosh et al. (2004), both of who find a long-term relationship between gold prices and inflation in the US. Our results also show a positive but insignificant relationship between share prices and inflation, implying that asset price inflation is thus far not a crucial issue in Pakistan. Moreover, the results verify the insignificant intermediary role of the capital market in Pakistan.

Our results also indicate that a one-percent increase in the exchange rate causes a 0.73 percent increase in inflation, and the variable is significant at 1 percent. These results correspond to numerous empirical studies that have found a long-run relationship between inflation and the exchange rate (see, for example, Choudhri & Khan, 2002; Honohan & Lane, 2003; Capie et al., 2004; Khalid, 2005; Khan & Schimmelpfenning, 2006; Dlamini et al., 2011; Laryea & Sumaila, 2011).

The reason for this positive relationship depends on a number of factors. For instance, a decrease in the exchange rate makes imported goods and services more expensive in the resident country. Producers then shift the higher cost of imported items and raw materials onto consumers in the form of high prices, thus causing cost-push inflation. The extent to which a depreciation of the domestic currency causes inflation depends on producers' degree of dependency on their imported components and on their willingness to forward higher costs to consumers.

The time path of cointegrated variables is affected by any deviation in the long-run equilibrium. The error correction term represents the percentage of correlation with any deviation in the long-run equilibrium and the deviation's speed of correction in the long run. The error correction term for overall inflation is negative and significant with a 0.2 percent speed of convergence towards equilibrium.

4.3.2. Inflation Hedging

Johansen (1998) and Johansen and Juselius's (1990) cointegration technique will now be used to assess the inflation-hedging properties of different assets, with particular focus on gold. As mentioned earlier, there are two prerequisites for applying this technique: (i) checking the stationary properties of the data and (ii) selecting an appropriate lag length. The ADF unit root test is performed to check the order of integration of each variable. Table 6 shows that all the variables are integrated of order one.

Table 6: Results of ADF test

	Level	1st difference	
Variable	**t- statistic**	**t- statistic**	**Order of integration**
Tg_t	0.55	-4.12*	I(1)
Tr_t	0.73	-8.70*	I(1)
Ts_t	0.48	-11.29*	I(1)
Fc_t	0.88	-5.78*	I(1)
Ie_t	-0.76	-4.62*	I(1)
Iu_t	-1.37	-2.06**	I(1)

Note: * = significant at 1%.

The second step is to determine the optimal lag length. We use the SIC, the results for which are given in Table 7. The appropriate lag length selected for equations (11), (12), (13), and (14) is 1.

Table 7: VAR lag order selection criteria

	Equation (11)	Equation (12)	Equation (13)	Equation (14)
Lags	**SIC**	**SIC**	**SIC**	**SIC**
0	14.28	13.29	13.91	27.79
1	14.05*	12.90*	13.79*	27.61*
2	14.54	13.00	13.93	27.65

Note: * = minimum value of SIC.

The trace test (λ_{trace}) and maximum eigenvalue test (λ_{max}) are used to determine the existence and number of cointegrating vectors. The results for all four equations are reported in Tables 8 to 11. The tables indicate the existence of one cointegrating vector for all the equations except equation (13), for which there are two cointegrating vectors.

Table 8: Cointegration test (Johansen's maximum likelihood method) for equation (11)

Equation (11): $Tg_t = p_0 + p_1 Ie_t + p_2 Iu_t + \varepsilon_t$

Null hypothesis	Alternative hypothesis			Critical values	P-values*
λ trace rank tests		Eigenvalues	λ trace rank value	0.05%	
H0: $r = 0$	H1: $r = 0$	0.48	52.13	42.91*	0.004
H0: $r = 1$	H1: $r = 1$	0.29	25.70	25.87	0.060
H0: $r = 2$	H1: $r = 2$	0.25	11.75	12.51	0.060
λ max. rank tests			λ max. eigenvalue		
H0: $r = 0$	H1: $r > 0$	0.48	26.42	25.82*	0.040
H0: $r \leq 1$	H1: $r > 1$	0.29	13.95	19.38	0.250
H0: $r \leq 2$	H1: $r > 2$	0.25	11.75	12.51	0.060

Note: * = MacKinnon-Haug-Michelis p-values. Trace test indicates one cointegrating equation at the 0.05 level. Max. eigenvalue test indicates one cointegrating equation at the 0.05 level.

Table 9: Cointegration test (Johansen's maximum likelihood method) for equation (12)

Equation (12): $Tr_t = \sigma_0 + \sigma_1 Ie_t + \sigma_2 Iu_t + \varepsilon_t$

Null hypothesis	Alternative hypothesis			Critical values	P-values*
λ trace rank tests		Eigenvalues	λ trace rank value	0.05%	
H0: $r = 0$	H1: $r = 0$	0.49	40.07	24.27*	0.0002
H0: $r = 1$	H1: $r = 1$	0.29	13.50	12.32*	0.031
H0: $r = 2$	H1: $r = 2$	0.0005	0.02	4.12	0.906
λ max. rank tests			λ max. eigenvalue		
H0: $r = 0$	H1: $r > 0$	0.49	26.57	17.79*	0.0019
H0: $r \leq 1$	H1: $r > 1$	0.29	13.48	11.22*	0.019
H0: $r \leq 2$	H1: $r > 2$	0.0005	0.02	4.12	0.90

* Note: * = MacKinnon-Haug-Michelis p-values. Trace test indicates one cointegrating equation at the 0.05 level. Max. eigenvalue test indicates one cointegrating equation at the 0.05 level.

Table 10: Cointegration test (Johansen's maximum likelihood method) for equation (13)

Null hypothesis	Alternative hypothesis			Critical values	P-values*
λ trace rank tests		Eigenvalues	λ trace rank value	0.05%	
H0: $r = 0$	H1: $r = 0$	0.49	45.57	42.91*	0.02
H0: $r = 1$	H1: $r = 1$	0.26	19.15	25.87	0.27
H0: $r = 2$	H1: $r = 2$	0.16	7.14	12.51	0.32
λ max. rank tests			λ max. eigenvalue		
H0: $r = 0$	H1: $r > 0$	0.49	26.41	25.82*	0.01
H0: $r \leq 1$	H1: $r > 1$	0.26	12.01	19.38	0.41
H0: $r \leq 2$	H1: $r > 2$	0.16	7.14	12.51	0.32

Note: * = MacKinnon-Haug-Michelis p-values. Trace test indicates one cointegrating equation at the 0.05 level. Max. eigenvalue test indicates one cointegrating equation at the 0.05 level.

Table 11: Cointegration test (Johansen's maximum likelihood method) for equation (14)

Equation (14): $Fc_t = \gamma_0 + \gamma_1 Ie_t + \gamma_2 Iu_t + \varepsilon_t$

Null hypothesis	Alternative hypothesis			Critical values	P-values*
λ trace rank tests		Eigenvalues	λ trace rank value	0.05%	
H0: $r = 0$	H1: $r = 0$	0.52	30.82	29.79*	0.03
H0: $r = 1$	H1: $r = 1$	0.34	11.44	15.49	0.18
H0: $r = 2$	H1: $r = 2$	0.02	0.64	3.84	0.42
λ max. rank tests			λ max. eigenvalue		
H0: $r = 0$	H1: $r > 0$	0.52	19.37	21.13*	0.08
H0: $r \leq 1$	H1: $r > 1$	0.34	10.80	14.26	0.16
H0: $r \leq 2$	H1: $r > 2$	0.02	0.64	3.84	0.42

Note: * = MacKinnon-Haug-Michelis p-values. Trace test indicates one cointegrating equation at the 0.05 level. Max. eigenvalue test indicates one cointegrating equation at the 0.05 level.

An asset serves as a complete hedge against expected or unexpected inflation if the Fisher coefficient (i.e., γ_1 and γ_2) in a regression of asset returns on expected and unexpected inflation is not statistically different from unity. The asset is an inflexible hedge for $\gamma < 0$; for $0 < \gamma < 1$, it is an incomplete hedge and a more-than-perfect hedge in case $\gamma > 1$.

The results for the VECM are given in Table 12.

Table 12: Results of VECM

	Eq. (11)		Eq. (12) 1st cointeg. eq.		Eq. (12) 2nd cointeg. eq.		Eq. (13)		Eq. (14)	
Variable	**Coeff**	**t-stat**	**Coeff**	**t-stat**	**Coeff**	**t-stat**	**Coeff**	**t-stat**	**Coeff.**	**t-stat**
I_E	-0.220	-0.29	1.26*	3.13	3.98		-4.28	-1.39	144.12	0.39
I_U	1.06*	5.24	-0.45*	-3.52	1.23*	2.91	4.09*	5.55	106.84	0.54
ECT	-0.221*	-3.23	-0.24*	-4.07	-0.24*	-4.07	-0.03***	-1.94	-1.17*	-4.3

Note: * = significant at 1%, ** = significant at 5%, *** = significant at 10%.

In investigating the hedging properties of different assets, we have focused on the return on gold. The results reveal a negative and insignificant relationship between the return on gold and expected inflation, and a significant and positive relationship with unexpected inflation. The Fisher coefficient for unexpected inflation is almost 1 (γ_2 = 1.06), implying that gold provides a complete hedge against unexpected inflation.

The Mercantile Exchange gives the following reasons for the tendency to invest in gold in Pakistan. Gold is highly liquid, and the historical evidence confirms that investing in gold is, universally, an ideal means of hedging uncertainty.[5] It allows portfolio diversification and risk mitigation, and is not affected by political or social stability because it never loses its intrinsic value. The estimate for the error correction term is –0.22 and significant at 1 percent, indicating a quick reversion toward the long-run relationship following a deviation.

The results for real estate are entirely different from those for gold: the first cointegrating equation shows that it provides a more-than-complete hedge against expected inflation with a significant Fisher coefficient equal to 1.26. The second cointegration equation reveals that real estate provides a more-than-complete hedge against both expected and unexpected inflation.

The estimate for the error correction term is –0.24, which shows a reversion toward the long-run relationship following a deviation. The rapid increase in real estate assets may be due to the global real estate bubble. These take longer to deflate and prices fall more slowly because the real estate market is less liquid. Investors then perceive that real estate assets will yield higher returns in the future. Moreover, in Pakistan, the income earned from a real estate investment trust is exempted from income tax and at least 90 percent of its income is distributed among the unit holders of the trust. This implies that people find it safe to invest in real estate assets, which is why their performance as an inflation hedge is noteworthy.

In the case of stock exchange securities, the results indicate a negative and insignificant relationship with expected inflation. The Fisher coefficient (4.09) shows that the return on stock exchange securities is not a hedge against expected inflation. Fama and Schwert (1977) suggest that

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[5] The Pakistan Mercantile Exchange is involved in the exchange of goods, specifically gold, silver, cotton, wheat, etc.

there is no proper explanation for this negative relationship but they do cite the possibility of unidentified phenomena or market inefficiency in delivering the available information on future stock prices. Our results support the proxy hypothesis, which posits that the negative relationship between stock returns and inflation reflects the harmful effects of inflation on real economic activity. However, securities provide a significantly more-than-complete hedge against unexpected inflation in Pakistan. The estimate for the error correction term is –0.03, thus implying a slow reversion to long-run equilibrium.

The performance of foreign currency as an inflation hedge in Pakistan is surprising. The magnitude of the Fisher coefficient for both expected and unexpected inflation are implausibly large: γ_1 =144.12 and γ_2 = 106.84, establishing that foreign exchange is a complete and outstanding inflation hedge although the results are insignificant. The hedging property of foreign currency is supported by the extent of dollarization, which is indicated by foreign currency deposits (FCDs) in the domestic country. Pakistan has faced a decline in FCDs for a number of reasons, including low exports, high imports, foreign debt, and low FDI. This large but insignificant role of foreign currency as an inflation hedge can be explained by the fact that currency swapping in Pakistan also occurs in informal markets, not just through banks or the central bank. Such currency swaps are usually not reported, thus undermining the importance of foreign currency as an inflation hedge. The insignificance can also be explained by the fact that holding foreign currency as an inflation hedge is a relatively new concept.

The results with reference to our focus variable, gold, are important. Some important and surprising findings emerge for stock exchange securities and foreign currency, which show that securities are potentially a crucial indicator of the capital market as they provide an over-hedge against unexpected inflation. The foreign exchange market in Pakistan, however, needs to be organized such that currency swaps and the inflation-hedging property of foreign currency are visible officially.

5. Conclusion and Recommendations

If not stabilized at a moderate level, inflation can have severe consequences for the economy. While an abundant body of literature exists on the determinants of inflation, this study contributes to the literature by investigating the role of gold prices and share prices in determining inflation. It also examines the inflation-hedging properties of gold and other assets.

We use time series data for 1960–2010 and apply cointegration and VECMs to assess the determinants of inflation and the inflation-hedging properties of different assets. The results reveal that gold prices are positively and significantly related to inflation in Pakistan. Share prices have a positive but insignificant relationship with inflation.

As gold prices are positively related to inflation, the study widens its analysis to examine the hedging properties of assets in the capital market against expected and unexpected inflation. The total returns on gold, stock exchange securities, real estate, and foreign currency are taken as dependent variables. The results suggest that all these assets, except foreign currency, provide a hedge against inflation (either expected and/or unexpected). Gold provides a complete hedge against unexpected inflation but not against expected inflation. Real estate provides a significantly more-than-complete hedge against expected inflation, but not against unexpected inflation where the relationship is negative and insignificant.

Stock exchange securities appear to potentially outperform other assets in providing a hedge against unexpected inflation. The results indicate a negative but insignificant relationship with expected inflation. The results for foreign currency imply a positive but insignificant relationship with both expected and unexpected inflation. The insignificant role of foreign currency as an inflation hedge can be explained by the fact that currency swapping in Pakistan also occurs through informal markets and not just through banks. Such currency swaps are not reported but their contribution is enormous.

Based on these findings, we present some recommendations:

- The study establishes that gold is a complete hedge against unexpected inflation in Pakistan. Therefore, speculative behavior regarding gold prices should be observed carefully since it could undermine the hedging properties of gold and lead to inflation. Gold prices are determined in organized markets internationally. Given that Pakistan is one of the largest consumers of gold, its gold market should be organized to eliminate the consequences of speculation.

- Given the insignificant relationship between share prices and inflation, investors should potentially be advised to invest in stocks rather than gold. The insignificant relationship implies that their expectations regarding share prices will not translate into inflation. In this way, stock investment can potentially increase without having any inflationary effect on the economy.

- The coefficient of foreign currency for both expected and unexpected inflation is insignificant but its magnitude is large, indicating that foreign exchange can potentially be a complete hedge. The coefficient is insignificant because most currency swaps take place in the informal market. Foreign remittances increased dramatically from USD 1 billion to USD 4.3 billion in 2002/03 as a result of tighter controls over illegal transfers and scrutiny of foreign accounts. Similar controls could also be imposed on illegal currency swapping to enable foreign currency to play its role significantly as a hedge against inflation.

The issue under study could be explored further by subsampling the data or using high-frequency datasets. This may yield many interesting results about the dynamics of inflation and the hedging properties of different assets. The exercise could also be replicated for different exchange rate regimes in Pakistan.

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Choice of Anchor Currencies and Dynamic Preferences for Exchange Rate Pegging in Asia

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Abstract

This paper attempts to answer two important questions in the context of Asian exchange rate regimes with respect to the choice of anchor currencies and dynamic preferences for exchange rate pegging. According to our results, the US dollar is the first choice of a de facto peg for many countries such as China, Hong Kong, the Philippines, and Pakistan. Other countries, apart from Korea and Indonesia, seem to prefer a basket peg comprising two or more anchor currencies with rapidly increasing weight attached to the euro. This shift from the US dollar to the euro reflects changes in the choices, preferences, and policies of these economies as a result of varying macroeconomic and global financial realities.

Keywords: Exchange rate regime, flexibility, pegging, Asia.

JEL classification: E42, E58, F31, F33, F41.

1. Introduction

Despite its several demerits, exchange rate pegging has remained the most popular exchange rate regime for many countries, particularly during the 1990s when formulating their external financial policies. Undoubtedly, it has played a significantly constructive role in the development of several economies across the globe. It was common practice for developing countries with less stable currencies to create a currency peg with a more stable currency to achieve a variety of benefits including but not limited to stable inflation, monetary policy discipline, and control over trade deficits.

Pegged exchange rates have different forms, including unilateral pegs, multilateral pegs, currency unions, and currency boards. In a unilateral peg, an individual country decides on the basis of its indigenous

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and unique circumstances to peg its domestic currency to a single anchor currency. The domestic country sets its own monetary policy, which is similar to that of the anchor currency country, and tries to target different monetary variables such as inflation, credit expansion, and the interest rate in the currency of the anchor country.

A more rigid form of the unilateral peg—in which the country establishing a peg is prepared to lose its monetary independence by sacrificing its ability to decide the size of its monetary base—might be categorized as a currency board. Multilateral pegs and currency unions are arrangements where several countries jointly agree to follow specific sets of exchange rate parities. The responsibility for protecting the parity and intervening if needed is shared by all member countries. Multilateral pegs can be better understood by looking at the monetary system of the European Union before and after the adoption of a unified currency and a single monetary policy (Dueker & Fischer, 2001).

There are various pros and cons of pegged exchange rate regimes. On one hand, such a regime is extremely beneficial because it serves as a tool for keeping inflation under control and provides policymakers with discipline. On the other hand, it has certain disadvantages when it curtails the freedom of an autonomous monetary policy and increases the country's exposure to shocks and speculative attacks transmitted from the anchor country. It also binds policymakers to pursuing any anti-inflationary measures and potentially exposes the country to financial crises by making it financially fragile (Mishkin, 1998).

However, the sole fact of being a pegged exchange rate regime does not make a country vulnerable to financial crisis and fragility. Other macroeconomic factors also play a role in making an economy vulnerable. These factors include the associated monetary structure, inconsistencies in the exchange rate policy itself, supplementary macroeconomic policies, futile policy execution, the failure to practice effective structural and macroeconomic policies that might have fortified the financial system, and the pursuit of hasty disinflation and obstinacy in managing shocks (Bubula & Otker-Robe, 2003).

Hard pegs may seem a viable option for countries with a weak monetary system, but they are certainly not a cure that can be applied to every economy. A country that wishes to adopt a hard peg must be able to meet the following criteria: an optimum currency region, a large portion of its total trade should be with the anchor currency's country, its inflation preferences should be similar to those of the anchor country, flexible labor markets, and a strong commitment to following the regulations.

It is also very important to find the right peg, i.e., one that ties the domestic country to the right anchor currency. For example, if the home country has diverse trade partners, pegging its currency to a basket of currencies would be a more viable option than pegging it to a single currency (Velasco, 1999). The rule is, nevertheless, to keep the basket simple, observable, and transparent because a multifaceted and fluctuating basket system can lead to added complexities.

When countries adopt a basket peg, they must also be ready to disclose certain facts to the public such as what currencies they have included in the basket, what individual weights are assigned to each country, and why and how often they will change the corresponding weights. At times, governments choose not to share this crucial information because of the pressure exerted by political and international financial institutions.

Following primarily the methodology proposed by Frankel and Wei (1994), this paper has two aims: (i) to determine the most popular anchor currency for different Asian economies by analyzing the degree of pegging and estimating the implicit weights that different countries assign to that anchor currency or currencies; and (ii) to examine the dynamic behavior of these implicit weights, reflecting changes in the choices, preferences, or policies of these economies as a result of varying macroeconomic and global financial realities. Our findings generally acknowledge the undeterred importance of the US dollar as the most favored anchor currency for many South Asian countries. However, we also find the euro receiving more attention in several countries when assigning weights to basket anchor currencies.

Section 2 provides a description of the data used and sample countries. Section 3 describes the methodology used. Section 4 presents our results based on a static and dynamic analysis and discusses policy implications. Section 5 concludes the study.

2. Description of Data

Our dataset is composed of ten Asian economies: China, Hong Kong, Indonesia, India, the Republic of Korea (referred to as Korea), Malaysia, the Philippines, Pakistan, Singapore, and Thailand. Monthly exchange rate data from 2001:M01 to 2009:M12 has been taken from the IMF's International Financial Statistics database. Data on the Swiss franc (CHF) and special drawing rights (SDR) have also been taken for the same time period to be used as the *numeraire* currency. For analysis purposes, each country is classified according to its de jure and de facto regimes, provided in Table 1. However, instead of using several categories, each country is placed in one of three broad categories: floating, managed float, and pegged. Crawling peg and currency board arrangements are both categorized as "pegged." The terms "floating," "free-floating," and "independently floating" are used interchangeably.

Table 1: Exchange rate regime classification in Asia [2001-2009]

Country	De jure regime	De facto regime
China	Managed floating	Crawling peg
Hong Kong	Pegged	Currency board arrangement
Indonesia	Floating	Managed floating with no predetermined path for the exchange rate
India	Managed floating	Managed floating with no predetermined path for the exchange rate
Korea	Floating	Independently floating
Malaysia	Managed floating	Managed floating with no predetermined path for the exchange rate
Philippines	Floating	Independently floating
Pakistan	Floating	Managed floating with no predetermined path for the exchange rate
Singapore	Managed floating	Managed floating with no predetermined path for the exchange rate
Thailand	Managed floating	Managed floating with no predetermined path for the exchange rate

Source: IMF's de facto classification of exchange rate regimes and monetary policy frameworks at http://www.imf.org/external/np/mfd/er/2008/eng/0408.htm

3. Methodology

Identifying the true de facto exchange rate regime adopted by any country is not a trivial task and there are several ways of explaining the de facto practices of different countries. Keeping in view, however, the limited and specific nature of this study, we use the popular Frankel-Wei regression method to identify the implicit weights a country assigns to its probable currency basket, thus signaling its de facto exchange rate regime (see Frankel & Wei, 1994, 1995, 2006). This can be explained with the help of the following equation:

$$\Delta e_t^i = \alpha + \beta_1 \Delta e_t^{\$} + \beta_2 \Delta e_t^{£} + \beta_3 \Delta e_t^{€} + \beta_4 \Delta e_t^{¥} + \varepsilon_t \quad (1)$$

where Δe_t^i is the log difference of the exchange rate in the country's domestic currency and $e_t^{\$}, e_t^{£}, e_t^{€}$, and $e_t^{¥}$ represent the log of the US dollar, pound sterling, euro, and Japanese yen's exchange rates.

It is important to note that Frankel and Wei use the CHF as a base (*numeraire*) currency for calculating all exchange rates. However, some recent studies have used SDR and gold as a robustness test in response to the argument that the CHF is strongly correlated with the euro (Frankel & Wei, 2007). The estimated coefficients in this equation β_1, β_2, β_3, and β_4 represent the implicit weight that each country assigns to the USD, GBP, EUR, and JPY on the basis of proportionate variations in the domestic currency explained by each of these benchmark currencies. However, the higher value of the estimated coefficient alone is not necessarily an indication of pegging since it may have resulted from the natural market-driven correlation between two currencies. It is important to look at the significance of each coefficient as well as the overall variations explained (R-squared) by the regressor currencies to infer the de facto degree of pegging or flexibility (Baig, 2001).

4. Results and Policy Implications

4.1. Static Analysis

Table 2 presents two sets of OLS estimates obtained from equation (1) for each country. The first set is based on a regression where the CHF is used as the *numeraire* currency; in the second set, it is replaced by SDR. The US dollar seems to have an uncontestable influence over each currency irrespective of the choice of *numeraire*. However, a significant decrease in R-squared and either an entire shift in significance from the euro to the yen or an increase in the significance of the yen is often observed when the CHF is replaced by SDR. Due to these instabilities, further interpretations are based on the regression where the CHF is used as the *numeraire.*

China, Hong Kong, and Pakistan seem to be strongly pegged to the USD, which is the single-most highly significant currency and explains more than 95 percent of the variations for China and Hong Kong and almost 75 percent for Pakistan. India, Singapore, and Thailand seem to follow a basket peg where three out of four currencies are highly significant and explain at least 70 percent or more of the variation in their domestic exchange rates. Malaysia stands somewhere in the middle as two currencies, the USD and euro with their highly significant weights, explain almost 85 percent of the variation in the domestic currency. The Philippines also assigns significant weight to the USD and euro but the proportion of explained variation is relatively low at 65 percent. The Korean won and Indonesian rupiah are the least influenced by these benchmark currencies. In spite of the individual significance of the USD, euro, or pound, the four benchmark currencies explain only 37 percent of the variation for Korea and only 16 percent for Indonesia.

These findings are consistent with those of Rizvi, Bordes, and Naqvi (2012)—the authors disagree with the IMF's classification of Indonesia and the Philippines and provide strong evidence showing that Indonesia does not intervene in its exchange rate market or at least intervenes less than the Philippines. This also implies that, since the IMF has categorized the Philippines as a de facto "floating" regime, Indonesia should have been classified either under the same regime or as a more flexible regime.

Table 2: Static estimates from Frankel-Wei regression

Country	Reg. type	Constant	USD	GBP	EUR	JPY	Adj. R-sq.
China	CHF coef.	-0.00161***	0.995896***	-0.036478*	-0.020006	-0.02264	0.978
	t-stat.	(-4.63)	(52.52)	(-1.75)	(-0.57)	(-1.16)	
	SDR coef.	-0.00167***	0.924904***	-0.045163*	-0.016221	-0.01943	0.914
	t-stat.	(-4.76)	(23.12)	(-1.77)	(-0.62)	(-0.901)	
Hong Kong	CHF coef.	-0.000016	0.987115***	-0.001456	-0.013543	0.01314*	0.997
	t-stat.	(-0.13)	(151.22)	(-0.2023)	(-1.11)	(1.96)	
	SDR coef.	-0.000025	0.986182***	-0.000287	-0.003038	0.01717**	0.9897
	t-stat.	(-0.21)	(72.28)	(-0.03)	(-0.33896)	(2.34)	
Indonesia	CHF coef.	0.00252	0.358325**	0.081375	0.987***	-0.1294	0.166
	t-stat.	(0.78)	(2.04)	(0.42)	(3.02)	(-0.72)	
	SDR coef.	0.00252	-0.270413	-0.046114	0.192939	-0.37988*	0.043
	t-stat.	(0.781)	(-0.74)	(-0.197)	(0.7998)	(-1.93)	
India	CHF coef.	0.00023	0.684167***	0.263455***	0.43749***	0.09432	0.707
	t-stat.	(0.17)	(9.21)	(3.23)	(3.16)	(1.23)	
	SDR coef.	0.000549	0.678939***	0.207088**	0.075706	-0.05117	0.193
	t-stat.	(0.39)	(4.22)	(2.02)	(0.72)	(-0.59)	
Korea	CHF coef.	0.000645	0.313147***	0.403447***	0.70885***	-0.00665	0.3699
	t-stat.	(0.298)	(2.65)	(3.11)	(3.22)	(-0.05)	
	SDR coef.	0.001028	-0.09932	0.229847	0.104233	-0.267306**	0.1089
	t-stat.	(0.48)	(-0.41)	(1.48)	(0.65)	(-2.03)	
Malaysia	CHF coef.	-0.00029	0.803385***	0.051305	0.19513**	-0.044526	0.852
	t-stat.	(-0.36)	(18.05)	(1.05)	(2.35)	(-0.97)	
	SDR coef.	-0.000301	0.579935***	-0.003202	0.006914	-0.11296**	0.444
	t-stat.	(-0.37)	(6.32)	(-0.05)	(0.11)	(-2.29)	
Philippines	CHF coef.	-0.00025	0.871139***	-0.045055	0.43259***	0.002329	0.6375
	t-stat.	(-0.152)	(9.92)	(-0.47)	(2.64)	(0.026)	
	SDR coef.	-0.00011	0.643005***	-0.124446	0.038673	-0.141788	0.2001
	t-stat.	(-0.0698)	(3.49)	(-1.060107)	(0.31947)	(-1.43)	
Pakistan	CHF coef.	0.00398***	0.911517***	-0.038313	0.088501	-0.098879	0.756
	t-stat.	(3.42)	(14.35)	(-0.548994)	(0.747928)	(-1.51)	
	SDR coef.	0.00393***	0.814724***	-0.03385	0.136077	-0.080413	0.376
	t-stat.	(3.42)	(6.22)	(-0.405707)	(1.581589)	(-1.14)	
Singapore	CHF coef.	-0.0005	0.52609***	0.06109	0.40939***	0.12838***	0.7996
	t-stat.	(-0.65)	(12.36)	(1.31)	(5.16)	(2.93)	
	SDR coef.	-0.00048	0.23157**	-0.01255	0.065937	0.00824	0.055
	t-stat.	(-0.602)	(2.61)	(-0.22)	(1.13)	(0.17)	
Thailand	CHF coef.	-0.0013	0.47809***	0.17798***	0.163956	0.268569***	0.702
	t-stat.	(-1.23)	(8.22)	(2.79)	(1.51)	(4.49)	
	SDR coef.	-0.0013	0.30961**	0.117691	-0.04411	0.186971***	0.134
	t-stat.	(-1.197)	(2.57)	(1.53)	(-0.56)	(2.88)	

4.2. *Time-Varying (Dynamic) Assessment*

For a deeper insight into how the influence of benchmark currencies over Asian currencies has evolved over the period under study, we augment the static Frankel-Wei estimates by calculating recursive OLS estimates. Recursive time-varying estimates are obtained through iterative estimations of the same regression presented in equation (1), where the sample size is increased by 1 in every iteration. These estimates help analyze the dynamic trend in the benchmark currency weighting that each country applies while determining the value of its own currency.

Figure 1 shows that, for China and Hong Kong, despite some variations, the USD remains the sole anchor even after 2005 when China announced it would follow a managed floating regime. There is some degree of substitution of the EUR for the USD in China immediately after 2005, but the dollar regains its weight during 2007–08.

Figure 1: Recursive coefficients from Frankel-Wei regression, China and Hong Kong

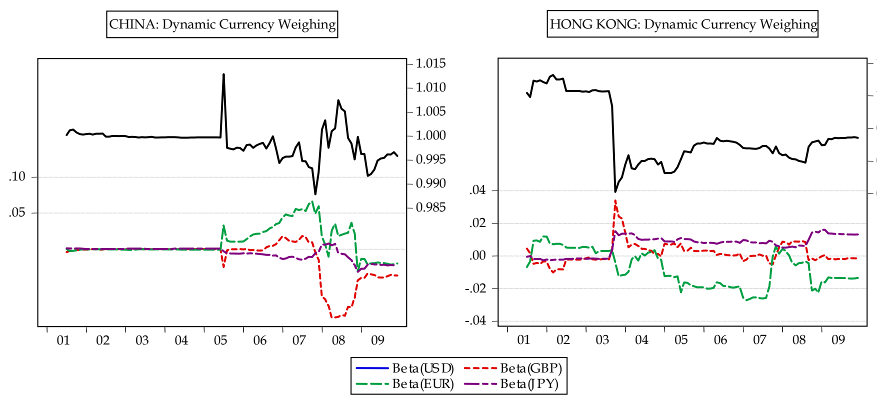
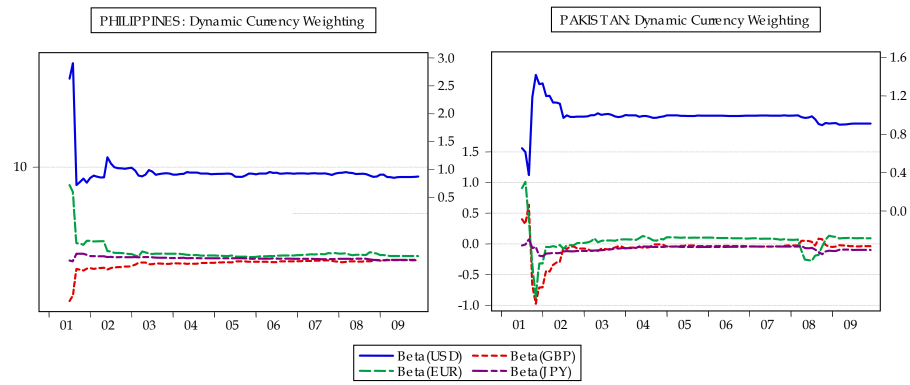


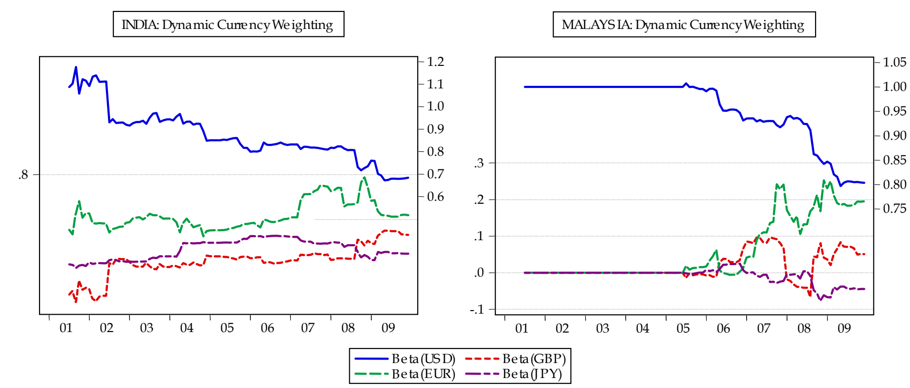
Figure 2 suggests that both the Philippines and Pakistan are also strongly influenced by the USD. However, unlike China and Hong Kong, they exhibit a very smooth pattern in dynamic estimates of all the anchor currencies. Indeed, the observed pattern indicates de facto pegging with very minor fluctuations at the end of the sample period. These fluctuations could be a natural outcome of increasing concerns about the weakening dollar following the negative impact of the subprime crisis in the US.

Figure 2: Recursive coefficients from Frankel-Wei regression, the Philippines and Pakistan



According to Figure 3, India seems to have been inclined toward the EUR since its inception while the USD's influence over the Indian rupee started to decline in 2002 with a moderate increase in the weight of the EUR. During this nine-year period, almost 40 percent of the USD's share moved toward other currencies with the EUR being a significant recipient. Exactly the same has happened in Malaysia but over significantly less time and at a faster pace: at times, the USD and EUR weights seem equivalent.

Figure 3: Recursive coefficients from Frankel-Wei regression, India and Malaysia



Figures 4 and 5 show that Singapore, Thailand, Indonesia, and Korea have relied on at least two or more currencies, negating the assumption of a hard peg with the USD. For Singapore and Thailand, the USD weight is round 0.5 on average over the whole sample period; for Indonesia and Korea, it is even less (approx 0.35) with positive weights allocated to at least two other benchmark currencies on a consistent basis.

Figure 4: Recursive coefficients from Frankel-Wei regression, Singapore and Thailand

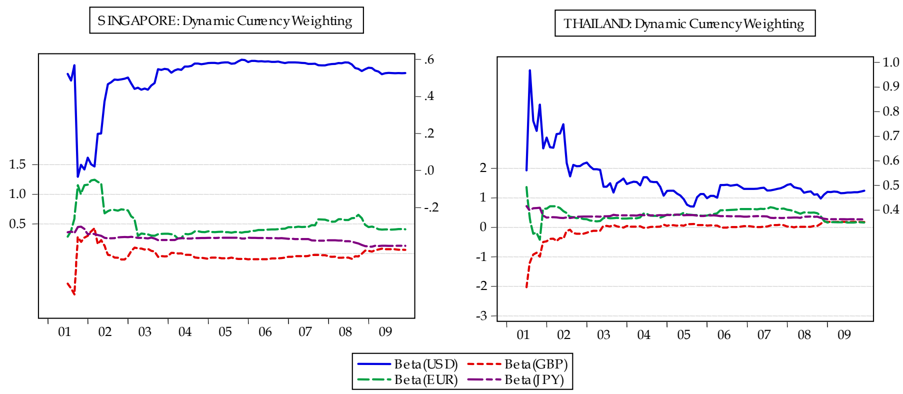
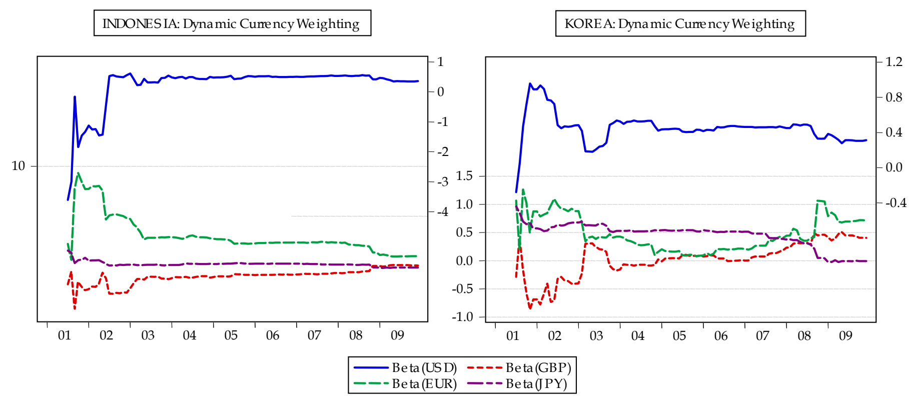


Figure 5: Recursive coefficients from Frankel-Wei regression, Indonesia and Korea



The anchoring of domestic currencies to multiple anchors may be an indication of a basket peg in the case of Singapore and Thailand because the proportion of explained variations in the domestic currency is about 70 percent. However, as for Indonesia and Korea, this proportion is only 16 and 36 percent, respectively, suggesting the existence of a market-driven natural correlation between the domestic currencies and benchmark currencies, thus precluding any possibility of a basket peg.

5. Conclusion

Applying a static and then dynamic Frankel-Wei analysis, we have assessed the degree and time-varying behavior of currency pegging for a selected sample of ten Asian economies. According to the results, the USD is the first choice of a de facto peg for many countries, including China, Hong Kong, the Philippines, and Pakistan. The other countries in our sample, except Korea and Indonesia, seem to have adopted a basket peg of two or more currencies with a rapidly increasing weight assigned to the euro.

The small R-squared term in the regression for Indonesia seems to imply that it is paying the penalty for some flaw in the IMF's methodology by being categorized as a managed float contrary to its de jure claim of being free-floating. On the other hand, the Philippines seems to be in an advantageous position as the IMF agrees with its de jure claim of being a free floater despite the fact that its real practices clearly indicate significant pegging to the basket of four currencies.

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Economic Returns to Education in France: OLS and Instrumental Variable Estimations

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Abstract

This article estimates the economic returns to schooling as well as analyzing other explanatory factors for the French labor market. It addresses the issue of endogeneity bias and proposes two new instruments for use in the instrumental variable two-stage least squares technique. Our results show that the proposed instruments are relevant and adequate, based on evidence from the available literature. After using the proposed instruments, we find that the OLS coefficients for schooling are biased downwards. Finally, we choose between the two proposed instruments.

Keywords: Endogeneity bias, instrumental variable, Mincerian model, two-stage least squares, wage regression, France.

JEL classification: C1, I2, J3, P5.

1. Introduction

In the labor market, individuals are rewarded depending on their skills, competencies, or knowledge. These skills and competencies are more typically referred to as human capital. Human capital theory (Mincer, 1958, 1974; Becker, 1964) states that education and training increase the productivity of individuals by augmenting their skills and knowledge. Education and training are, therefore, key factors in determining the economic performance of an individual. The rates of return to different human capital factors can help individuals make decisions regarding their investment in education, based on their possible future earnings. It is also important for policymakers to allocate resources such that they discourage or eliminate discrimination in economic rewards on the basis of geography, ethnicity, sex, or age.

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Mincer's (1974) econometric wage regression model has been commonly used to estimate the returns to schooling. Some notable early work on this model includes studies by Chiswick (1983), Kuch and Haessel (1979), and Tomes (1983). Despite its wide-ranging applications, the model's simple estimation can be biased due to econometric problems of endogeneity, measurement error, and sample selection bias (Ashenfelter & Krueger, 1994; Card, 1999; Griliches, 1977, 1979).

In this article, we focus on the bias arising from the endogeneity of schooling, which can restrict the estimation of the true causal effect of schooling on wages (Griliches, 1977). The endogeneity problem arises due to the violation of an assumption under the Mincerian model that treats all individuals as identical (with respect to ability, opportunities, and environment) other than differences in education and training. This violation is inevitable given that different people cannot be identical with respect to certain unobservable characteristics such as social environment, location, and family background. For example, ability can be seen as a determinant of wages in the labor market on the one hand, and may also be correlated with schooling on the other, i.e., more able people tend to acquire more schooling, will be more productive at work, and hence paid better.

If unobserved ability affects both schooling and wages, then an ordinary least squares (OLS) estimation will yield biased results (Griliches, 1977; Card, 2001). In such a situation, the schooling variable will be correlated with the error term in the wage equation and, as a result, the coefficient associated with schooling will be biased. This type of bias is known as an endogeneity bias and can be tackled by using the instrumental variable two-stage least squares (IV2SLS) estimation approach. For this, we need instrumental variables that affect schooling but are otherwise uncorrelated with wages.

Different types of instruments have been used to counter the endogeneity bias, including parental or spouse education (see Blackburn & Neumark, 1993; Ozdural, 1993; Trostel, Walker, & Woolley, 2002; Zhang, 2011, among others) and the availability of (or distance to) a nearby college or school (see Card, 1993; Maluccio, 1998, among others). Many other instruments can also be found in the literature (see Angrist & Krueger, 1991; Brunello & Miniaci, 1999; Harmon & Walker, 1995; Ismail, 2007). The consensus in the literature is that the returns to schooling obtained from IV2SLS estimation are typically higher than those from OLS estimation.

Inspired by numerous studies on the relationship between earnings/wages and their determinants in different countries, a number of studies have been carried out using French data to investigate different aspects of the French labor market (see, for example, Abdelkarim & Skalli, 2005; Boumahdi & Plassard, 1992; Fougère, Goux, & Maurin, 2001; Selz & Thélot, 2004). However, few studies address the issue of the endogeneity of schooling in this context (e.g., Boumahdi and Plassard, 1992; Steunou, 2003; Viger, 2007).

We aim to eliminate the possible endogeneity bias in returns to schooling when applied to French data with the help of certain new instruments. Generally, researchers working with French data have tackled the issue of endogeneity bias by using parental education and family background variables as instruments. However, such instruments have been justifiably criticized due to their direct influence on wage determination in the labor market (Flabbi, 1999). This study is different from previous studies in that it tries to address the endogeneity problem using new instrumental variables that are expected not to have a direct impact on wages in the work market.

2. New Instrumental Variables and French Data

As with the correlation between an individual's schooling and his/her parents' education, it is reasonable to suppose that the former could also depend on the general trend in and motivation for schooling among that individual's family members. Given the likely similarities in environment (social and otherwise) and era, one's education may have a higher correlation with that of one's siblings relative to one's parents; this qualifies the variable as a more reasonable instrument than parents' education. The natural question to ask is which sibling's education should be used as an instrument, which leads us to use average household schooling as an instrument for endogenous schooling.

Based on this notion, we introduce the first instrument (Z1) as "the household's average number of years of education." This variable reflects the fact that different persons in a family may have different educational inclinations and levels of ability. Thus, averaging the schooling of all family members eliminates the bias of ability. Since it is a data-generated instrument, its key benefit is that it can be used in any study that provides some information on household education. The instrument combines the effects of parents' and siblings' education on the schooling of an individual.

The educational decisions of an individual may also be affected by different factors at the time that he/she decides whether to work or to continue schooling. Thus, our second proposed instrument (Z2) is "the average years of schooling for a specific gender of a specific age group in the particular year that they join the labor force." In this case, the ability bias is eliminated as different people come from different areas and different families (some with low ability and some with high ability). Any effect of school proximity is also balanced out by using this overall average. Technically, Z2 combines several instruments: ability, school proximity, family background, school quality, etc. The average schooling for the period 1950–2010[1] has been calculated by Barro and Lee (2010).

To estimate the Mincerian wage regression for the French labor market, we use data from the Labor Force Survey (Enquête d'Emploi Continu) carried out by the Institut National de la Statistique et des Etudes Économiques in 2007. Although the Mincerian model's initial formulation was used to regress only the log of wages on a linear term for schooling and linear and quadratic terms for experience (or potential experience), most studies now use an extended version of the model, which contains other explanatory variables in addition to schooling and experience.

Our set of explanatory variables therefore includes different variables that can also affect wage determination, especially in the context of the French labor market. Table 1 defines all the variables used in our estimation. However, to compute the second instrument (Z2), Barro and Lee's (2010) measures interact with the gender, year, and age group variables relevant to an individual joining the labor force.

[1] These measures are available at http://www.barrolee.com/ for many countries of the world, including France.

Table 1: Variables in estimation of Mincerian model for French data

Variable	Description
Response variable	
LNWAGE	Natural logarithm of the monthly wage of an individual from main job
Explanatory variables	
SCH	Variable for education, measured in number of years of completed schooling
BEFEX	Experience (in labor market before current job), measured in number of years
BEFEX2	Experience squared
EXP	Current job seniority (labor market experience within present job), measured in number of years
EXP2	Current job seniority squared
HOURS	Number of hours associated with monthly salary (i.e., hours worked per month)
DGENDER	Gender of individual (male = 1, female = 0)
RDRURAL0	Dummy variable indicating whether individual is resident of rural area or not (reference category)
DNPARIS	Dummy variable indicating whether individual is resident of urban area other than Paris region or not (non-Paris urban = 1, otherwise = 0)
DPARIS	Dummy variable indicating whether individual is resident of Paris region or not (Paris region = 1, otherwise = 0)
DTYPDIP	Dummy variable indicating whether individual has degree/diploma in professional education or general education (professional = 1, general = 0)
DPUBLIC	Dummy variable indicating whether individual works in public sector or private sector (public = 1, private = 0)
RDTMPCT	Dummy variable indicating that individual is engaged in temporary work (reference category)
DFIXCT	Dummy variable indicating that individual works under fixed-term contract (fixed-term contract = 1, otherwise = 0)
DPERCT	Dummy variable indicating that individual works under permanent contract (permanent contract = 1, otherwise = 0)

3. Results and Discussion

Table 2 presents the results of human capital wage regressions based on the French data, using OLS and IV2SLS. In the first-stage regressions, the lower p-values associated with the instruments suggest the relevance of both instruments. The results show that the endogeneity problem causes a downward bias in returns to schooling for the OLS estimates. The schooling coefficient in the IV2SLS estimation (with Z1) is about 14 percent higher than that in the OLS estimation. This finding is in accordance with the literature, where endogeneity-corrected (IV2SLS) estimates are 10–30 percent (or approximately two percentage points) higher than OLS estimates (see Card, 1994; Ashenfelter, Harmon, & Oosterbeek, 1999). The schooling coefficient in the IV2SLS estimation (with Z2) is nearly double the corresponding OLS coefficient. The higher returns to schooling yielded by the IV2SLS approach are confirmed by several other researchers using French data (see Boumahdi & Plassard, 1992; Viger, 2007).

Before interpreting the coefficients in relation to the other explanatory variables, we need to choose between the two instruments Z1 and Z2. From the IV2SLS estimates based on Z1 and Z2, it is clear that the OLS estimates are downward-biased. The magnitude of difference between the OLS and IV2SLS estimates for the effect of education depends on the instrument used for schooling in the first-stage schooling equation. The explanatory power of Z1 is greater than that of Z2 in terms of the first-stage R^2 (see Table 2).

Table 2: OLS and IV2SLS estimates

		IV2SLS-Z1		IV2SLS-Z2	
Variable	**OLS**	**First-stage**	**Second-stage**	**First-stage**	**Second-stage**
INTERCEPT	4.7342	0.9943	4.6386	6.9604	4.0385
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
SCH	0.0670	–	0.0766	–	0.1371
	(<0.0001)		(<0.0001)		(<0.0001)
BEFEX	0.0133	-0.0494	0.0145	-0.1482	0.0227
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
BEFEX2	-0.0002	0.0005	-0.0002	0.0014	-0.0003
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
EXP	0.0184	-0.0121	0.0189	-0.0089	0.0221
	(<0.0001)	(0.0001)	(<0.0001)	(0.1277)	(<0.0001)
EXP2	-0.0001	-0.0003	-0.0001	-0.0008	-0.0001
	(<0.0001)	(0.0003)	(<0.0001)	(<0.0001)	(0.0086)
HOURS	0.0064	0.0012	0.0063	0.0047	0.0060
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
DGENDER	0.1687	-0.2163	0.1734	-0.4695	0.2027
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
DNPARIS	0.0117	0.0621	0.0082	0.3552	-0.0137
	(0.0159)	(0.0045)	(0.0915)	(<0.0001)	(0.0312)
DPARIS	0.2212	0.2631	0.2056	1.5910	0.1077
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
DTYPDIP	-0.0498	0.6557	-0.0619	1.2536	-0.1378
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
DPUBLIC	0.0049	0.0594	0.0014	0.3518	-0.0210
	(0.4736)	(0.0555)	(0.8423)	(<0.0001)	(0.0129)
DFIXCT	0.3415	0.6011	0.3231	1.4442	0.2079
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
DPERCT	0.4963	0.4792	0.4760	1.5706	0.3487
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
Instrument	–	0.8916	–	0.2922	–
		(<0.0001)		(<0.0001)	
Hausman exogeneity test		262.58		64.66	
		(<0.0001)		(<0.0001)	
R^2	0.5528	0.7459	0.5435	0.2856	0.4355

Note: p-values given in parentheses.
Source: Authors' calculations.

In order to choose the more appropriate instrument, we look at the nature and behavior of both Z1 and Z2. Since Z1 represents the average number of years of education for the household, there may be some issues when using this instrument. For example, an individual's ability may be related to his/her family situation or to a genetically determined inclination to pursue education. Furthermore, some household members might face similar schooling costs or opportunities (in terms of distance or the availability of educational institutions, etc.). Common demographics and geography can also affect schooling attainment because members of the same household are more likely to share a social and community environment. On the other hand, the impact of similar demographics, geography, schooling costs, opportunities, and genetic ability when using Z1 is eliminated in the case of the second instrument (Z2) because it is an overall average (as defined in Section 2).

A further comparison of Z1 and Z2 can be made with the help of a correlation matrix, given in Table 3. The correlation between Z1 and SCH is higher than that between Z2 and SCH. However, the correlation between the response variable (LNWAGE) and Z1 is also greater than that between LNWAGE and Z2. Although Z2 has a relatively low correlation with schooling, it also has a very low correlation with the response variable. Finally, keeping in view the correlation matrix and the more compact definition of Z2, we find this instrument more suitable for IV2SLS estimation. We refer to IV2SLS-Z1 and IV2SLS-Z2 in the IV2SLS estimations using the instruments Z1 and Z2, respectively.

Table 3: Correlation matrix for response, endogenous variable, and instruments

Variable	LNWAGE	SCH	Z1	Z2
LNWAGE	1.0000	0.3399	0.3360*	-0.1360
SCH		1.0000	0.8481*	0.2728*
Z1			1.0000	0.2633
Z2				1.0000

Note: * = significant at 5 percent level of significance.
Source: Authors' calculations.

Having selected the second instrument, Z2, as the more appropriate of the two, we now discuss the results of IV2SLS-Z2. The returns to schooling are about 13.7 percent for each additional year of schooling, which is twice that obtained from the OLS estimation. The returns to labor market experience are also higher than the corresponding OLS estimate. Each additional year of experience yields a 2.27 percent increment. Similarly, each extra year of current job seniority increases wages by approximately 2.21 percent.

The impact of work duration is almost identical in all specifications. The results reveal that women earn 20 percent less than their male counterparts and this gap is three percentage points higher in the IV2SLS-Z2 estimation. The wage differential due to regional differences changes considerably in the IV2SLS-Z2 estimation when compared to the OLS results. The differential between non-Paris urban and rural workers favors rural workers by a magnitude of 1.4 percent; in the OLS estimation, however, this differential favors non-Paris urban workers. The wage gain for workers in the Paris region over rural workers is 10.77 percent, which is 11–12 percentage points lower than the OLS estimate.

The coefficient of the professional diploma dummy variable indicates that professional diploma holders earn 13.8 percent less than those with a general education degree. This finding contradicts the previous evidence from other studies that have used French data (see Tansel, 1994; Simonnet, 1996). Individuals working in the private sector earn roughly 2 percent more than those working in the public sector; this difference is significant only in the IV2SLS-Z2 specification. We also find that individuals on fixed-term and permanent contracts earn 21 percent and 35 percent more, respectively, than their counterparts who work as temporary workers. These wage premiums increase by about 15 percent when endogeneity is taken into account. The direction of these findings is consistent with Araï, Ballot, and Skalli (1996) but the magnitude of the penalty for temporary workers is much larger in our investigation.

4. Conclusion

When estimating the economic returns to education using the Mincerian wage regression model, the problem of endogeneity bias is liable to arise. The 2SLS method with some potential instruments is commonly used in this situation. For the French data, we have used two new instruments: (i) the household's average number of years of education and (ii) the average schooling for a specific gender and specific age group in the particular year that an individual joins the labor force. The second instrument is found to be more appropriate.

Using this instrument, we find that the returns to schooling are about 13.7 percent for each additional year of schooling. Each year of experience yields an increase of about 2.27 percent in wages. French men are likely to earn 20 percent more than women. Workers in the public sector earn more than in the private sector, while temporary workers earn considerably less.

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Appendix

Table A1: Summary statistics

Variable	Mean	Median	SD	SE	Kurtosis	Skewness
LNWAGE	7.26	7.24	0.49	0.003	1.50	-0.33
SCH1	11.66	11.00	2.93	0.02	0.58	-0.05
BEFEX2	9.51	6.00	9.76	0.06	0.31	1.08
BEFEX22	185.62	36.00	305.92	1.86	5.77	2.31
EXP2	10.23	6.60	10.19	0.06	0.15	1.09
EXP22	208.33	43.56	336.95	2.05	3.67	2.05
HOURS3	144.65	151.00	31.21	0.19	3.52	-1.76
Z1	11.51	11.25	2.54	0.02	0.52	0.16
Z2	10.21	10.77	1.95	0.01	0.13	-0.95
N = 27,136						

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On the (Ir)Relevance of Monetary Aggregate Targeting in Pakistan: An Eclectic View

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Abstract

This study attempts to identify a stable money demand function for Pakistan's economy, where the monetary aggregate is considered the nominal anchor. With evolving financial innovations and regulations, the stability of money demand has been the focus of numerous debates. Where earlier studies have provided conflicting explanations due to inadequate specifications and imprecise estimations, we find that money demand in Pakistan is stable, if specified properly. For developing countries such as Pakistan, it is important to target monetary aggregates or respond to deviations from the desirable path if monetary policy is to be effectively implemented and communicated; this should remain, if not a primary, then an auxiliary target in the monetary policy framework.

Keywords: money demand, stability, monetarism.

JEL classification: C20, E12, E41, E5.

1. Introduction

The role of monetary aggregate(s) as a virtue has long been debated in the conduct of monetary policy. It is generally believed that money carries important information about the underlying state of an economy and can help predict discernible monetary facts. Although most central banks have used monetary aggregates as an intermediate target in their monetary policy frameworks, the usefulness of monetary aggregates has diminished, particularly since the early 1990s when the money demand function became subject to structural changes (see, for instance, Mishkin, 2000; Woodford, 1998, 2008).

It is generally argued that successful monetary aggregate targeting requires a stable, or at least predictable, relationship between money growth and inflation. However, this relationship has become more

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obscure, particularly in advanced economies, since the 1990s with the evolution of the financial sector.[1] Financial deregulation and innovation have significantly changed the preferences of households and the financial sector and thus destabilized the money demand function (Arrau, De Gregorio, Reinhart, & Wickham, 1995; Bernanke, 2006).[2] As a consequence, many developed and developing countries have changed their nominal anchor and switched from a monetary aggregate targeting regime to inflation targeting or price level targeting.

Nevertheless, monetary aggregates contain important information and their significance in monetary policy frameworks cannot be ignored. A detailed assessment of monetary trends and their relationships with goal variables (output growth and inflation) can provide useful information on the demand pressures in an economy—this, in turn, drives our analysis of the specification and stability of the money demand function in Pakistan.

The core functional specification of money demand is derived from a set of intertemporal optimal decisions made by households and firms in a dynamic stochastic general equilibrium (DSGE) setting. This specification is then estimated empirically using various econometric techniques while investigating other potential determinants of money demand in terms of their goodness of fit. This is because the preferences of households and firms have, to some extent, changed and the operational scope of the financial sector widened over the last two decades (see State Bank of Pakistan, n.d.). However, the development of structured financial instruments is still in an evolutionary phase, making it necessary to analyze different empirical specifications for a robustness check.

Another important empirical contribution of this paper is that it provides an eclectic stability analysis both in terms of models and parameters. Although numerous empirical estimations of the money demand function have focused on the stability of money demand and the velocity of money in the context of Pakistan, the results are mixed.[3] Some studies find a robust relationship between money and the goal variables (see Qayyum, 2005; Omer, 2010; Azim, Ahmed, Ullah, Zaman, & Zakaria, 2010) while other find an unstable money demand function (see

[1] When households started to invest in bonds and mutual funds.
[2] Lieberman (1977) argues that the "increased use of credit, better synchronization of receipts and expenditures, reduced mail float, more intensive use of money substitutes, and more efficient payments mechanisms will tend to permanently decrease the transaction demand for money over time."
[3] A quick review of earlier empirical studies on the money demand function in Pakistan is given in Appendix I.

Moinuddin, 2009; Omer & Saqib, 2009). It is interesting to note that, even when using almost the same specification and methodology but a different sample range, researchers have arrived at conflicting results.[4] In addition, while earlier studies talk about model stability, they do not evaluate the parameter stability of money demand. Among other variables, the consistency of interest rate sensitivity is important for the stability of money demand: changing interest rate sensitivity over time would mean that a money demand estimate in one period may not be able to predict money demand as well in other period (Mishkin, 1995).

There are a number of ways to empirically determine the stability of money demand. Earlier studies show that the mere existence of a long-run relationship between monetary aggregates and their determinants is a sign of stable money demand. However, this argument does not qualify for stability and requires more statistical tests to determine whether both long-run and short-run elasticities remain stable over time (Bahmani-Oskooee & Rehman, 2010). These tests include the recursive estimation of coefficients and analysis of evolution. If the coefficients vary significantly—both in magnitude and sign—as more information is added to a sample, then this would indicate instability.

The problem of instability may not necessarily be due to the incorrect specification of a long-run function but due to the inadequate modeling of short-term dynamics (see, for example, Laidler, 1993). Therefore, it is important that the money demand function should be correctly specified both in the long term and in the short term.

Another way to determine money demand stability is to analyze the velocity of monetary aggregates. Omer and Saqib (2009) model the velocity of money (M0, M1, M2) in a univariate fashion and argue that each series of monetary aggregates is not mean-reverting and is integrated of high order, i.e., I(1). They conclude that the velocities are nonstationary at level, which signifies instability in money demand. However, this analysis does not qualify as determining money demand stability because most economic time-series depicting a trend (and thus nonstationarity) can be explained by structural changes[5]—an empirical point highlighted by

[4] Qayyum (2005) and Moinuddin (2009) use real M2, real GDP, and the call money rate for the period 1960–99 and 1974–06, respectively.

[5] During the last two decades, Pakistan's economy, particularly its financial sector, has undergone structural changes that include the opening of the equity market and long-term and short-term bond market for government securities, the introduction of foreign currency accounts, the liberalization (to some extent) of external accounts, and opening of new domestic and foreign private banks.

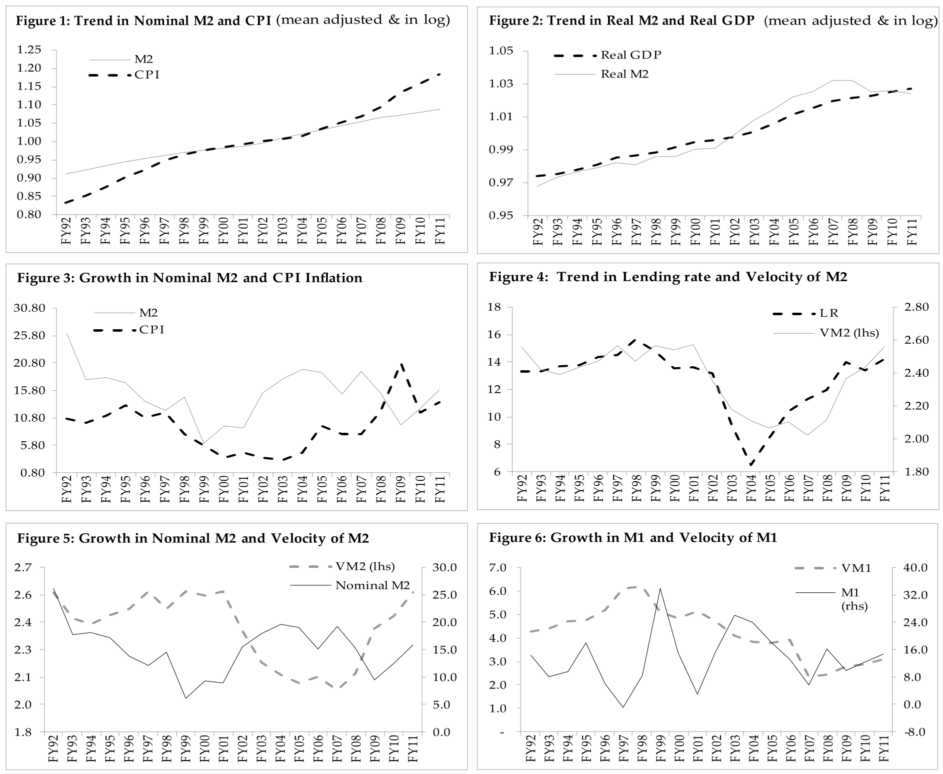
Ericsson, Hendry, and Prestwich (1997). Therefore, if an empirical specification were modeled properly, the result could be different.

Section 2 gives some stylized facts about monetary aggregates in Pakistan during 1992–2011. Section 3 describes our empirical methodology in detail. Section 4 discusses the empirical estimation, stability issues, and results of the money demand function, and Section 5 concludes the study.

2. Some Stylized Monetary Facts in Pakistan

Before investigating model specifications and technical details, it is important to visualize the various temporal developments of selected macro/monetary economic indicators in Pakistan. Figures 1–6 show the trends in broad money (M2) and the consumer price index (CPI); the variables are mean-adjusted and in log form in order to bring in one scale. Over the last two decades, nominal money has increased around fourteen-fold while consumer prices have risen four times from the level of FY1991, which witnessed a significant increase in M2.

Figure 1–6: Developments in monetary aggregates in Pakistan

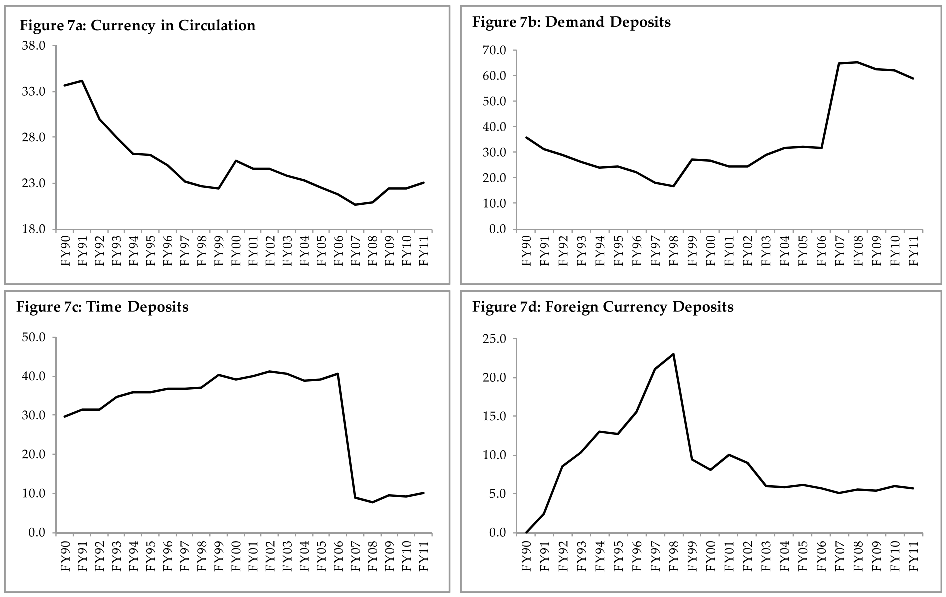


In Figure 2, real money increases parallel to real GDP, gaining pace in 2001 and onward mainly due to a significant increase in foreign inflows (although sterilized by some inflows from the central bank through FY2004) and subdued inflation. This phenomenal increase in nominal and real money has had serious repercussions for the economy in terms of high inflation and protracted low economic growth.

Figure 3 shows that consumer price inflation stabilized once monetary growth was controlled (particularly in 1995, when money targeting was institutionalized). However, the significant growth in nominal money in the early 2000s induced inflationary pressure in the subsequent period. During the last decade, nominal money increased by 15.2 percent per annum on average while CPI inflation increased to 8.4 percent per annum on average. This inflationary pressure was more pronounced after FY2008 due to frequent government borrowing from the State Bank of Pakistan (SBP) for budgetary support (inflationary in nature) coupled with an exorbitant rise in international commodity prices—particularly energy prices—in FY2008 and the erosion of domestic currency.

Money velocity (VM2) fell moderately from 2.6 in FY2001 to around 2.0 in FY2007 and the interest rate almost touched bottom in FY2004 (Figure 4). Interestingly, the velocity of narrow money shows more volatility than that of broad money due to the impulsive nature of demand deposits (DD) and currency in circulation (CIC) (Figures 5 and 6).

Figure 7: Components of broad money



In Pakistan, broad money (M2) includes CIC, deposits with the SBP, demand and time deposits, and foreign currency deposits. Figure 7 shows the contribution of the four major components of broad money, where the share of CIC declines from a little over 34 percent in 1991 to 22 percent in 2010. The contribution of DD to broad money also declines during the 1990s, but bottoms out in 1998 after the freezing of resident foreign currency deposit (RFCD) accounts.

The earlier decline in CIC and DD was primarily due to the introduction of RFCDs in 1991; their share in broad money increased to 23 percent in 1998. However, subsequent to Pakistan's nuclear tests, foreign currency accounts were frozen for fear of foreign sanctions and capital flight. On the other hand, time deposits are seen to increase moderately over time. The data indicate a clear shift from time deposits to DD in 2007 due to a change in classification: deposits with a six-month to one-year tenure (previously time deposits) are now reported as DD. Together, demand and time deposits constitute around 75 percent of broad money.

3. Empirical Methodology

The standard money demand function in linear form is written as:[6]

$$\widetilde{\left(\frac{M_t}{P_t}\right)} = \varphi_0 + \varphi_1(\widetilde{Y_t}) + \varphi_2(\tilde{\imath}_t) + \tilde{\varepsilon}_{m,t}$$

Where φ_1 and φ_2 are the output and interest elasticities of money demand, respectively. From a theoretical perspective, the money demand function is positively related to output and negatively to the nominal interest rate.

This specification of money demand is consistent with the basic version of Friedman's money demand function. However, Friedman and his followers also consider other potential determinants of money demand: $\mathrm{M_d/P = f(W, r - r^e, \pi^e, h)}$, where $\mathrm{M_d/P}$ denotes real money balances, W is wealth, r is the interest rate, $\mathrm{r^e}$ is the expected change in the interest rate, is the expected change in price level, and h is the ratio of human to nonhuman wealth. In practice, however, it is difficult to determine the volume of wealth in the economy and, therefore, a scale variable (usually real GDP or, in some instances, real consumption) is used as a proxy as we also observe in the final specification of the above micro-founded model.

There is widespread agreement in the empirical literature on the choice of scale variable although various theories of money demand have highlighted the importance of different scale variables. For instance, transaction-related theories suggest real income while portfolio approaches emphasize financial wealth (Calza, Gerdesmeier, & Levy, 2001). Regarding empirical estimates of income elasticity (i.e., the responsiveness of the demand for money to changes in real income), the quantity theory suggests a one-to-one relationship between real balances (M2) and real income while the Baumol-Tobin model of the transaction demand for money specifies 0.5 (Calza et al., 2001). However, in developing countries, income elasticity is much higher (more than 1 percent) mainly due to insufficient avenues for alternative financial assets and the pervasive monetization of the economy (Adam, Kessy, Nyella, & O'Connell, 2010). Some empirical studies use stock prices or volatility as an additional variable in the money demand function to substantiate the wealth effect (Bruggeman, Donati, & Warne, 2003).

Economic theory also considers the interest rate an important variable that reflects the opportunity cost of holding money, but it provides

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[6] Appendix A deals with the derivation and analytical foundations of the money demand function outlined here.

little guidance on selecting an appropriate interest rate (Laidler, 1993). The empirical literature uses a variety of interest rates, including the short-term market or bond rate, the long-term rate, and the rate of return on alternative financial assets. In portfolio decision-making, economic agents treat a variety of assets as alternatives to money, and therefore a wide spectrum of rates of return should be included in the money demand function (Heller & Khan, 1979). However, this raises some statistical issues (i.e., most interest rates are co-linear) and complicates the estimation of the model. Since most interest rates move in more or less the same direction, researchers are best restricted to limited rates of return (Calza et al., 2001).

In the case of developing countries, where the financial market (particularly the long-term bond and equity markets) is not fully developed, the money demand function should include the short-term market interest rate. Further, economic agents may choose short-term financial assets as an alternative to money in high inflationary environments while opting for long-term assets when economic conditions are stable and predictable.

Since a significant portion of monetary aggregates (M2 or M3) is remunerative (including demand and time deposits, foreign currency deposits, and various saving schemes), the own rate of return on monetary aggregates cannot be ignored. Bruggeman et al. (2003) use the weighted average rate of return on different components of M2/M3 to calculate the own rate for each country in the euro area. It is generally expected that the coefficient of the own rate of interest will have a nonnegative value, i.e., an increase in the own rate will raise the demand for real balances (Laidler, 1993). However, a tight monetary stance may raise the own rate of interest, which, in turn, increases the demand for money. This seems contradictory to the essence of monetary policy (Calza et al. 2001) and, therefore, most studies use interest rate spreads (the market rate minus the own rate or the bond rate minus the own rate) instead of just the own rate.

The expected change in price level enters a money demand function as the opportunity cost of holding money along with the nominal interest rate. Importantly, the change in price level affects the rate of return on the inventory of goods as high expected inflation induces economic agents to shift from money to goods, i.e., to stocking inventory, due to high profit incentives. However, Golineli and Pastorello (2002) do not include inflation as a measure of the opportunity cost of holding money in their long-run specification of the money demand function, arguing that it has no additional explanatory content compared to the long-term interest rate.

Other money demand specifications include the exchange rate (or depreciation) in order to capture the effect of currency substitution. Exchange rate depreciation can have a positive or negative effect on real balances (Bahmani-Oskooee & Rehman, 2010). A depreciation of the domestic currency increases the value of foreign assets in terms of the domestic currency, and if it is perceived as an increase in wealth, it will have a positive impact on real balances. However, if the depreciation enhances expectations of further depreciation, then currency substitutions may take place and reduce real money balances (Bahmani-Oskooee & Shin, 2002). The exchange rate variable is important in open economies' money demand function where claims denominated in foreign currency are high and currency conversion is prevalent.

Since the contribution of this paper is fairly empirical, after carefully identifying a theoretical specification with potential determinants of money demand (suggested both by a micro-founded model and the empirical literature), we attempt to search for a stable money demand function for Pakistan by applying various econometric techniques. These techniques include the residual-based cointegration approach, the autoregressive distributed lag (ARDL) modeling approach, the recursive Johansen cointegration approach, and the Bayesian estimation approach with Markov-Chain Monte-Carlo (MCMC) simulations. Appendix B outlines these econometric models.

4. Results and Discussion

This section provides and interprets estimation results based on the four methodologies mentioned in the previous section. The variables included in all the different empirical specifications are: nominal and real money (M2), the industrial production index (IPI), real GDP, the CPI, expected inflation, the weighted average deposit rate as the own rate of broad money, weighted average six-month market treasury bills (MTBs), the ten-year bond rate (federal investment bonds [FIBs] and Pakistan investment bonds [PIBs]), the short-term interest rate spread (six-month MTB minus deposit rate), long-term interest rate spread (ten-year bond rate minus deposit rate), the weighted average lending rate, the exchange rate, and the expected depreciation of the exchange rate.

A detailed description of each variable and its data source is provided in Table C1, Appendix C. All variables are expressed in log form except the interest rate, depreciation of exchange rate, and inflation rate.

The augmented Dickey-Fuller (ADF) test is used to check the stationarity of the variables (see Table C2, Appendix C).

4.1. Estimation Results of Engle-Granger Approach

In this approach, we use static ordinary least squares to estimate real and nominal money demand. A simple analysis shows that the income elasticity is greater than unity for both real and nominal money demand, suggesting a relatively high flow of money in the economy (see Appendix D, Table D1). This high income-elasticity could be due to structural changes in the economy that have resulted in large changes in currency and deposits (Adam et al., 2010).

In order to capture this structural change, we devise a principal component that includes the effect of five variables: services and manufacturing as a percent of GDP, imports and investment as a percent of GDP, government consumption as a percent of GDP, and credit to the private sector as a percent of broad money (M2)[7] (see Appendix H, Figures H1a to H1e). Income elasticity decreases with the inclusion of the first principal component, though marginally. However, its coefficient is not significantly different from 0, which signifies a limited structural change over this period[8] (see Appendix H, Figure H1f). On the other hand, the price elasticity of nominal money is close to unity, reflecting a one-to-one relationship between the GDP deflator and M2.

To measure the opportunity cost of money demand, we use the weighted average lending rate, which is also a weak proxy for the rate of return on alternative assets because it represents the interest rate on money. This variable is introduced in view of its availability for the entire period of analysis. The interest rate coefficient carries the correct sign but its magnitude is small. Inflation, representing opportunity cost, is also included in real balances (see Appendix D, Table D1, columns 3 and 4), and has a small effect. The residual from each equation is tested for stationarity using the ADF test and the null hypothesis of nonstationarity is rejected at the 1, 5, and 10 percent levels.

[7] The choice of these variables is based on their transaction-intensive nature and reflects changes in the demand for liquid services (Adam et al., 2010).
[8] The result may be different if we were to include principal components that reflect supply-side changes. During the last two decades, particularly since 2000, the reach of the financial sector in Pakistan has increased (in terms of more branches, the entrance of foreign banks, the privatization of public banks, internet banking, ATMS, the development of the equity and bond markets, and initiation of microfinance enterprises/banks); this has significantly reduced financial costs.

A dynamic error correction model is estimated using differenced variables (with two lags) and the lag of the estimated error term. The error correction term is highly significant and has the correct sign. Its large magnitude suggests faster adjustment toward long-run equilibrium (see Appendix D, Table D2). Real and nominal money have a strong inertial effect (Table D2, columns 1 to 4). The short-run effect of real output is large but only marginally significant, while the short-term dynamic effect of opportunity cost is not significantly different from 0.

We also estimate real and nominal money using quarterly data from 1992Q1 to 2011Q2 (see Appendix D, Table D3). Since quarterly data on real GDP is not available, we use the average of the monthly IPI instead. In the nominal demand function, the IPI coefficient is very low, signifying that it captures a small share of real output (around 15 to 18 percent). The coefficients of the CPI and lending rate are quite significant and have the expected signs. However, the estimated model is spurious and suffers from serial correlation and heteroskedasticity.

4.2. Cointegration Results Based on ARDL Approach

This approach considers seven alternative model specifications of the real money demand function. In all the estimated models, the coefficient of the scale variable has the correct sign and is significant, but its magnitude is less than unity as expected (see Appendix E, Table E1). In Model 1 (M-1 in Table E1), the inflation rate (representing opportunity cost) has the expected sign but a lower magnitude. The lending rate is used to proxy the rate of return on alternative assets; it has the correct sign but is not significantly different from 0.

During the early 1990s, the financial sector underwent many changes, including the introduction of foreign currency accounts. These provided an alternative avenue for parking money, explaining why the exchange rate has a role to play in money demand. In Model 2 (M-2 in Table E1), the money demand function is estimated using the exchange rate as an additional variable; it has the correct sign but is highly insignificant. The underlying reason for this insignificance could be the volatility of the exchange rate coupled with the freezing of foreign currency accounts in 1998 and huge inflows of foreign currency in 2001 onward. These interventions in the foreign exchange market may have dampened the impact of the exchange rate.

In most money demand equations, the short-term bond rate (government's MTBs) is used to signify an alternative to money holding.

Therefore, we include the weighted average rate of six-month T-bills (M-3 in Table E1) instead of the lending rate, which yields the expected sign but is insignificant; the rationale for this is that it was not used by economic agents until 2010.

With evolving financial sector reforms, particularly in the late 1990s, the initiation of fiscal and monetary coordination—though not very effective until now—and frequent monetary policy communication has changed the economic perspective of economic agents. These agents are now more rational and take into account the economic outlook when making an economic decision. Cognizant of this changing behavior, we augment the money demand function with expected inflation and the expected depreciation of the exchange rate as potential opportunity cost variables (see M-4 in Table E1). The long-run coefficients of both variables have the expected signs and are highly significant. This may also explain the central bank's tendency to intervene frequently in the foreign exchange market in order to stabilize the exchange rate.

Since the bulk of broad money is remunerative, the own rate of M2 cannot be ignored. The weighted averaged deposit rate is used to capture the own rate of M2 (see Appendix E, Table E1, M-5 and M-6). This not only has the incorrect sign, it is also insignificant in both models. Hypothetically, the coefficient of the own rate should be positively related to money demand, i.e., an increase in the own rate should increase the demand for money (Laidler, 1993). The underlying reason for the nonresponsiveness of the own rate is the sluggish movement of deposit rates in the lowest panel of the interest rates corridor, i.e., the deposit rate is very low (until recently when a minimum floor was introduced in 2005) and barely moves in tandem with other market interest rates (M. H. Khan & Khan, 2010).

The long-term bond market also provides an alternative avenue for the transaction demand for money. We therefore include the weighted average rate of FIBs and PIBs in our specification to examine its impact on money demand (see Appendix E, Table E1, M-6). The coefficient of the long-term bond rate has the expected sign but is not significantly different from 0. This may be due to the fact that the long-term bond market in Pakistan is shallow and restricted: only a few large banks and financial companies are allowed to transact government bonds.

Since most market interest rates move in tandem, and including each interest rate variable in the money demand function might complicate the model and make it difficult to interpret, we include only the short-term (six-month T-bills minus deposit rate) and long-term (ten-year FIBs/PIB rate minus deposit rate) spread to capture the effect of the own rate on broad money and opportunity cost (see Appendix E, Table E1, M-7). The coefficients of the interest rate spreads have the expected sign but are insignificant.

We also estimate the short-run dynamics of each model to examine the process of convergence toward a long-run path. The coefficient of the error correction term is significant and has the correct sign, reflecting a moderate pace of adjustment. Most of the variables in (see Appendix E, Table E2, M-1 to M-4) are significant at conventional levels and indicate a convergence toward equilibrium once they have deviated from the long-run path. On the other hand, the coefficients of the own rate, short-term rate, long-term rate, and interest rate spread are insignificant (see Appendix E, Table E2, M-5 to M-7).

4.3. Cointegration Results Based on Johansen's Approach

In the cointegration approach, we estimate unrestricted VAR models with various specifications of nominal and real money demand. At the outset, the variables are checked for unit roots; the ADF test confirms that all the variables are nonstationary at level and stationary at first difference, and are thus integrated of order one [I(1)] (see Appendix C, Table C2). For the model with nominal money as the dependent variable, the lag length selected under Schwarz's Bayesian criterion (SBC) and the Akaike information criterion (AIC) is p = 3 and p = 4, respectively. Using quarterly data from 1992Q1 to 2011Q4 and a lag length of 3 helps maintain parsimonious selection. On the other hand, in the models with real money, we select a lag length of 4 using the SBC (see Appendix F, Tables F1 and F2).

In Johansen's cointegration model, the long-run determinants of nominal M2 are the IPI, CPI, and lending rate; the determinants of real money are the IPI, expected inflation, and the lending rate. We capture the model's short-run dynamics by taking quarterly changes. This dynamism is introduced by incorporating past changes in each economic determinant in explaining M2 growth. In the short run, the determinants of nominal and real M2 growth are the last quarter's values for economic growth, inflation, and changes in interest rates. Further, past deviations of M2 from its stable long-run path are also incorporated as an explanatory variable.

The cointegration relationship is determined by using trace statistics and maximum eigenvalues. However, it is important to make certain assumptions regarding the deterministic trend specification and drift term before estimating the rank. The two specifications common in the literature are (i) a restricted intercept without a deterministic trend in a cointegration relationship and (ii) an unrestricted intercept with a linear deterministic trend in short-run equations. We use both specifications in our analysis.

Using the specification of an unrestricted intercept and linear trend in the cointegration equation for nominal money, we find that both the trace statistics and maximum eigenvalues show one cointegrating vector. Accordingly, the estimated long-run relationship for nominal money is:

$$\log(M2) = \underset{(0.15)}{1.21}*\log(IPI) + \underset{(0.27)}{1.88}*\log(CPI) - \underset{(0.006)}{0.03}*\text{lending rate}$$
$$-\underset{(0.006)}{0.02}*\text{trend} + 5.55$$

The coefficients of all the variables are significant (see standard errors in parentheses) and have the expected signs. Income elasticity has the same magnitude as expected, i.e., more than unity. The estimated empirical realization of the adjustment parameter ($\hat{\alpha}$) in the long-run equilibrium has the following values:

$$\hat{\alpha} = \begin{bmatrix} -0.150\ (0.049) \\ 0.351\ (0.129) \\ -0.075\ (0.019) \\ -3.785\ (1.161) \end{bmatrix}$$

The first element of the column shows the error correction parameter of the estimated money demand function, which indicates a rapid adjustment toward equilibrium following a shock. The adjustment parameter changes slightly when the short-run dynamic equations for nominal money and other variables are re-parameterized to estimate projections using parsimonious relationships. Each equation is then extended by incorporating seasonal variables and the volatility of oil prices; this incorporates the short-term effects of oil price changes and the seasonal demand for money.

Similarly, for real money balances, with the specification of an unrestricted intercept without a linear deterministic trend, both the trace and eigenvalue statistics indicate one cointegrating vector. Their long-run relationship is as follows:

log(real M2) = 0.88*log(IPI) – 0.08*exp(inflation) – 0.16*lending rate + 13.75

(0.21) (0.02) (0.04)

The coefficients of all the variables are significant (see standard errors in parentheses) and have the expected signs. The vector of the adjustment parameter is as follows:

$$\hat{\alpha} = \begin{bmatrix} -0.023(0.008) \\ -0.073\ (0.019) \\ -0.804\ (0.543) \\ 0.015\ (0.210) \end{bmatrix}$$

The error correction term for real money is very low, indicating slow adjustment toward long-run equilibrium. However, the error correction terms for IPI and expected inflation are relatively high, implying rapid adjustment. The ADF test indicates that the residuals of the short-run equations are stationary and normal.

A range of models of real money demand with different specifications is estimated (see Appendix F, Table F3). The lag length of the unrestricted VAR model is based on the SBC. In Models 1 to 3, we use the weighted average rate of six-month T-bills instead of the lending rate as it represents the short-term bond rate. The coefficient of the MTB rate is highly significant but has the incorrect sign. In Model 4, we include the exchange rate variable; the trace statistics indicate two cointegrating equations. Although all the long-run coefficients are significantly different from 0, their effects are contrary to theory.

In Model 5, the MTB rate is replaced with the lending rate as an opportunity cost variable and the exchange rate included. The coefficient of the opportunity cost variable is highly significant and has the expected sign. However, the effect of the exchange rate is more pronounced, i.e., a one-percent increase in the exchange rate (depreciation) will increase real money by more than 9 percent. Nevertheless, the positive expected inflation contradicts the theory.

Model 6 includes a risk premium variable, which is the difference between the lending rate and risk-free rate (MTB rate). The coefficient of the risk premium has the expected sign but is insignificant at conventional levels. This model is extended by including the exchange rate (Model 7): in this case, all the variables, except the exchange rate, are significantly different from 0 and have the expected signs. The adjustment parameters for real money demand in all these models are either very low or explosive, signifying little or no convergence to a long-run equilibrium path.

4.4. Bayesian Estimation Results

The Bayesian estimation approach uses prior information on key structural parameters before taking the model and data to the simulation stage. According to Canova (2007), these reflect researchers' confidence about the likely location of the model's structural parameters. In practice, priors are chosen based on observation, fact, and the empirical literature.

For our study, two parameters, α (the share of capital in the production function) and β (the subjective discount factor) are fixed at 0.46 and 0.99. The parameter value of the discount factor (β) is set in order to obtain the historical mean of the nominal interest rate in the steady state. Following Haider and Khan (2008), the degree of price stickiness (θ) is assumed to be 0.74. This value is consistent with the latest survey-based finding on firms' optimal pricing behavior in Pakistan (see Choudhary, Naeem, Faheem, Hanif, & Pasha, 2011). The elasticity parameter of money demand with respect to output (φ_1) is taken as 0.86 whereas with respect to the interest rate (φ_2) it is –0.018. These values are consistent with ARDL long-run estimates (see Appendix E, Table E1). Prior information about other selected parameters is given in Table G1, Appendix G.

After selecting the priors, we apply Bayesian simulation algorithms by combining the likelihood distribution, which leads to an analytically intractable posterior density. In order to sample from the posteriors, we use a random walk Metropolis-Hastings algorithm to generate 150,000 draws from the posteriors. We report the posterior results (parameter estimates) in the second column of Table G2 in Appendix G. Figures G1 and G2 give the kernel estimates of the priors and the posteriors of each parameter. The results show that the prior and posterior means are not far from each other. To some extent, this confirms the stability of the money demand parameter estimates. However, we also use various formal tests of parameter stability, the results of which are discussed in the next section.

4.5. Parameter Stability Tests

As we observed at the estimation stage, the elasticity parameter of money demand with respect to the interest rate is sensitive to alternative specifications of money demand. Therefore, we need to test parameter stability over time for which we select the best models from each individual approach. Ideally, parameter stability should be checked using different methods, including the empirical realization of income elasticity and opportunity cost variables with a changing sample period or recursive estimations of the coefficients of each model. To verify the models' stability, we apply the CUMUS and CUMUS-square of the residual.

We first select Model 2 in the cointegrated VAR specification. This model is re-estimated using Johansen's procedure with a lag length of 4 and trend specification but changing the sample period, i.e., the end of each fiscal year from 2004 to 2011. The results of this recursive estimation show that income elasticity and the opportunity cost coefficients change significantly—both in magnitude and sign—with the changing sample period, signifying an unstable money demand function (see Appendix F, Table F5).

However, when the same procedure is applied to the extended model (Model 7), the income elasticity and opportunity cost coefficients display the correct signs and are also steady over time. This confirms that the money demand function with this specification is stable. The trace statistics and maximum eigenvalues posit one or two cointegrating relationships. Figures H3a and H3b in Appendix H show the cointegration relation estimated with Johansen's procedure using Models 2 and 7, where the former signifies instability and the latter indicates stability. The mean-reverting properties are evident in the graphical representation.

Parameter consistency can also be checked by recursively estimating the coefficient of the money demand function, where coefficient of each variables ($\hat{\beta}_{rt}$) estimated by adding more data to the equation (flexible window). Figures H4a to H4d (Appendix H) show the recursively estimated log-run coefficient of Model 7 using Johansen's cointegration method. To avoid the large uncertainty associated with the initial estimates, just slightly less than half the estimates for each coefficient are displayed (from FY2006Q3 to FY2011Q2).

Figure H4a depicts stable income elasticity (around 0.7 percent) over the changing sample period. Figure H4b shows the recursive estimates of expected inflation: its coefficient remains highly stable during the initial period, but later shows significant variation and jumps markedly in FY2009. This was the period when economic conditions weakened considerably due to an uncertain domestic environment coupled with the external financial crisis and an unexpected increase in international commodity and energy prices that seeped into high domestic inflation. On the other hand, the recursive estimates of the exchange rate and risk premium remain more or less stable with slight variation in FY2009 and onward. We also recursively estimate the coefficients of the other cointegration model, but they are not stable and display large variations.

The models' stability is also tested using the CUSUM of both the residual and squared residual. We estimate the residuals recursively for all the ARDL models, where the CUSUM of the residual for Model 4 lies within a 5 percent significance level (see Appendix H, Figure H2). It is important to note that the ARDL model is slightly different from Model 7 (Johansen's cointegration): the exchange rate variable in the latter is replaced with the expected depreciation of the exchange rate as it is stationary at level and the ARDL procedure can be applied irrespective of a zero or one order of integration. The risk premium variable is replaced with the lending rate. Both variables in the ARDL model represent the opportunity cost of money holding.

Finally, the parameter stability of money demand in the NK-DSGE model specification is assessed using the global sensitivity toolkit. This toolkit uses the Smirnov test of stability, which shows the significance of individual parameters for the whole model. The cumulative plots for stability and instability behavior provide us with useful information on the fitness of each structural parameter. Figures H5 and H6 (see Appendix H) show that all the structural parameters of money demand are stable and properly fitted with respect to the data.

5. Conclusion

This study has attempted to investigate the money demand function for Pakistan, where monetary aggregates are considered a nominal anchor. Importantly, monetary aggregates as an operational and intermediate target have contributed significantly to the implementation and communication of monetary policy, but the stability of money demand has long been debated, given evolving financial innovations and regulations. Earlier studies on the subject have provided conflicting explanations due to the use of inadequate specifications and imprecise empirical estimations of the money demand function.

This study finds that money demand in Pakistan is stable if correctly specified (see ARDL Model 4, unrestricted cointegrated VAR Model 7, and NK-DSGE model) and concludes that monetary aggregates should remain, if not the primary, the secondary targets in the monetary policy framework. Although financial innovations have changed the preferences of economic agents (money holders) in developed countries, they have had a limited impact on economic behavior in developing countries such as Pakistan.

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Appendix A

Theoretical Foundations of the Money Demand Function

This section deals with the micro-foundations of the standard money demand function. To derive the theoretical specification, we consider a typical new Keynesian DSGE model as given in Galí (2008) and Walsh (2010). The model setup starts with households' optimal decision to maximize their intertemporal utility function subject to a lifetime budget constraint. The utility function depends on consumption,, leisure, , and real money balances . Firms, on the other hand, produce output by using labor as a standard input of production and attempt to maximize profit in a monopolistic competitive fashion. Other agents include the central bank, which conducts monetary policy, and the government, which deals with fiscal policy-related issues.

Since we are interested in the theoretical specification of money demand, we start with household optimization. The lifetime utility function is:

$$U_0 = E_t \sum_{i=0}^{\infty} \beta^i u \left[C_{t+i}, (1 - l_{t+i}), \frac{M_{t+i}}{P_{t+i}} \right]$$

where is the subjective discount factor. For analytical simplicity, the utility function is assumed to be separable and is specified as:

$$u \left[C_{t+i}, (1 - l_{t+i}), \frac{M_{t+i}}{P_{t+i}} \right] = \frac{(C_{t+i})^{1-\gamma}}{1-\gamma} + \frac{(1-l_{t+i})^{1-\upsilon}}{1-\upsilon} + \frac{\xi_{m,t+i}}{1-\omega} \left(\frac{M_{t+i}}{P_{t+i}} \right)^{1-\omega}$$

where is the parameter of risk aversion, is the inverse elasticity of labor supply, is the inverse interest elasticity of money demand, and is a stochastic shock to money demand. The other usual assumptions regarding the utility function are: ∂∂ and ∂∂, which implies that it is increasing but diminishing over time in each case. Households want to maximize this utility function subject to the following budget constraint:

$$C_t + \frac{M_t}{P_t} + \frac{B_t}{P_t} = \frac{W_t}{P_t} l_t + \frac{M_{t-1}}{P_t} + (1 + i_{t-1}) \frac{B_{t-1}}{P_t} + \Pi_t$$

where denotes the general price level, is interest-bearing assets with as the nominal gross return on assets at time , is nominal labor income, and is real dividends.

The household optimization process solves the following problem as:

$$\mathcal{L} = E_t \sum_{i=0}^{\infty} \beta^i \left[\begin{array}{c} \left\{ \frac{(C_{t+i})^{1-\gamma}}{1-\gamma} + \frac{(1-l_{t+i})^{1-\upsilon}}{1-\upsilon} + \frac{\xi_{m,t+i}}{1-\omega} \left(\frac{M_{t+i}}{P_{t+i}} \right)^{1-\omega} \right\} \\ +\lambda_{t+i} \left\{ \begin{array}{c} \frac{W_{t+i}}{P_{t+i}} l_{t+i} + \frac{M_{t-1+i}}{P_{t+i}} + (1+i_{t-1+i}) \frac{B_{t-1+i}}{P_{t+i}} + \Pi_{t+i} \\ -C_{t+i} + \frac{M_{t+i}}{P_{t+i}} + \frac{B_{t+i}}{P_{t+i}} \end{array} \right\} \end{array} \right]$$

where is the Lagrange multiplier associated with the household budget constraint. The solution to this optimization process yields the following first-order conditions (FOCs):

$$C_t^{-\gamma} = \lambda_t$$

$$\xi_{m,t} \left(\frac{M_t}{P_t} \right)^{-\omega} = \lambda_t \left(1 - \beta E_t \frac{\lambda_{t+1}}{\lambda_t} \frac{P_t}{P_{t+1}} \right)$$

$$(1 - l_t)^{-\upsilon} = \frac{W_t}{P_t} \lambda_t$$

$$1 = \beta (1 + i_t) E_t \left(\frac{\lambda_{t+1}}{\lambda_t} \frac{P_t}{P_{t+1}} \right)$$

We define the gross inflation rate as . Solving the above FOCs simultaneously yields two important results.

The first is the intertemporal Euler equation of consumption:

$$1 = \beta \frac{(1 + i_t)}{(1 + \pi_{t+1})} E_t \left(\frac{C_t}{C_{t+1}} \right)^{\gamma}$$

The second is the nonlinear money demand function:

$$\xi_{m,t} \left(\frac{M_t}{P_t} \right)^{-\omega} = C_t^{-\gamma} \left(\frac{i_t}{1 + i_t} \right)$$

The economy-wide aggregate resource constraint in the model can be written as Therefore, the final nonlinear version of the money demand function can be simplified as:

$$\xi_{m,t} \left(\frac{M_t}{P_t} \right)^{-\omega} = Y_t^{-\gamma} \left(\frac{i_t}{1 + i_t} \right)$$

For empirical estimation purposes, we consider the linear version of this money demand function. It can be log-linearized around the deterministic steady-state using first-order Taylor approximation and any linearized variable can be defined as $\tilde{x}_t = \log(x_t) - \log(\bar{x})$.

$$1 + \tilde{\xi}_{m,t} + (-\omega)\widetilde{\left(\frac{M_t}{P_t}\right)} = 1 - \gamma\widetilde{Y}_t + \tilde{\imath}_t$$

or

$$\widetilde{\left(\frac{M_t}{P_t}\right)} = \frac{\gamma}{\omega}(\widetilde{Y_t}) - \frac{1}{\omega}(\tilde{\imath}_t) + \frac{1}{\omega}(\tilde{\xi}_{m,t})$$

If we consider, $\varphi_1 = \frac{\gamma}{\omega}$, $\varphi_2 = -\frac{1}{\omega}$ and $\tilde{\varepsilon}_{m,t} = \frac{1}{\omega}(\tilde{\xi}_{m,t})$, where, φ_1 is the output elasticity of money demand according to the restriction on deep parameters, it appears to be positive while is negative. This signifies that money demand is positively related to output and negatively to nominal interest rates. The final simplified linear version of money demand with the inclusion of an interest term, for econometric consideration, is given as:

$$\widetilde{\left(\frac{M_t}{P_t}\right)} = \varphi_0 + \varphi_1(\widetilde{Y_t}) + \varphi_2(\tilde{\imath}_t) + \tilde{\varepsilon}_{m,t}$$

Appendix B

Econometric Modeling Setups

This study estimates the money demand for Pakistan with different specifications in light of a theoretical micro-founded model and the empirical literature. Accordingly, we apply various econometric techniques to analyze whether money demand is stable. More specifically, we consider the residual-based cointegration approach (Engle & Granger, 1987), the ARDL approach (Pesaran & Shin, 1995; Pesaran, Shin, & Smith, 2001), the recursive Johansen cointegration approach (Johansen, 1988; Johansen & Juselius, 1990), and the Bayesian estimation approach with MCMC simulations (Canova, 2007).

B.1. Engle-Granger Modeling Setup

From the theoretical model derived in Section 3, we have the following standard linear specification of the money demand function:

$$\widetilde{\left(\frac{M_t}{P_t}\right)} = \varphi_0 + \varphi_1(\widetilde{Y_t}) + \varphi_2(\tilde{\imath}_t) + \tilde{\varepsilon}_{m,t}$$

This specification suggests that real money demand depends on real output and the nominal interest rate, where φ_1 and φ_2 are the elasticity parameters of money demand with respect to output and the interest rate. However, in practice, econometricians have attempted to estimate both nominal and real versions of the money demand function. The nominal version considers actual inflation an explanatory variable while the real money demand function takes expected inflation as a possible determinant to tackle future expectations. The literature argues that, due to structural changes in currency and deposits in developing economies, the income elasticity of money demand will be high (see, for instance, Adam et al., 2010).

In order to tackle this possible empirical issue, we construct an index of structural change based on the principal component technique. This index includes the effect of five variables: services and manufacturing as a percent of GDP, imports and investment as a percent of GDP, government consumption as a percent of GDP, and credit to the private sector as a percent of broad money (M2). This variable is then augmented as an explanatory variable in both regressions of money demand. Thus, we have the following four specifications:

Nominal money demand:

$$\widetilde{(M_t)} = \varphi_0^{S1} + \varphi_1^{S1}(\widetilde{Y_t}) + \varphi_2^{S1}(\tilde{\imath}_t) + \varphi_3^{S1}(\tilde{\pi}_t) + \tilde{\varepsilon}_{M,t}^{S1} \quad \text{(EG – S1)}$$

$$\widetilde{(M_t)} = \varphi_0^{S2} + \varphi_1^{S2}(\widetilde{Y_t}) + \varphi_2^{S2}(\tilde{\imath}_t) + \varphi_3^{S2}(\tilde{\pi}_t) + \varphi_4^{S2}(\widetilde{StructChange_t}) + \tilde{\varepsilon}_{M,t}^{S2} \quad \text{(EG – S2)}$$

Real money demand:

$$\widetilde{\left(\frac{M_t}{P_t}\right)} = \varphi_0^{S3} + \varphi_1^{S3}(\widetilde{Y_t}) + \varphi_2^{S3}(\tilde{\imath}_t) + \varphi_3^{S3}(\tilde{\pi}_t^e) + \tilde{\varepsilon}_{m,t}^{S3} \quad \text{(EG – S3)}$$

$$\widetilde{\left(\frac{M_t}{P_t}\right)} = \varphi_0^{S4} + \varphi_1^{S4}(\widetilde{Y_t}) + \varphi_2^{S4}(\tilde{\imath}_t) + \varphi_3^{S4}(\tilde{\pi}_t^e) + \varphi_4^{S4}(\widetilde{StructChange_t}) + \tilde{\varepsilon}_{m,t}^{S4} \quad \text{(EG – S4)}$$

Where, as usual, [φ_1^{S1}, φ_1^{S2}, φ_1^{S3}, and φ_1^{S4}] >0 and [φ_2^{S1}, φ_2^{S2}, φ_2^{S3}, and φ_2^{S4}] <0 are the long-run elasticities of income and the interest rate, and can be estimated using static ordinary least squares (SOLS).

In order to test the long-run dynamics, we need to check the stationarity of each variable. Once we have confirmed that the selected series are integrated and of the same order, we apply the Engle-Granger test of cointegration. This test checks the stationarity of the estimated residuals of each specification. If an estimated residual is stationary at level, we can conclude that long-run dynamics exist. Using the Granger representation theorem, any cointegrated regression can be mapped onto its error correction mechanism (ECM), which deals with short-run dynamics.

In general, the ECM version of any of the above specifications can be written as:

$$\Delta\widetilde{z_t} = \alpha + \beta\widetilde{ECM}_{t-1} + \sum_{i=1}^{p-1} \gamma_i \Delta\widetilde{z_{t-\iota}} + \sum_{i=1}^{p-1} \Psi_i \Delta\widetilde{Y_{t-\iota}} + \tilde{\varepsilon}_{z,t}$$

where $\beta \in (-1,0)$, is the short-run correction parameter. One important shortcoming of the Engle-Granger cointegration test based on SOLS is that the estimates have an asymptotic distribution, that is, they are generally non-Gaussian and exhibit a large sample bias. Since conventional testing procedures are not valid unless modified substantially, SOLS is generally inappropriate for making inferences about the cointegrating vector. For this reason, we use ARDL VAR models for cointegration and Bayesian simulation approaches.

B.2. ARDL Modeling Setup

Pesaran and Shin (1995) and Pesaran et al. (2001) have developed a bounds test using ARDL to find the long-run relationship between variables irrespective of their order of integration.[1] This technique does not require pretesting the unit roots of the variables as the ADF unit root test become redundant in the presence of structural breaks. It allows us to estimate the cointegration relationship with OLS once the lag order to the model is identified. Therefore, in general, we can write the VAR(p) model as:

$$\widetilde{Y}_t = b + ct + \sum_{i=1}^{p} \Phi_i \widetilde{Y}_{t-i} + \tilde{\varepsilon}_{Y,t}$$

where $\widetilde{Y}_t$ represents a vector of variables. The above expression can be written as a simple VECM:

$$\Delta\widetilde{Y}_t = B + Ct + \Theta\widetilde{Y}_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta\widetilde{Y}_{t-i} + \tilde{\varepsilon}_{Y,t}$$

Where, $\Theta = -(\mathrm{I}_{k+1} - \sum_{i=1}^{p} \Phi_i)$ and $\Gamma_i = -\sum_{j=i+1}^{p} \Phi_j$, *i*=1,…,*p*-1; are the (*k*+1)×(*k*+1) matrices of the long run multipliers and the short run dynamic coefficients. By making the assumption that there is only one long run relationship amongst the variables, Pesaran and Shin (1995) and Pessaran *et al.,* (2001) focus on the VECM equation and partition $\widetilde{Y}_t$ into a dependant variable, $\widetilde{z}_t$ and a set of forcing variables, $\widetilde{X_t}$. Under such conditions, the matrices B, C, Γ and, most importantly, Θ, the long run multiplier matrix can also be partitioned conformably with the partitioning of $\widetilde{Y}_t$.

$$\Theta = \begin{pmatrix} \theta_{11} & \theta_{12} \\ \theta_{21} & \theta_{22} \end{pmatrix} \quad B = \begin{pmatrix} b_1 \\ b_2 \end{pmatrix} \qquad C = \begin{pmatrix} c_1 \\ c_2 \end{pmatrix} \qquad \Gamma = \begin{pmatrix} \gamma_{11} & \gamma_{12} \\ \gamma_{21} & \gamma_{22} \end{pmatrix}$$

The key assumption, that $\widetilde{X_t}$ is long run forcing for $\widetilde{z}_t$, then implies that the vector, θ_{21}=0, that is that there is no feedback from the level of $\widetilde{z}_t$ on $\Delta\widetilde{X_t}$. As a result, the conditional model for $\Delta\widetilde{z}_t$ and $\Delta\widetilde{X_t}$ can be written as:

$$\Delta\widetilde{z}_t = b_1 + c_1 t + \theta_{11}\tilde{z}_{t-1} + \theta_{12}\tilde{X}_{t-1} + \sum_{i=1}^{p-1} \gamma_{11,i}\Delta\tilde{z}_{t-i} + \sum_{i=1}^{p-1} \gamma_{12,i}\Delta\tilde{X}_{t-i} + \tilde{\varepsilon}_{1,t}$$

[1] This technique is applicable to variables that are either stationary at level or integrated of order one. However, the relationship becomes explosive when one or all the variables are of high order, i.e., I(2).

$$\Delta\widetilde{X_t} = b_2 + c_2 t + \theta_{22}\tilde{X}_{t-1} + \sum_{i=1}^{p-1}\gamma_{21,i}\Delta\tilde{z}_{t-i} + \sum_{i=1}^{p-1}\gamma_{22,i}\Delta\tilde{X}_{t-i} + \tilde{\varepsilon}_{2,t}$$

Under standard assumptions about the error terms in above representations,[2] Pessaran *et al.,* (2001) re-write the system as:

$$\Delta\widetilde{z_t} = \alpha_1 + \alpha_2 t + \phi\tilde{z}_{t-1} + \delta\tilde{X}_{t-1} + \sum_{i=1}^{p-1}v_i\Delta\tilde{z}_{t-i} + \sum_{i=1}^{p-1}\varphi_i\Delta\tilde{X}_{t-i} + \tilde{\varepsilon}_t$$

They term this an unrestricted error correction model. Note that, in this expression, a long-run relationship will exist among the level variables if the two parameters ϕ and δ are both non-zero, in which case, for the long-run solution, we obtain:

$$\widetilde{z_t} = -\frac{\alpha_1}{\phi} - \frac{\alpha_2}{\phi}t - \frac{\delta}{\phi}\tilde{X}_t$$

Pesaran et al. (2001) choose to test the hypothesis of no long-run relationship between $\widetilde{z_t}$ and $\widetilde{X_t}$ by testing the joint hypothesis that $\phi = \delta = 0$ in the context of the above VECM. The test they develop is a bounds-type test with a lower bound calculated on the basis that the variables in $\widetilde{X_t}$ are I(0) and an upper bound calculated on the basis that they are I(1). They provide critical values for this test from an extensive set of stochastic simulations under differing assumptions regarding the appropriate inclusion of deterministic variables in the ECM.

If the calculated test statistic (which is a standard F-test for testing the null hypothesis that the coefficients on the lagged levels' terms are jointly equal to 0) lies above the upper bound, the result is conclusive and implies that a long-run relationship does exist between the variables. If the test statistic lies within the bounds, no conclusion can be drawn without prior knowledge of the time series properties of the variables. In this case, standard methods of testing have to be applied. If the test statistic lies below the lower bound, no long-run relationship exists.

We estimate the money demand function based on the following general form:

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[2] Essentially that they are independently normally distributed with a positive definite variance covariance matrix.

$$\Delta\widetilde{\left(\frac{M_t}{P_t}\right)} = \alpha_1 + \alpha_2 t + \phi\widetilde{\left(\frac{M_{t-1}}{P_{t-1}}\right)} + \delta\tilde{X}_{t-1} + \sum_{i=1}^{p-1}\upsilon_i\Delta\widetilde{\left(\frac{M_{t-\iota}}{P_{t-\iota}}\right)} + \sum_{i=1}^{p-1}\varphi_i\Delta\tilde{X}_{t-i} + \tilde{\varepsilon}_t$$

where $\widetilde{X_t}$ is a set of exogenous variables. We attempt seven alternative specifications of the ARDL model to test the underlying stability hypothesis. The main objective is to correctly identify the real money demand function and to determine the long-run relationship vis-à-vis the short-run dynamic. The vector of exogenous variables in each specification is listed below:

$\widetilde{X_t}$ = [*industrial output; weighted avg. lending rate; inflation*] (ARDL-S1)

$\widetilde{X_t}$ = [*industrial output; weighted avg. lending rate; inflation; exch. rate*] (ARDL-S2)

$\widetilde{X_t}$ = [*industrial output; six-month T-bill rate; expected inflation*] (ARDL-S3)

$\widetilde{X_t}$ = [*industrial output; weighted avg. lending rate; expected inflation; exch. rate*] (ARDL-S4)

$\widetilde{X_t}$ = [*industrial output; own rate; six-month T-bill rate; inflation*] (ARDL-S5)

$\widetilde{X_t}$ = [*industrial output; own rate; six-month T-bill rate; ten-year bond rate; inflation*] (ARDL-S6)

$\widetilde{X_t}$ = [*industrial output; risk premium; expected inflation; exch. rate*] (ARDL-S7)

These specifications are applied to quarterly data from 1991Q1 to 2011Q4 for the following variables: the IPI as a scale variable, the opportunity cost variable CPI inflation (QoQ), the exchange rate, the variant interest rate, i.e., the lending rate, the six-month treasury bond rate, the long-term rate (weighted average of ten-year FIBs/PIBs), the short-term interest rate spread, and the long-term interest rate spread. We also consider the weighted average of the deposit rate as the own rate of broad money (M2). The SBC is applied to select a lag order for each model.

B.3. Johansen Modeling Setup

This setup is an extension of the Engle-Granger modeling framework in a VAR fashion, and also follows the assumption that all the series are integrated of the same order. The starting point is similar to that of the ARDL setup with the construction of a VAR(p) model as follows:

$$\widetilde{Y_t} = \boldsymbol{b} + \boldsymbol{c}t + \sum_{i=1}^{p}\Phi_i\widetilde{Y}_{t-i} + \tilde{\varepsilon}_{Y,t}$$

where $\widetilde{Y_t}$ represents a vector of variables of the same order of integration—let us say, I(1). The above expression can be written as a simple VECM:

$$\Delta\widetilde{Y}_t = \boldsymbol{B} + \boldsymbol{C}t + \boldsymbol{\Theta}\widetilde{Y}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta\widetilde{Y}_{t-i} + \tilde{\varepsilon}_{Y,t}$$

Where, $\boldsymbol{\Theta} = -(\mathrm{I}_{k+1} - \sum_{i=1}^{p} \Phi_i)$ and $\Gamma_i = -\sum_{j=i+1}^{p} \Phi_j$, i=1,…,p-1; are the (k+1)×(k+1) matrices of the long run multipliers and the short run dynamic coefficients. Since, $\Delta\widetilde{Y}_{t-i}$, …, $\Delta\widetilde{Y}_{t-i+1}$ are I(0) but $\widetilde{Y}_{t-1}$, is I(1). So in order to do this equation consistent, $\boldsymbol{\Gamma}_i$ should not be of full rank. Consider its rank is r<(k+1). We can decompose matrix, $\boldsymbol{\Theta} = \boldsymbol{\alpha\beta}'$, where, $\boldsymbol{\alpha}$ is (k+1) x r matrix of error correction terms (speed of adjustment parameters) and $\boldsymbol{\beta}'$ is r x (k+1) matrix of coefficients of co-integrating vector. Now, we need to estimate two residual series, first by regressing, $\Delta\widetilde{Y}_t$ on $\sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta\widetilde{Y}_{t-i}$ and name it $\xi_{o,t}$ and second by regressing $\widetilde{Y}_t$ on $\sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta\widetilde{Y}_{t-i}$ and name it, $\xi_{i,t}$. Now, in next step, we form a regression on estimated residuals as: $\xi_{o,t} = \boldsymbol{\alpha\beta}'\xi_{i,t} + \varsigma_t$. The variance/co-variance matrix of this regression can be written as:

$$\Omega = \begin{pmatrix} S_{00} & S_{01} \\ S_{10} & S_{11} \end{pmatrix}$$

where S_{11} is the sum of the squares of ξ, S_{00} is the sum of the squares of ξ and S_{01} is the sum of the product of ξ and ξ. It is important to note that OLS is not applicable to estimating VAR because of cross-equation restrictions. We therefore use the maximum likelihood technique, where the maximum of the likelihood function is obtained by solving the eigenvalue problem:

$$\left|S_{10}S_{00}^{-1}S_{01} - \lambda S_{11}\right| = 0$$

This is equivalent to finding the eigenvalue of . We can then obtain the eigenvalues that are the root of this equation. These represent the γ canonical correlation between ξ and ξ. Thus, the maximum of the likelihood function is given by:

$$-2logLmax\left\{\alpha N \sum_{i=1}^{k+1} \ln{(1-\lambda_i)}\right\}$$

Using this expression, we can define the trace test statistic and maximum eigenvalue as:

$$\lambda_{trace} = -N \sum_{i=1}^{k+1} \ln{(1-\widehat{\lambda}_\iota)}$$

and

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

The null hypothesis of the maximum eigenvalue test is γ cointegrating vector(s), whereas the alternative hypothesis is γ + 1 cointegrating vector(s).

To apply this modeling procedure to the estimation and stability of the money demand function, we consider seven alternative specifications:

$\widetilde{Y}_t$ = [*real.m2;Industiral.Output; Weighted Avg.Ledning Rate; Inflation*] (Johansen-S1)

$\widetilde{Y}_t$ = [*real.m2;Industiral.Output; Weighted Avg.Ledning Rate; Expected Inflation*] (Johansen-S2)

$\widetilde{Y}_t$ = [*real.m2;Industiral.Output; 6-month.TBill.Rate; Expected Inflation*] (Johansen-S3)

$\widetilde{Y}_t$ = [*real.m2;Industiral.Output; 6-month.TBill.Rate; Expected Inflation;Exch.Rate*] (Johansen-S4)

$\widetilde{Y}_t$ = [*real.m2;Industiral.Output; Weighted Avg.Ledning Rate; Inflation,Exch.Rate*] (Johansen-S5)

$\widetilde{Y}_t$ = [*real.m2;Industiral.Output; Risk Premium; Inflation*] (Johansen-S6)

$\widetilde{Y}_t$ = [*real.m2;Industiral.Output; Risk Premium; Expected Inflation;Exch.Rate*] (Johansen-S7)

To estimate the short-run dynamics of the money demand function, we extract VECMs for each specification for which the dependent variable is real money demand as follows:

$$\Delta\widetilde{\left(\frac{M_t}{P_t}\right)} = \alpha_1 + \alpha_2 t + \phi\widetilde{\left(\frac{M_{t-1}}{P_{t-1}}\right)} + \delta\tilde{X}_{t-1} + \sum_{i=1}^{p-1}\upsilon_i\Delta\widetilde{\left(\frac{M_{t-\iota}}{P_{t-\iota}}\right)} + \sum_{i=1}^{p-1}\varphi_i\Delta\tilde{X}_{t-i} + \tilde{\varepsilon}_t$$

B.4. Bayesian Modeling Setup

In order to estimate the money demand function using the Bayesian estimation approach, we consider the complete new Keyensian monetary model in log-linearzed form as derived in Section 3.[3] This model consists of six structural equations: (i) forward-looking IS equation, (ii) new Keynesian Phillips curve, (iii) production function, (iv) money demand equation, (v) evolution of the natural rate of interest, and (vi) Taylor-type monetary policy rule. Furthermore, three stochastic shocks—a productivity shock, money demand shock, and monetary policy shock—are taken into account. The outline of the NK model is given as follows:

$$\tilde{y}_t = E(\tilde{y}_{t+1}) - \frac{1}{\gamma}(\tilde{\imath}_t - E\tilde{\pi}_{t+1} - \tilde{r}_t^n) \quad \text{[NKIS]}$$

$$\tilde{\pi}_t = \beta E(\tilde{\pi}_{t+1}) + \kappa\tilde{y}_t \quad \text{[NKPC]}$$

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[3] For simplicity, we have not provided the micro-foundations of the supply side (firms and the cost channel of monetary policy). These derivations are available on request.

Where, $\kappa = \frac{(1-\theta)(1-\beta\theta)}{\theta}$ is slope of NKPC.

$\tilde{r}_t^n = \rho + \gamma\psi(\rho_a - 1)\tilde{a}_t$ [Eq. of Natural Rate of interest]

$\tilde{y}_t = \tilde{a}_t + (1-\alpha)\tilde{l}_t$ [Eq. of production function]

$\widetilde{m}_t = \varphi_1(\widetilde{Y_t}) + \varphi_2(\tilde{\imath}_t) + \tilde{\varepsilon}_{m,t}$ [Eq. of real money demand]

$\tilde{\imath}_t = \rho + \chi_1(\widetilde{\pi_t}) + \chi_2(\tilde{Y}_t) + \tilde{\varepsilon}_{i,t}$ [Monetary policy rule]

$\tilde{a}_t = \rho_a \tilde{a}_{t-1} + \xi_{a,t}$ [Productivity Shock]

$\tilde{\varepsilon}_{m,t} = \rho_m \tilde{\varepsilon}_{m,t-1} + \xi_{m,t}$ [Money demand shock]

$\tilde{\varepsilon}_{i,t} = \rho_i \tilde{\varepsilon}_{i,t-1} + \xi_{i,t}$ [Monetary policy Shock]

Following Canova (2007), we try to fit this model, which consists of placing a prior distribution on the structural parameters, , the estimates of which are then updated using the data according to the Bayes rule:

$$p\left(\frac{\Gamma}{Y^T}\right) = \frac{p\left(\frac{Y^T}{\Gamma}\right)}{p(Y^T)} \propto L\left(\frac{\Gamma}{Y^T}\right)p(\Gamma)$$

where $p\left(\frac{Y^T}{\Gamma}\right) = L\left(\frac{\Gamma}{Y^T}\right)$ is the likelihood function, $p\left(\frac{\Gamma}{Y^T}\right)$ is the posterior distribution of parameters, and is the marginal likelihood defined as:

$$p(Y^T) = \int p\left(\frac{Y^T}{\Gamma}\right)p(\Gamma)d\Gamma$$

Any NK model will form a linear system with rational expectations, the solution to which takes the form

$$R_t = B_1(\Gamma)R_{t-1} + B_2(\Gamma)\mu_t$$

$$\mu_t = B_3(\Gamma)\mu_{t-1} + B_4(\Gamma)\varepsilon_t$$

where R_t is a vector of endogenous variables, μ_t is a vector of stochastic disturbances, and ε_t is a vector of innovations to stochastic shocks and coefficient matrices A_i depending on the parameters of the model.

The measurement equations linking the observable variables used in the estimation with the endogenous variables can be written as , where *C* is the deterministic matrix. The equations' expressions form the state-space representation of the model, the likelihood of which can be evaluated using the Kalman filter. An analytical solution to the whole system may not be obtainable in general, but the sequence of posterior draws can be obtained using the MCMC simulation methodology (see Gelman, Carlin, Stern, & Rubin, 2006; Koop, Poirier, & Tobias, 2007). Finally, the random walk Metropolis-Hastings algorithm is used to generate Morkov chains for the model parameters.

After obtaining the Bayesian estimation results, we use the global sensitivity analysis (GSA) toolkit[4] to assess the fitness and stability of the model's structural parameters. This toolkit consists of MATLAB routines, which use the Smirnov test for stability analysis. Ratto (2007) provides a detailed discussion on using this toolkit with various applied examples.

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[4] http://www.dynare.org

Appendix C

Results

Table C1: Variables and data sources

Variables	Source
Nominal M2	Statistical Bulletins, SBP (various issues)
Real GDP (Annual data)	Pakistan Economic Survey, MOF (various issues)
Industrial production index (IPI)	Statistical Bulletins, SBP (various issues)
Consumer price index (CPI)	Statistical Bulletins, SBP (various issues)
Inflation (percent change in CPI)	Statistical Bulletins, SBP (various issues)
Exchange rate	Statistical Bulletins, SBP (various issues)
ER App/dep (+/-)	Statistical Bulletins, SBP (various issues)
Lending rate	Statistical Bulletins, SBP (various issues)
Own rate (wt. avg. deposit rate)	Statistical Bulletins, SBP (various issues)
Short-term rate (wt. avg. 6-month MTB)	Statistical Bulletins, SBP (various issues)
Long-term rate (wt. avg. 10-year FIB/PIB)	Statistical Bulletins, SBP (various issues)
Risk premium (lending rate - MTB)	Statistical Bulletins, SBP (various issues)
Short-term spread (6-month MTB - deposit rate)	Statistical Bulletins, SBP (various issues)
Long-term spread (10-year bond - deposit rate)	Statistical Bulletins, SBP (various issues)

Note: SBP = State Bank of Pakistan, MOF = Ministry of Finance.

Table C2: Unit root test

	Augmented Dickey-Fuller test			Dickey-Fuller GLS test		
Variables	**Level**	**Difference**	**Order of integration**	**Level**	**Difference**	**Order of integration**
Nominal M2	-0.423	-2.004**	I(1)	-0.423	-2.004**	I(1)
Real M2	-1.810	-8.811*	I(1)	-0.848	-8.702*	I(1)
Real GDP (annual data)	-0.724	-6.366*	I(1)	0.749	-6.156*	I(1)
IPI	-1.934	-13.432*	I(1)	-1.012	-7.278*	I(1)
CPI	1.567	-3.6315*	I(1)	0.977	-3.284*	I(1)
Inflation	-2.556	-5.545*	I(1)	-0.978	-3.380*	I(1)
log (exchange rate)	-0.300	-3.435**	I(1)	0.161	-3.393*	I(1)
ER depreciation	-3.435**	-	I(0)	-3.416*	-	I(0)
Lending rate	-1.483	-3.369*	I(1)	-1.452	-2.977*	I(1)
Own rate (wt. avg. deposit rate)	-1.838	-2.228	I(1)	-0.527	-2.047**	I(1)
Short-term rate (wt. avg. 6-month MTB)	-1.274	-3.916*	I(1)	-1.341	-2.605**	I(1)
Long-term rate (wt. avg. 10-year FIB/PIB)	-1.942	-3.490**	I(1)	-1.481	-3.532*	I(1)
Risk premium (lending rate - MTB)	-1.913	-5.380*	I(1)	-0.801	-3.592*	I(1)
Short-term spread (6-month MTB - deposit rate)	-2.49156	-6.779*	I(1)	-2.270**	-	I(0)
Long-term spread (10-year bond - deposit rate)	-2.356	-5.715*	I(1)	-2.011**	-	I(0)

Note: Level of significance = * at 1%, ** at 5%, and *** at 10%. Critical values are from Mackinnon (1996). SBC is used for lag selection. Test estimation includes intercept. All variables are in log form except variant of interest rate, exchange rate depreciation, and inflation.

Appendix D

Engle-Granger Modeling Results

Table D1: Long-run static estimation of money demand (based on annual data)

Dependent variable: Log of broad money (sample range: 1978 to 2011)

	Nominal money		Real money	
Regressors	**(1)**	**(2)**	**(3)**	**(4)**
Log(GDP real)	1.24	1.13	1.26	1.14
	(7.54)	(7.32)	(59.35)	(24.2)
log(GDP deflator)	0.98	1.00		
	(9.61)	(10.31)		
Lending rate	-0.02	-0.01	-0.01	-0.01
	(-2.79)	(-2.08)	(-2.18)	(-1.81)
Inflation (averaged)			-0.01	-0.002
			(-2.65)	(-1.0)
Structural changes		0.03		0.03
		(0.22)		(1.58)
Constant	-8.96	-7.48	-4.55	-2.92
	(-4.31)	(-3.87)	(-14.9)	(-4.32)
Adj. R^2	0.99	0.99	0.99	0.99
DW	1.25	1.07	0.81	0.90
ADF of residual	-4.22*	-3.3**	-2.74***	-2.97**
Serial corr./hetro.	yes	Yes	Yes	yes

Note: t-statistics in parenthesis. *, **, *** = significant at 1%, 5%, and 10%, respectively. First principal component represents structural change.

Table D2: Error correction models of broad money (based on annual data, sample: 1973–2011)

	D(Log(M2)		D(Log(real M2)	
Dependent variable/regression	**(1)**	**(2)**	**(3)**	**(4)**
$ECM_{(t-1)}$	-0.63	-0.75	-0.50	-0.61
	(-4.49)	(-3.65)	(-4.10)	(-4.18)
D(LOG(M2(-1)))	0.64	0.70		
	(3.91)	(3.64)		
D(LOG(M2(-2)))	0.19	0.28		
	(1.29)	(1.42)		
D(LOG(real M2(-1)))			060	064
			(4.80)	(3.91)
D(LOG(real M2(-2)))			0.18	
			(1.61)	
D(LOG(YR))	0.75	0.54	0.56	1.20
	(2.77)	(1.41)	(2.15)	(3.25)
D(INF)			-0.01	-0.01
			(-5.30)	(-4.36)
D(LR(-1))		-0.01		
		(-1.52)		
D(LR(-2))	0.01	0.01	0.01	0.01
	(1.49)	(1.84)	(1.19)	(1.71)
D(structural change)		0.05		
		(1.73)		
D(structural change(-2)				0.01
				(0.63)
Constant	-0.01	-0.03	-0.01	-0.04
	(-0.40)	(-0.77)	(-0.78)	(-1.92)
Adj. R^2	0.43	0.39	0.65	0.67
SE	0.03	0.03	0.03	0.03
F-statistic	5.61	3.53	11.41	10.43
	[0.001]	[0.01]	[0.00]	[0.00]
DW	1.87	1.95	1.97	

Note: t-statistics in parentheses. Structural change is the first principal component.

Table D3: Long-run static estimation of money demand (based on quarterly data)

Dependent variable: Log of M2 (sample range: 1992Q1–2011Q4)

	Nominal money		Real money	
Regressors	**(1)**	**(2)**	**(3)**	**(4)**
Log(IPI)	0.35 (6.61)	0.35 (6.13)	0.98 (22.48)	0.53 (9.53)
Log(CPI)	1.53 (37.44)	1.53 (16.89)		
Lending rate	-0.02 (-7.63)	-0.02 (-7.19)	-0.01 (-0.89)	-0.02 (-4.66)
Inflation (Q/Q)			-0.01 (-1.78)	-0.004 (-1.75)
Log(ER)		0.01 (0.94)		0.47 (9.53)
Constant	5.84 (41.43)	5.84 (38.08)	9.64 (40.06)	10.1 (59.84)
Adj. R^2	0.99	0.99	0.89	0.95
DW	0.74	0.75	1.45	1.10
ADF of residual	-4.22*	-4.28*	-2.74*	-5.5*
Standard error	0.07	0.07	0.12	0.08
Serial corr./hetro.	yes	yes	yes	yes

Note: t-statistics in parentheses. *, ** , and *** = significant at 1%, 5%, and 10%.

Appendix E

ARDL Modeling Results

Table E1: Long-run relationship of real money (ARDL approach)

Dependent variable: log real M2 (sample range: 1992Q1–2011Q4)

Variables		M-1	M-2	M-3	M-4	M-5	M-6	M-7
Log(IPI)		0.86	0.69	0.92	0.88	0.87	0.83	0.93
		(7.91)	(6.74)	(7.70)	(12.59)	(6.77)	(6.41)	(13.04)
Lending rate		-0.02	-0.03		-0.02			
		(-1.51)	(-3.89)		(-1.89)			
Inflation		-0.02	-0.003	-0.02		-.0.02	-0.01	
		(-1.85)	(-0.43)	(-1.37)		(-1.38)	(-1.23)	
Expected inflation					-0.02			-0.02
					(-2.73)			(-1.82)
Log(exchange rate)			0.22					
			(1.98)					
Expected depreciation of ER					0.03			0.04
					(2.18)			(2.40)
Own rate (wt. avg. deposit rate)						-0.04	0.01	
						(-0.76)	(0.17)	
Short-term rate (wt. avg. 6-month MTB)				-0.02		-0.01	0.003	
				(-1.37)		(-0.38)	(0.12)	
Long-term rate (wt. avg. 10-year FIB/PIB)							-0.03	
							(-1.05)	
Short-term spread (6-m MTB - deposit rate)								
Long-term spread (10-y bond - deposit rate)								
Risk premium (lending rate – 6-m MTB)								0.01
								(0.53)
Constant		11.75	11.48	11.36	11.62	11.61	11.94	11.18
		(20.45)	(28.30)	(20.82)	(30.70)	(19.07)	(18.16)	(35.59)
F-statistics [probability]		4.552	4.150	3.369	5.9502	3.805	3.452	4.526
		[0.001]	[0.003]	[0.014]	[0.000]	[0.004]	[0.005]	[0.002]
Bounds test	Upper value	4.378	4.049	4.378	4.049	4.049	3.805	4.049
	Lower value	3.793	2.85	3.793	2.85	2.85	2.649	2.85
Lag length selection	SBC and AIC	1 lag & 6 lags	1 lag & 5 lags	1 lag & 5 lags	1 lag & 5 lags	1 lag & 5 lags	1 lag & 7 lags	1 lag & 5 lags
ARDL lag selection		(1,0,0,0)	(3,0,0,1,1)	(2,0,1,0)	(3,0,0,0,1)	(1,0,0,0,0)	(1,0,0,0,0,0)	(1,0,0,0)
Observations		79	77	79	77	72	72	77

Note: SBC is used to select the optimum number of lags in the ARDL model.

Table E2: Short-run error correction model
Dependent variable: Δlog real M2 (sample range: 1992Q1–2011Q4)

Variables	M-1	M-2	M-3	M-4	M-5	M-6	M-7
ECM $_{(-1)}$	-0.10	-0.17	-0.09	-0.12	-0.10	-0.11	-0.11
	(-3.49)	(-4.55)	(-2.65)	(-4.92)	(-2.71)	(-2.89)	(-4.53)
DLIPIA	0.09	0.11	0.09	0.11	0.09	0.09	0.11
	(2.91)	(4.77)	(2.58)	(4.36)	(2.51)	(2.59)	(4.14)
DLR	-0.002	-0.004		-0.002			
	(-1.61)	(-3.09)		(-1.84)			
DINF	-0.002	-0.01	-0.002		-0.002	-0.002	
	(-1.89)	(-3.23)	(-1.57)		(-1.57)	(-1.37)	
D Exp(INF)				-0.002			-0.002
				(-2.80)			(-1.90)
DLER		-0.29					
		(2.70)					
D EXP(ER)				0.001			0.001
				(0.46)			(0.87)
D (own rate)					-0.003	0.001	
					(-0.77)	(0.17)	
D short-term			-0.001		-0.000	0.0002	
(6-m MTBs)			(-1.21)		(-0.004)	(0.12)	
D long-term (10-y bond rate)						-0.004	
						(-1.02)	
DLRM21		-0.45		-0.49			-0.42
		(-4.41)		(-4.77)			(-4.24)
DLRM22		0.29		0.24			0.32
		(3.01)		(2.48)			(3.47)
D(risk premium)							0.001
							(0.53)
Constant	1.21	1.91	1.07	1.43	1.14	1.29	1.27
	(3.73)	(4.81)	(2.68)	(5.12)	(2.78)	(2.96)	(4.70)
Adj. R^2	0.15	0.56	0.11	0.53	0.10	0.1	0.51
F-statistics	F(4,74)	F(5,69)	F(4,68)	F(7,69)	F(5,66)	F(6,65)	F(7,69)
	7.35[0.003]	15.4[0.00]	3.18[0.02]	13.46[0.00]	2.60[0.033]	2.34[0.014]	12.45[0.00]

Table E3: Forecasting of broad money

	M1	M-2	M-3	M-4	M-5	M-6	M-7
Q1-FY11	2.58%	0.77%	3.23%	0.49%	3.00%	3.08%	1.12%
Q2-FY11	1.88%	6.08%	2.76%	6.74%	2.63%	2.16%	7.1%
Q3-FY11	2.95%	2.54%	2.18%	0.80%	2.05%	1.71%	1.4%
Q4-FY11	2.21%	3.48%	2.23%	4.19%	2.15%	1.82%	4.77%
FY11	10.0%	13.43%	10.82%	12.66%	10.20%	9.07%	15.11%

Note: For forecasting purposes, we use the actual data on CPI averaged inflation (13.7), averaged lending rate (13.9%), IPI growth (0.4%), expected depreciation (-0.8%), and M2 growth (15.89) for FY2011.

Appendix F

Johansen Modeling Results

Table F1: VAR lag order selection criteria

Endogenous variables: LOG(M2) LOG(IPIA) LOG(CPIA) LR
Exogenous variables: Constant, sample 1991Q1 to 2011Q4, observations = 74

Lag	LogL	LR	FPE	AIC	SC	HQ
0	-67.973	NA	0.000	2.027	2.155	2.078
1	352.744	782.178	0.000	-9.373	-8.736	-9.120
2	440.986	154.112	0.000	-11.408	-10.261	-10.952
3	475.577	56.515	0.000	-11.932	-10.275*	-11.273
4	508.812	50.556*	4.94e-11*	-12.417*	-10.250	-11.556*
5	520.439	16.375	0.000	-12.294	-9.617	-11.230
6	527.591	9.268	0.000	-12.045	-8.858	-10.778
7	548.447	24.675	0.000	-12.182	-8.485	-10.712

Note: * indicates lag order selected by the criterion.
LR = sequential modified LR test statistic (each test at 5% level), HQ = Hannan-Quinn information criteria. FPE = final prediction error, AIC = Akaike information criterion, SC = Schwarz information criterion.

Table F2: VAR lag order selection criteria

Endogenous variables: LOG(RM2) LOG(IPIA) INF(1) LR
Exogenous variables: Constant, sample 1992Q1 to 2011Q4, observations = 74

Lag	LogL	LR	FPE	AIC	SC	HQ
0	-336.383	NA	0.171486	9.588	9.716	9.639
1	7.800	639.890	1.66E-05	0.344	0.981	0.597
2	81.685	129.038	3.26E-06	-1.287	-0.140	-0.831
3	111.799	49.201	2.22E-06	-1.684	-0.027	-1.025
4	154.695	65.250*	1.06e-06*	-2.442	-0.275*	-1.580*
5	171.856	24.171	1.06E-06	-2.475*	0.202	-1.410
6	181.227	12.142	1.35E-06	-2.288	0.899	-1.021
7	194.098	15.228	1.59E-06	-2.200	1.497	-0.730

Note: * indicates lag order selected by the criterion.
LR = sequential modified LR test statistic (each test at 5% level), HQ = Hannan-Quinn information criteria. FPE = final prediction error, AIC = Akaike information criterion, SC = Schwarz information criterion.

Table F3: Money demand function with different specifications
Dependent variable: Log(real M2), Sample: 1992Q1 to 2011Q2

Model	Estimated long-run relation	SBC/AIC lag-length	Cointegration: Trend specifications	Trace	Max. eigenvalue
1	log(M2) = 1.21*Log(IPI) + 1.88*Log(CPI) - 0.03*Lending rate -0.02*trend + 5.55	2 & 4	Unrestricted intercept with linear deterministic trend	1	1
2	log(real M2) = 0.88*log(IPI) – 0.08*Exp(inflation) – 0.16*Lending rate + 13.75	4 & 5	Restricted intercept without trend	**1**	**1**
3	log(RM2)=1.13*LOG(IPI) + 0.05*Exp(Inflation) + 0.06*MTB + 9.05	4 & 4	Restricted intercept without trend	1	0
4	log(RM2)=-1.53*LOG(IPI) +2.52*ER + 0.07*Exp(Inflation) - 0.02*MTB - 0.05*trend -0.07	2 & 7	Unrestricted intercept with linear deterministic trend	2	1
5	log(RM2)=1.25*LOG(IPI) + 0.02*Exp(Inflation) + 1.16*log(ER) - 0.02*LR - 0.02*trend - 5.99	2 & 5	Unrestricted intercept with linear deterministic trend	1	1
6	log(RM2)=0.48*LOG(IPI) + 0.002*Exp(Inflation) - 0.01*Risk +0.01*trend+ 12.4	3 & 5	Unrestricted intercept with linear deterministic trend	1	2
7	log(RM2)=1.38*LOG(IPI) - 0.08*Exp(Inflation) - 0.34*log(ER)- 0.11*Risk +11.23	2 & 7	Restricted intercept without linear deterministic trend	1	2

Note: Weighted average rate of 6-month MTB, lending rate (LR), averaged exchange rate (ER), risk = LR - MTB.

Table F4: Estimation of money demand (Johansen procedure)

Sample end period	η_{ipi}	$\eta_{E(inf)}$	η_{LR}	Cointegration Trace	Cointegration Max. eigenvalue
FY05-Q4	-3.10	-0.02	-0.21	1	1
	(-1.58)	(-0.76)	(-4.08)		
FY06-Q4	-2.06	-0.04	-0.31	3	1
	(-0.92)	(-1.18)	(-3.97)		
FY07-Q4	-0.13	-0.11	-0.48	4	1
	(-0.05)	(-1.57)	(-4.94)		
FY08-Q4	-0.19	-0.002	-0.11	1	1
	(-0.32)	(-0.09)	(-4.98)		
FY09-Q4	0.45	0.02	-0.04	1	1
	(1.68)	(-3.28)	(4.18)		
FY10-Q4	-0.41	0.08	0.20	1	1
	(-0.38)	(3.33)	(4.35)		
FY11-Q4	0.88	-0.08	-0.16	2	1
	(4.27)	(-3.69)	(-4.45)		

Table F5: Estimation of money demand (Johansen procedure)

Sample end period	η_{ipi}	$\eta_{E(inf)}$	η_{ER}	η_{risk}	Cointegration Trace	Cointegration Max. eigenvalue
FY05-Q4	0.78	-0.01	0.31	-0.04	1	1
	(6.17)	(-2.17)	(3.14)	(-5.54)		
FY06-Q4	0.66	-0.01	0.36	-0.05	1	1
	(6.90)	(-2.01)	(3.74)	(-5.75)		
FY07-Q4	0.68	-0.01	0.34	-0.04	1	1
	(9.03)	(-2.26)	(3.91)	(-6.05)		
FY08-Q4	0.65	-0.01	0.36	-0.04	1	1
	(9.59)	(-1.98)	(4.51)	(-5.68)		
FY09-Q4	0.65	-	0.47	-0.03	2	2
	(12.97)		(8.38)	(-5.20)		
FY10-Q4	0.84	-0.02	0.26	-0.05	1	1
	(7.18)	(-3.84)	(2.20)	(-4.68)		
FY11-Q4	1.35	-0.08	0.30	-0.09	1	2
	(6.56)	(-6.64)	(1.41)	(-4.35)		

Appendix G

Bayesian Modeling Results

Table G1: Benchmark prior estimates

Parameters	Description	Benchmark priors	Source
α	Share of capital in production function	0. 46	Haider and Khan (2008)
β	Subjective discount factor	0.99	Ahmed, Haider, and Iqbal (2012)
ρ	Real interest rate in steady state	0.025	Authors' calculations
θ	Measure of price stickiness	0.75	Haider and Khan (2008)
κ	Slope coefficient in NKPC	$\frac{(1-\theta)(1-\beta\theta)}{\theta}$	Haider and Khan (2008)
γ	Parameter of risk aversion	0.587	Ahmed et al. (2012)
φ_1	Output elasticity of money demand	0.860	Authors' calculations
φ_2	Interest elasticity of money demand	-0.018	Authors' calculations
χ_1	Sensitivity of central bank with respect to inflation	1.2	Authors' calculations
χ_2	Sensitivity of central bank with respect to output	0.31	Authors' calculations
ρ_a	Persistence of technology shock	0.97	Authors' calculations
ρ_m	Persistence of money demand shock	0.47	Authors' calculations
ρ_i	Persistence of monetary policy shock	0.32	Authors' calculations

Table G2: Model prior and posterior distribution results

	Prior distributions			Posterior distribution			
Parameters	Distribution	Mean	SD	Distribution	Mean	5th percentile	95th percentile
φ_1	gamma	0.860	0.045	gamma	0.859	0.798	0.917
φ_2	gamma	-0.018	0.005	gamma	-0.024	-0.026	-0.009

Note: The posterior means of all the estimation parameters are delivered by 150,000 runs of the Metropolis-Hastings algorithm. We have used MATLAB toolbox Dynare 4.1 for this simulation purpose.

Figure G1: Bayesian prior vs. posterior distribution plots

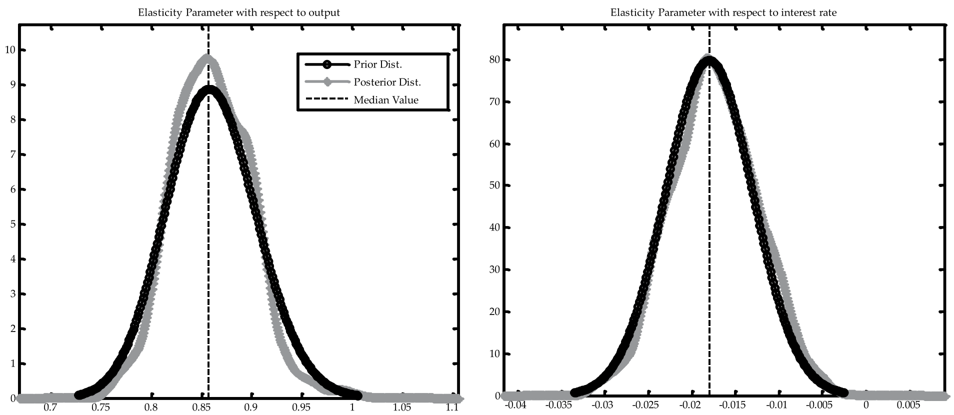
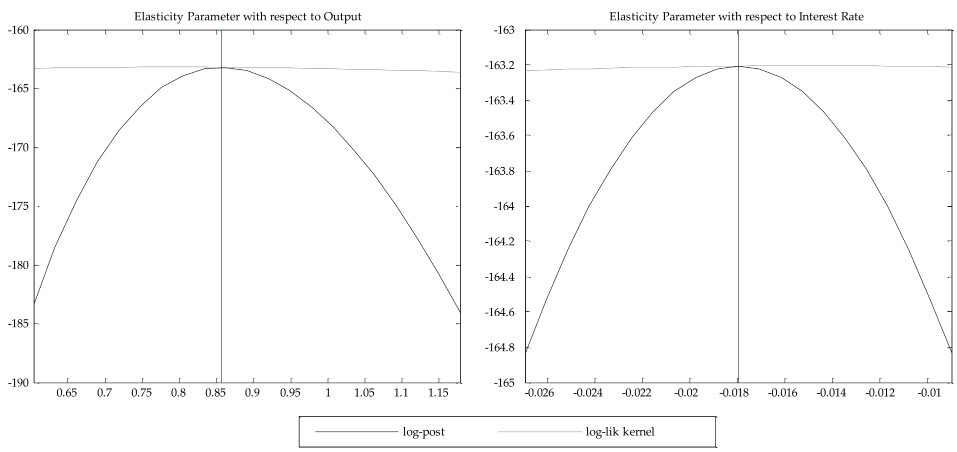


Figure G2: Distribution plots of Bayesian posterior kernel and log-likelihood



Appendix H

Stability Results

Figure H1: Trends in economic activity and first-principal component

Figure H1a: Investment as % of GDP

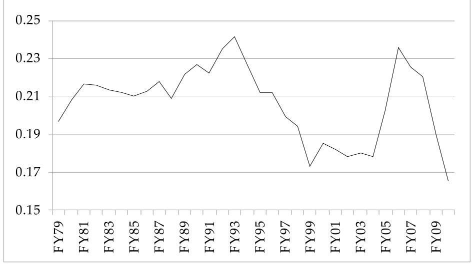


Figure H1b: Imports as % of GDP

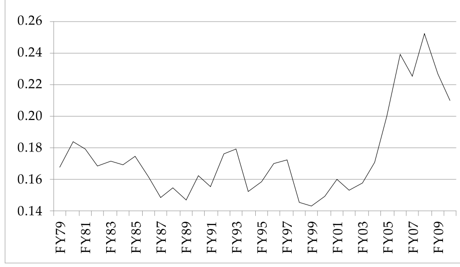


Figure H1c: Credit to Private sector as % of M2

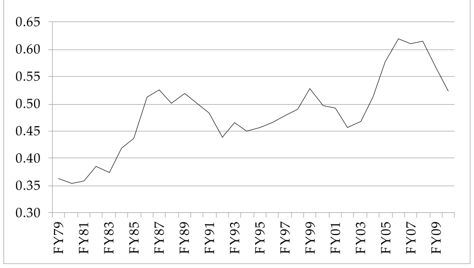


Figure H1d: Manuf. & Services Sector as % of GDP

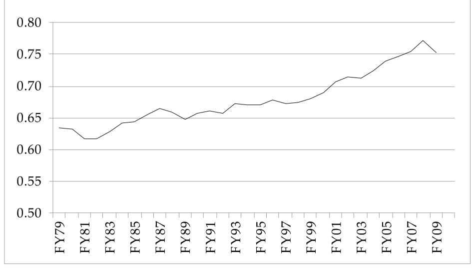


Figure H1e: Government Consumptions as % of GDP

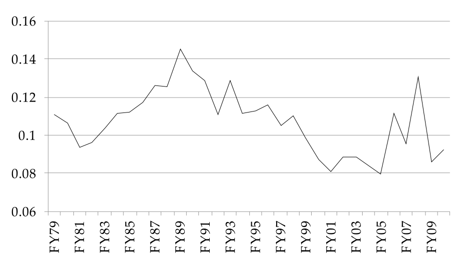


Figure H1f: First Principle Component of five series

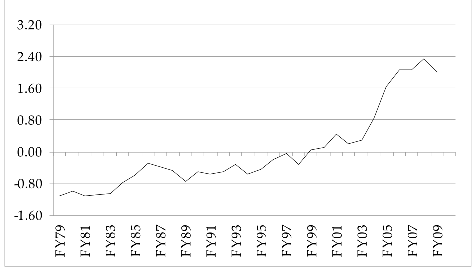
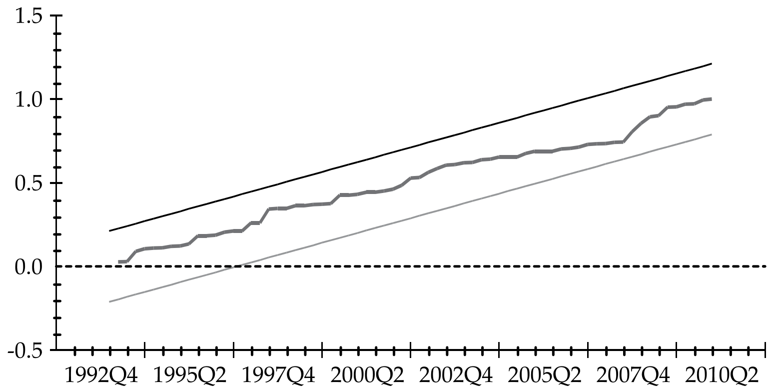
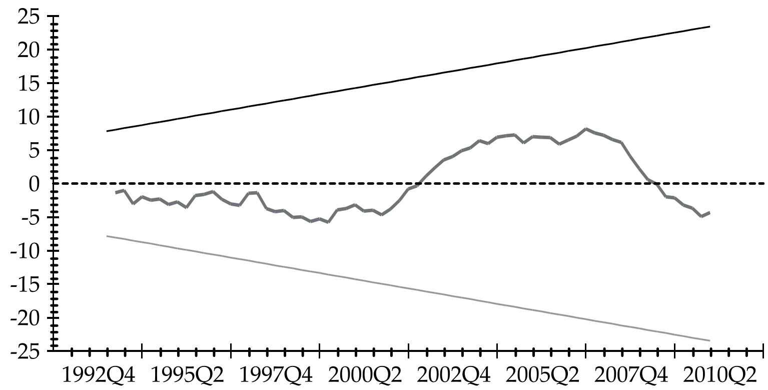


Figure H2: CUSUM and CUMUS square of ARDL Model 4



The straight lines represent critical bounds at 5% significance level



The straight lines represent critical bounds at 5% significance level

Figure H3: Cointegration relation of Models 2 and 7 (Johansen procedure)

Figure H3a: Cointegration Relation of Model-2

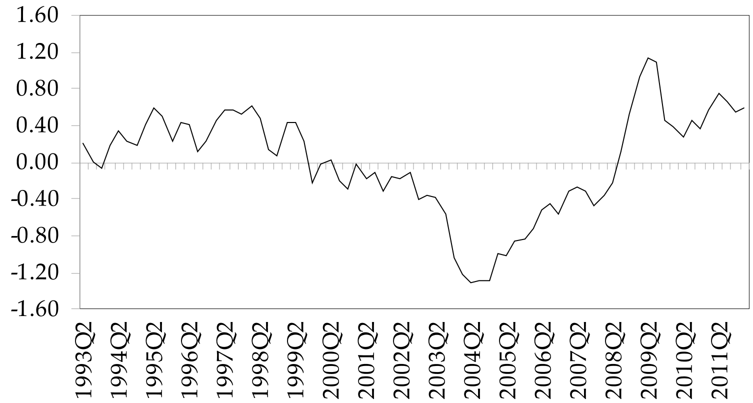


Figure H3b: Cointegration Relation of Model-7

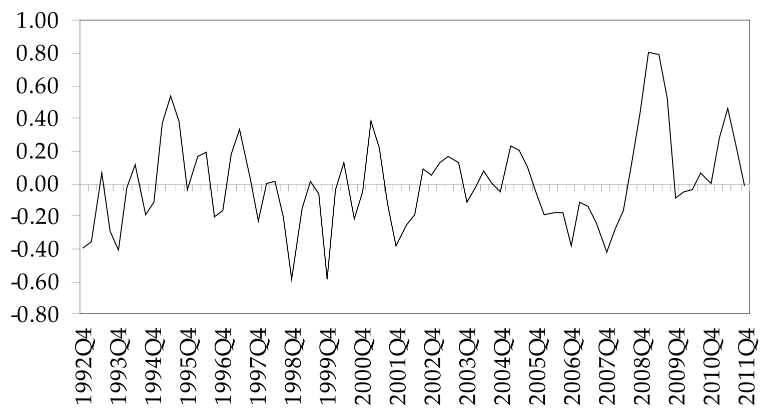


Figure H4: Recursive stability estimates of Model 7 (Johansen procedure)

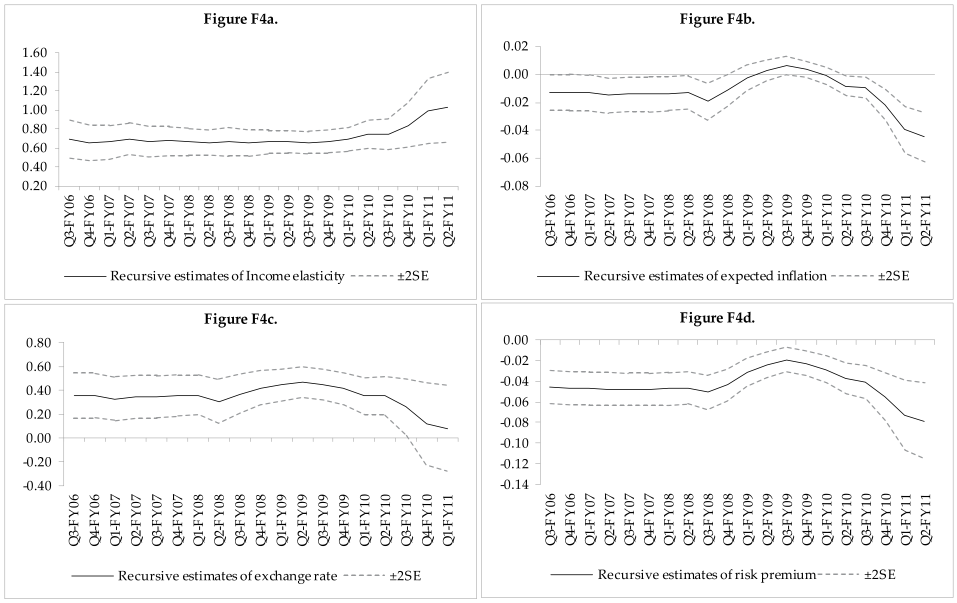


Figure H5: Bayesian reduced-form screening of parameters

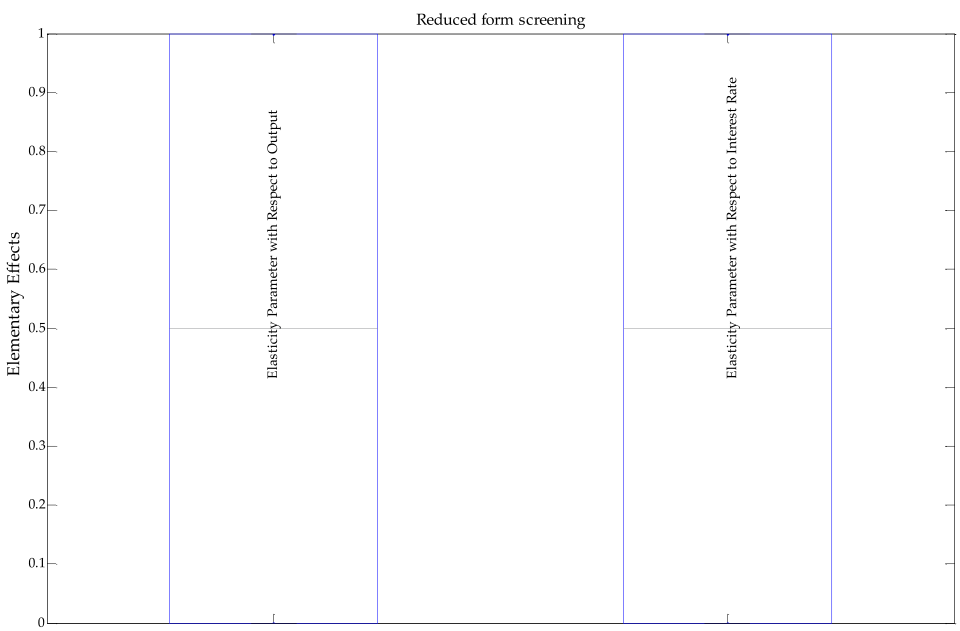
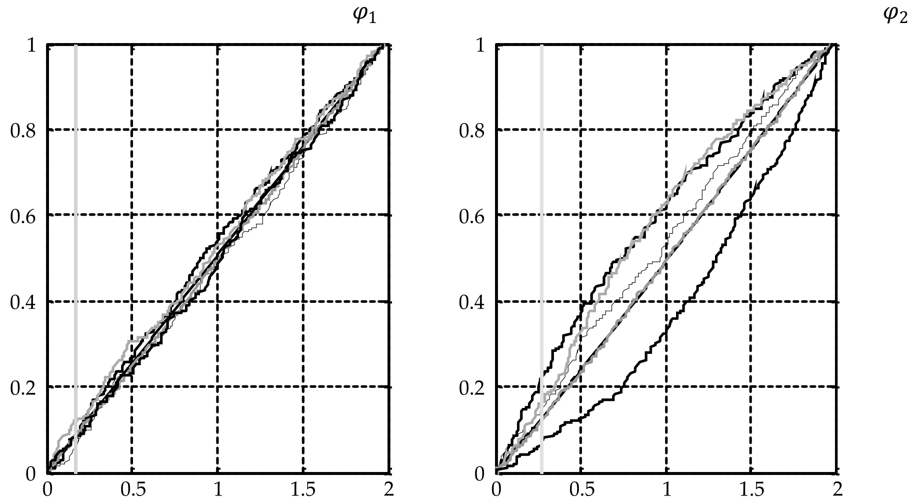


Figure H6: Bayesian posterior stability plots using GSA toolkit



Note: Since we have used different specifications of the nominal interest rate, we find different stability paths for the parameters.

Appendix I

Selected Empirical Literature Review

No.	Study	Coverage of variables	Functional form	Findings
1	Akhtar (1974) Sample space: 1951–70	M1 M2 Interest rates Real GDP Inflation	Log linear	Interest rates and income are found to be the most important determinants of money demand function. Long-term interest rates are significant and robust with unit elasticity of income.
2	Abe, Fry, Min, Vongvipanond, and Yu (1975) Sample space: 1951–70	M1 per capita M2 per capita GDP per capita Interest rates Inflation	Log linear	Various models for different definitions of interest rates and national income. Main determinants of M1 and M2 are interest rate, inflation, and income. Interest rate and inflation rate are statistically significant with negative sign. Income is statistically significant with positive sign.
3	Khan (1980) Sample space: 1960–78	M1 M2 Income (measured) Income (permanent) Interest rates Bank branches	Log linear	Income, rate of interest on time deposits, inflation, and degree of monetization are the important explanatory variables explaining almost 99 percent of the variation in money demand. No evidence for the hypothesis that permanent income is a better explanatory variable than measured income.
4	Khan (1982) Sample space: 1960–78	M1 M2 Income (measured) Income (permanent) Interest rates	Log linear	Study conducted for six developing countries including Pakistan. Importantly, it does not alter the main findings of Khan (1980). However, it finds that permanent income and expected inflation are better explanatory variables in the case of Sri Lanka compared to measured income and actual inflation, respectively.
5	Nisar and Aslam (1983) Sample space: 1960–79	M1 M2 GNP Interest rates	Log linear	Various models estimated with alternative measures of interest rates and money stocks. Main conclusion is that interest rates are significant and robust with unit money elasticity.
6	Ahmad and Khan (1990) Sample space: 1960–87	M1 M2 Real GDP Interest rate	Log linear and time-varying parametric approach	Money demand function remains robust up to 1980 and unstable thereafter.

No.	Study	Coverage of variables	Functional form	Findings
7	Hossain (1994) Sample space: 1951–91	M1 M2 Real GDP Yield on govt. bonds Market call rate Inflation	Log linear	Stable money demand function determined through Johansen's cointegration tests. Interest rate has a significant negative impact on money stock with unit elasticity of income.
8	Qayyum (1998)	Real money demand Real income Yield on long-term govt. bonds Measured inflation Seasonal dummies	Error correction model	Long-run money-income proportionality hypothesis is accepted.
9	Qayyum (2005)	M2 Nominal income Interest rates Inflation	Johansen's cointegration and dynamic error correction	Inflation is an important determinant of money demand. Rates of interest, market rate, and bond yield are important for long-run money demand behavior.
10	Moinuddin (2009)	Real M2 Real GDP Real interest rates	Log linear	Money demand function is unstable in Pakistan and therefore monetary aggregate targeting is not suitable.
11	Omer and Saqib (2009)	M2 Real GDP Inflation	ARDL	The quantity theory is an inadequate explanation of inflation, income velocity of money is unstable, and money is endogenous. These results suggest the need to rethink monetary targeting strategy in Pakistan.
12	Narayan, Narayan, and Mishra (2009)	M2 Yd ER Interest rates (foreign and domestic)	Panel cointegration and panel long-run estimation	Panel Granger causality test suggests short-run causality running from all variables, except foreign interest rate, to money demand. Money demand functions are all stable except for Nepal.
13	Azim et al. (2010) Sample space: 1973–2007	Broad money (M2) Real GDP Inflation Exchange rate	ARDL	Long-run relationship between broad money and goal variables. Money demand function is stable in the case of Pakistan, using CUSUM and CUSUM square test.
14	Omer (2010) Sample space: 1975–2006	Reserve money (M0) Narrow money (M1) Broad money (M2) Velocity of money (M0, M1, M2) Call money rate Nominal GDP Per capita income CPI inflation	ARDL	Velocity of base and broad money is insensitive to interest rate changes, but responsive to income and business cycle fluctuations. However, velocity of narrow money (M1) depends on interest rate changes, income, and business cycle fluctuation. Money velocities of all three models are stable using CUSUM and CUSUM square test. Money demand is stable.

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Forward-Looking and Backward-Looking Taylor Rules: Evidence from Pakistan

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Abstract

This study uses the forward-looking rule and backward-looking Taylor rule to investigate the conduct of monetary policy in Pakistan during 1971–2011. We compare the pre- and post-reform periods, and find that the estimates obtained using the generalized method of moments indicate that no interest rate rule was being followed. This explains the inability of monetary policy to control inflation and minimize the output gap. Although monetary policy was not very active in the pre- and post-reform periods, the post-reform quarterly data show some interest rate inertia and smoothing. Monetary policy was less accommodating of the cyclical nature of the output gap. We conclude that the behavior of the State Bank of Pakistan was not very different under forward- or backward-looking rules.

Keywords: Taylor rule, forward-looking behavior, backward-looking policy, monetary policy, generalized method of moments.

JEL classification: C22, E52, E58.

1. Introduction

Pakistan's inflation rate fluctuated widely between 3 and 27 percent during 1971:Q1–2011:Q4. The average inflation rate was 9.5 percent with a standard deviation of 6 percent. With this highly volatile rate, every episode of high inflation was followed by a tamed inflationary regime. However, the inflation rate and growth rate followed a mixed trend: the high inflationary environment of the 1980s was accompanied by high growth, but after 1990, the regime yielded contrary results. The State Bank of Pakistan (SBP) introduced several reforms in the financial sector during this time, which influenced monetary policy. As a result, the SBP made some adjustments to the interest rate but found itself still facing fiscal dominance. Although these policy changes brought about some success in restraining inflation, the instability continued.

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In this paper, we ask how monetary policy is conducted in such an environment. Does it rely on discretion or the use of a sophisticated rule? In this context, discretion can be seen as "muddling through." Lagged reactions may exacerbate an inflationary situation whereas rules are determined and set such that the central bank follows a particular rule regardless of the situation. However, this rigidity limits policy options and any deviation from the set path can result in short-run fluctuations.

Taylor's (1993b) rule relates movements in the interest rate to the inflation rate and output gap. The rule is simply an equation indicating that the larger the coefficients, the more aggressive monetary policy will be. Taylor's rule approximates the path of the quarterly federal funds rate during 1987:Q1–1992:Q3. The debate that followed this study theorized a rule-based monetary policy as the only mandate of a central bank. The effectiveness of the central bank was judged on the basis of policy reaction functions in controlling inflation. Taylor (2000) clarifies that a monetary policy rule is nothing more than a contingency plan that describes as precisely as possible the circumstances in which a central bank can change the instruments of monetary policy. Taylor's earlier (1999a) study advocates a historical rather than model-based approach to monetary policy, arguing that historical rules evolve slowly but allow the separation of policy influences. Subsequently, Rudebusch (2002) has held that monetary policy sluggishness or gradualism is an illusion.

Many studies have raised questions about adopting rules. Goodhart (1984, 1991) notes that rules are rigid: they fail to address the complexities of changing circumstances and tend to collapse in extraordinary circumstances. The targets of a government with damaged credibility become irrelevant, leaving the field open for economic indicators to guide policy. An important question relates to the determination of the interest rate. Is the Taylor rule based on past movements in inflation or on the expected inflation rate? Kerr and King (1996), Bernanke and Woodford (1997), and Clarida, Galí, and Gertler (1997, 1998, 2000) suggest that the future inflation rate acts as a policy guide for central banks to help them avoid inconsistencies. When households and businesses start to expect higher inflation in the economy, this generates more inflation. In this case, the past inflation rate is not a good guide for the monetary authority.

Taylor (1993a, 1993b) points out that numerous generalizations can make these rules more responsive but also more complex. He gives the example of estimating the expected inflation rate where one would need to use futures markets, the term structure of interest rates, and various surveys. Problems inherent in measuring the output gap—such as predictions about productivity, labor force participation and changes in the natural rate of unemployment—mean that the rule must be kept simple.

Mishkin (1996) argues that it is dangerous to invariably associate monetary policy easing or tightening with a rise or fall in short-term nominal interest rates because the nominal interest rate does not always pass through to the real interest rate. Other asset prices, e.g., stock prices and land prices, are also based on information about the central bank's monetary policy stance. Open market operations are also an important tool for monetary policy that works by increasing inflation and smoothing other asset prices, thus increasing output.

Price stability is crucial for business decisions because inflation can affect economic growth adversely. Many central banks pursue a strategy of proactively raising interest rates to prevent a rise in inflation in an overheated economy. The policy's success requires the monetary authority to make an accurate assessment of the timing and effect of its policies on the economy. Woodford (2000) explains the problems of forecasting private sector expectations when determining the future rate of inflation. In forward-looking models, an optimal policy depends on the path of a target that evolves over time. These principles are applicable only in a controlled environment; this does not exist in the case of monetary policy, which requires the private sector to be able to respond in the presence of shocks.

This paper aims to address the following questions. How does the SBP conduct monetary policy to achieve price stability and minimize the output gap? Do interest rate changes reasonably approximate the inflation rate in Pakistan? If so, do they follow the Taylor rule based on past or future inflation?

Section 2 provides an overview of monetary policy in Pakistan. Section 3 presents an analytical framework and specifies the Taylor rule for Pakistan; it also describes the data used and their sources. Section 4 carries out estimations using the annual and quarterly time series, subdividing the latter into five monetary regimes to examine the relative autonomy of the SBP. We apply the Chow test to ascertain the presence of any structural breaks across regimes and differences in policy response. Section 5 traces the rules path and estimates the social loss function. Section 6 presents some concluding observations.

2. Monetary Policy in Pakistan: An Overview

The SBP became relatively autonomous in the early 1990s with the onset of structural adjustment and liberalization reforms. During 1989–92, Pakistan implemented the World Bank's Financial Sector Deepening and Intermediation Project. To support this reform program, the IMF initiated a three-year structural adjustment program (see Janjua, 2004). This section compares the conduct of monetary policy in the pre-reform and post-reform periods over 1971–2011.

We distinguish between different monetary regimes on the basis of changes in leadership (governor) at the SBP. Subdividing the quarterly data series into pre- and post-reform regimes enables us to look closely at the stability and reaction function of monetary policy in periods of relative autonomy. The post-reform period is also subdivided into five regimes on the basis of governors' tenures, though not necessarily coinciding with changes in the political regime. It is interesting to note that most of these governors were appointed by caretaker regimes (see Appendix). The rapid turnover of political regimes makes it difficult to apply any sophisticated statistical analysis in order to understand the issue of the SBP's autonomy.

The duration of the SBP governor's tenure has not always been constant. For the sake of statistical analysis, therefore, we adjust some overlapping tenures. I. A. Hanfi assumed the governor's office on 17 August 1988. Following his resignation, Kasim Parekh was appointed governor; his term ended on 30 August 1990, after which Hanfi was reappointed from 1 September 1990 to 30 June 1993. Given the small number of observations for this period, we merge Parekh's tenure with Hanfi's second tenure. Thus, the period 1989:Q1–1993:Q3 is identified as Hanfi's tenure. The second regime under Mohammad Yaqub lasted two full terms from July 1993 to November 1999. Ishrat Husain, who followed, also completed two full terms from 2 December 1999 to 1 December 2005. Shamshad Akhtar, the SBP's first woman governor, succeeded Husain from 2006:Q1 to 2008:Q3, while the present regime has already seen three governors (Salim Raza, Shahid Kardar, and the present governor, Yasin Anwar).

Under the State Bank of Pakistan Act, the scope for independent action increased incrementally, although the SBP has been perceived as acting more or less autonomously depending on the governor's strength of personality. The first noteworthy attempt to gain some autonomy for the SBP was made by S. U. Durrani at the 23rd general board meeting on 18 September 1971. The effort was short-lived as the SBP soon became virtually attached to the finance ministry. On 28 November 1989, Hanfi informed the board that the SBP had no effective control over monetary policy; he went on leave and eventually resigned in protest. Hanfi was not the only governor who had to resign from office. The period 1986–93 was highly destabilizing not only in political terms and changes in government, but also for the SBP.

The first formal step toward autonomy was taken in 1993 when the SBP was detached from the finance ministry (Janjua, 2004). Several significant changes took place during this period of relative autonomy (1989:Q1–2011:Q4). In 1991, permission to open new commercial banks was granted. Nonetheless, Hanfi's policies, similar to those of his predecessor, were less active in both directions: the SBP neither stabilized prices nor tried to reduce the output gap. The finance ministry continued to make decisions even on routine matters of the SBP.

In August 1993, the caretaker government of Prime Minister Moeen Qureshi recognized the need for the central bank's autonomy to improve macroeconomic management. It proposed the separation of fiscal and monetary management. This was a period of low external financial assistance and mounting debt servicing. While a tight monetary policy may not have been the ideal choice, it was the only option left.

Under Yaqub's governorship during this period, the SBP's autonomy was never fully absorbed. A new government followed the caretaker regime and the Monetary and Fiscal Policy Coordination Board was formed, allowing the finance ministry back in the driver's seat. Yaqub resigned three times from the governorship because of his commitment to financial liberalization. In 1995, maximum lending rates—except on concessionary finance schemes—were abolished. Minimum lending rates were abolished in 1997. These price ceilings and floors were the main reason for the prevailing market rigidities and distortions.

The Husain regime was no different from that of Hanfi in incorrectly estimating the state of the economy. Monetary policy remained discretionary. The SBP's autonomy was diluted once again in the name of better financial regulation. The President became responsible for appointing the bank's governor while the federal government appointed its deputy governors.

Under Shamshad Akhtar, the SBP prepared a ten-year strategy paper on banking sector reforms. The paper recommended making changes in the State Bank of Pakistan Act to redefine and strengthen the SBP's role in making and executing monetary policy. It also sought a clear-cut role for the central bank in advising the government on fiscal policy and domestic debt management (SBP, 2009). However, Akhtar's term ended before any concrete steps could be taken.

Governors are appointed initially for a period of three years, which may be extended for another three years. Since Akhtar's departure, the actual tenures have become shorter. Some reforms have been introduced, such as the separation of liquidity and debt management, a corridor framework for the overnight money market rate, and the institution of a representative monetary policy committee to improve transparency credibility. The frequency of monetary policy announcements has also increased. However, the changes required in the State Bank of Pakistan Act to ensure the central bank's autonomy have not taken place (see SBP, n.d.).

3. Analytical Framework

This section describes the framework within which we estimate a Taylor-type rule for Pakistan's economy.

3.1. The Taylor Rule

Taylor (1993b) explains monetary policy as an interest feedback rule, where the percent federal funds rate (i_t) is a function of the percent inflation rate (π_t) and the percent change in output gap (y_t)

$$i_t = 0.04 + 1.5\,(\pi_t - 0.02) + 0.5(y_t - y_t^*) \tag{1}$$

If the central bank follows this rule strictly, it must have a 2 percent inflation and interest rate rule. The federal funds rate rises when there is an increase in the inflation rate from 2 percent or when real GDP exceeds the trend. When the central bank achieves its inflation and real GDP targets, then the federal funds rate will be equal to 4 percent. The Taylor rule is considered a fairly good explanation of US monetary policy and a prescription for desirable policy rule or an indicator for assessing policy behavior (Woodford, 2001).

Buzeneca and Maino (2007) argue that developing countries apply a rule-based monetary policy more intensively because of their shallow markets. Taylor's (2000) rule-based policy is a better instrument for developing countries because of velocity shocks. Monetary aggregates are preferable only if measuring the real interest rate is difficult or if major shocks to investment occur. Orphanides and Wieland (2013) modify the Taylor rule and give it a more generalized form as follows:

$$i_t = i_{t-1} + 0.5\left(\pi_{(t+3|t)} - \pi^*\right) + 0.5\left(q_{(t+2|t)} - q^*_{(t+2|t)}\right) \tag{2}$$

where i_t stands for the federal funds rate set by the central bank, π denotes the inflation rate, π^* denotes the target inflation rate, q stands for the GDP growth rate, q^* denotes potential GDP, t is a time subscript representing one quarter, and t+2,3 indicates the second and third quarter forecasts, respectively. This equation shows that the central bank adjusts its policy rate based on deviations in the forecasted inflation rate from the target inflation rate and on deviations in forecasted GDP growth from the estimated potential GDP growth. The regression coefficient 0.5 implies that a one-percentage point deviation in the target inflation rate or output growth requires the policy rate to be adjusted by 50 points. This can be converted into a simple formula for estimation:

$$r = r* + C\pi(\pi - \pi*) + CY\, y \tag{3}$$

where r is the real interest rate, and $C\pi$ and CY are the coefficients on the policy rules.

The real interest rate is calculated as $r = i - \pi$, after substituting the value of r in the above equation and obtaining

$$i = (r^* - C_\pi\pi) + (1 + C_\pi)\pi + C_y y_t \tag{4}$$

where $(r^* - C_\pi\pi^*) = C$

$$i = C + (1 + C_\pi)\pi + C_y y_t \tag{5}$$

This rule assumes that the real interest rate is adjusted around the target rate, and the inflation rate and output gap deviate from the target, which is assumed to be π^* and 0. The equation then takes the form

$$r = C + C_\pi\pi + C_y y_t \tag{5a}$$

$$C > 0, (1 + C_\pi) \geq 1, C_y \geq 0$$

If monetary policy follows a Taylor rule-like prescription, the intercept value (C) has to be positive and the inflation target $(1 + C_\pi)$ greater than 1. This means that the interest rate has a positive relationship with inflation and shifts in the same direction with a change in inflation. The output gap (C_y) has to be positive and greater than 0.

$$i = 1 + 1.5\,\pi + 0.5\text{y} \qquad (5b)$$

In a strict sense, these values must be 1, 1.5, and 0.5. Deviations from these values explain the behavior of monetary policy. The low value of R-squared indicates a discretionary monetary policy and the low response of the central bank in controlling inflation (Tchaidze, 2001).

Clarida et al. (2000) have criticized the Taylor rule where the federal funds rate is a function of lagged inflation and the output gap. They suggest that the conduct of monetary policy changes with the shift in macroeconomic variables. Monetary policy rules must adjust the federal funds rate in accordance with expected inflation and output at their target rates. The central bank has not only to adjust the interest rate but also to predict the expected inflation rate. If there are expectations of high inflation, the central bank should take a proactive stance. In the authors' version of policy rules, the Taylor rule becomes a special case.

3.2. Baseline Reaction Function

Clarida et al (2000) formulate the following forward looking rule:

$$r_t^* = r^* + \beta\left(E\{\pi_{t,k}|\Omega_t\} - \pi^*\right) + \gamma\, E\,\{x_{t,q}|\Omega_t\} \qquad (6)$$

$\pi_{t,k}$ is the annual inflation rate measured as the percentage difference in the price level between two time periods *t* and *k*. π^* is the target inflation rate, $x_{t,q}$ is the output gap measured as the percentage difference of the log of real GDP, *E* is the expectations operator, Ω_t is the information operator, r_t^* is the federal funds rate, and r^* s the target nominal interest rate for targeted inflation and the output gap.

3.3. Implied Real Rate Rule

Clarida et al (2000) consider following implied real rate rule for the real interest rate target *rr*:

$$rr_t^* = rr^* + (\beta - 1)\left(E\{\pi_{t,k}|\Omega_t\} - \pi^*\right) + \gamma\, E\,\{x_{t,q}|\Omega_t\} \qquad (7)$$

$rr_t^* = r_t - E\{\pi_{t,k}|\Omega_t\}$ and $rr^* = r^* - \pi^*$ represent the real interest rate, which is stationary and determined by nonmonetary factors.

The benchmark rate for β is 1 and for γ is 0. The sign and magnitude both play an important role. When $\beta > 1$, the interest rate rule tends to stabilize the economy; when $\beta \leq 1$, the interest rate rule is likely to destabilize or accommodate shocks to the economy. If $\gamma > 1$, the economy is likely to be stabilized and when $\gamma \leq 1$, the economy will tend to stabilize. In each period, the central bank adjusts the funds rate as a function of the gap between the current target rate and a linear combination of past values of the interest rate.

$$r_t = \rho(L)r_{t-1} + (1-\rho)r_t^* \tag{8}$$

Continuing to follow Clarida et al (2000), inserting an interest rate-smoothing equation into the target equation yields

$$r_t = (1-\rho)\{rr^* - (\beta - 1)\pi^* + \{\beta\pi_{t,k} + \gamma\, x_{t,q}\} + \rho(L)r_{t-1} + \varepsilon_{t,} \tag{9}$$

Clarida et al (2000) assume the error term to be a linear combination of forecasted errors and is considered orthogonal, yielding:

$$E\{r_t - (1-\rho)\{rr^* - (\beta - 1)\pi^* + \{\beta\pi_{t,k} + \gamma\, x_{t,q}\} + \rho(L)r_{t-1}\}z_t = 0 \tag{10}$$

where z_t is a set of instruments—measured by the generalized method of moments (GMM)—with an optimal weighting matrix that accounts for possible serial correlation. When the number of instruments equals more than four restrictions, we can test the overidentification assumption. We then impose the restriction that the sample average should equal the real equilibrium interest rate. All other assumptions are the same as in Clarida et al. (2000).

3.4. Method of Estimation

We use the GMM to estimate the Taylor rule for Pakistan. This method is applied to a dataset where the shape of the distribution is not known (see Hansen, 1982). Given that GMM estimators are considered consistent, asymptotically normal, and efficient, we can replace the population parameter moment condition with its sample analogy. We also assume that the orthogonality condition holds, implying that the instruments used are exogenous and uncorrelated with the error term. Should the number of instruments exceed the number of parameters, there will be no unique solution; we would then assume an objective function by introducing a weighting matrix. Roodman (2009) also points out that, when the GMM uses weak instruments or too many instruments, its distribution is not reliable.

3.5. Specifications and Data

We use both annual and quarterly data to estimate the SBP's policy stance. The annual dataset extends from the fiscal year 1971 to the fiscal year 2011. The inflation rate is measured in terms of the GDP deflator, and the log of real GDP is used to measure the output gap as the percent difference between actual and potential GDP. Potential GDP is measured using the Hodrick-Prescott (HP) filter at power 4 (see Ravn & Uhlig, 2002). The discount rate is used as the interest rate in light of Harvey and Jaeger (1993) and Harvey and Trimbur's (2008) criticism that the HP filter uses an inappropriate smoothing constant. Our experiment shows that a default value of 16,000p^4 best explains the trend in this series.

First, we test the traditional Taylor rule equation, which is a linear combination of the interest rate, inflation rate, and output gap. We then test the same rule incorporating the modifications given in Clarida et al. (2000), using the GMM. We use four lags for the inflation rate and growth rate in broad money supply, the difference between the short-run and long-run interest rate spread, and the consumer price index (CPI) interest rate as instruments. These lags are selected on the basis of Hansen's J-statistic, which is used to test over-identified specifications. The difference-in-Hansen test is applied to check the orthogonality of weak instruments.

The second dataset, the quarterly data series, spans the calendar years 1970:Q1 to 2011:Q4. All data are taken from the International Monetary Fund's International Financial Statistics (2011). For this dataset, we use the quarterly inflation rate measured on the basis of the percent change in the CPI on a year-to-year basis. Since Pakistan does not yet have official quarterly estimates for GDP, we use the index of large-scale manufacturing as a proxy for output growth. Although this is a source of specification bias, various authors have used the same proxy due to the same constraint.[1]

The HP filtering technique is used to remove the trend when computing potential GDP. This is a standard data de-trending technique

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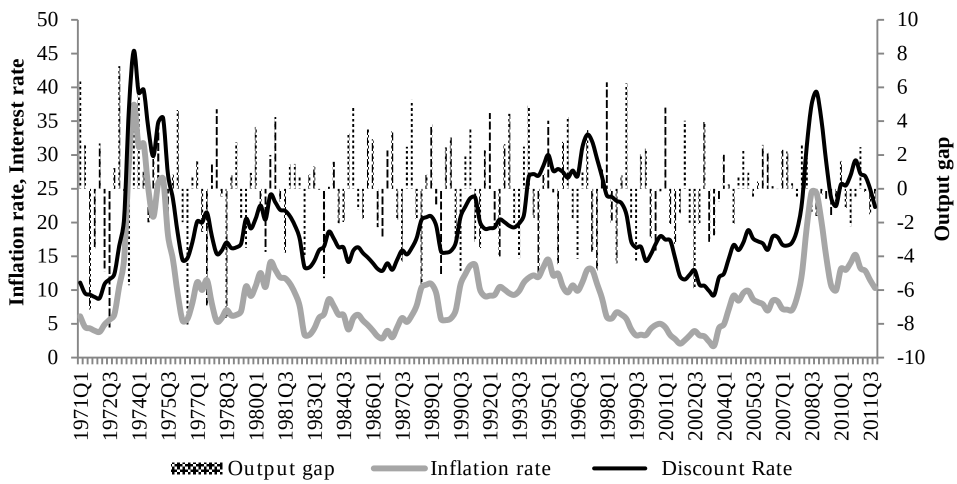
[1] Khan and Schimmelpfenning (2006) use an interpolated series of nominal and real GDP and a large-scale manufacturing index to measure economic activity. Ahmad and Ahmed (2006) also use the quantum index of industrial production as a proxy for GDP. They cite Shanmugam, Nair, and Li (2003) and Nell (2000/01) with regard to using these proxies.

that yields robust results. The call money rate, rather than the discount rate, is taken as the interest rate, given that the SBP used the fixed discount rate as a policy measure in the pre-reform period. The use of adjustments in the discount or policy rate arose after Hanfi's tenure.

4. Estimation

Figure 1 shows the cyclical nature of the relationship between the output gap, inflation rate, and interest rate. Even in the absence of shocks, these fluctuations prove to be persistent and self-fulfilling.

Figure 1: Cyclical behavior of output, inflation, and interest rate in Pakistan



4.1. Forward-Looking Monetary Policy Rules

We estimate the Taylor rule for Pakistan by using the GMM to analyze annual data for 1971–2011 on the interest rate, inflation rate, output gap, and interest rate-smoothing constant p. We then estimate the same model using quarterly data for 1971:Q1–2011:Q4, subdividing the data into the reform period 1989:Q1–2011:Q4. The target horizon used is one year for the annual data and one quarter for the quarterly data. The results are summarized in Table 1. Clarida et al. (2000) find that a value of less than 1 for the inflation rate indicates the failure of monetary policy in controlling inflation. It implies that the monetary policy was poorly designed as a result of incorrect estimates of the state of the economy. When $\beta \leq 1$, the interest rate rule tends to destabilize or accommodate shocks to the economy.

Table 1: Forward-looking monetary policy rules (baseline results)

Period	Inflation rate	Output gap	Constant	H–J $\chi 2$ P-value
1971–2011 (GMM)	0.319**	-0.143*	1.816***	8.07152
	(0.0063)	(0.077)	(0.0764)	(0.1523)
1971:Q1–2011:Q4 (GMM)	0.021**	0.030*	1.90***	21.35
	(0.007)	(0.0197)	(0.0964)	(0.0003)
1989:Q1–2011:Q4 (GMM)	0.049**	0.032*	1.637***	9.52302
	(0.010)	(0.014)	(0.160)	(0.0493)

Note: Standard errors are reported in parentheses. The set of instruments includes four lags for the inflation rate, short-long spread, CPI rate, and M2 growth rate. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.
Source: Author's calculations.

The value of the coefficient of the inflation rate is 0.32 with a standard error (SE) of 0.0063, i.e., below 1 and significant. The output gap is negative and has a value of –0.14 with an SE of 0.07. Hansen's J chi-square value is 8.07152 with a p-value of 0.1523, indicating interest rate inertia during 1971–2011. We conclude that there is no adjustment to the interest rate when the inflation rate is rising or when output deviates above the target. The absolute values of the output gap estimates do not justify the tightening of monetary policy.

The GMM results show that the SBP did not resort to an aggressive policy in order to control inflation. The central bank did not actively reduce the output gap and inflation rate with adjustments in the interest rate, which is important for influencing aggregate demand. We conclude that the SBP's monetary policy during 1971–2011 was discretionary and less active. All the variables are significant but their signs are not as expected. The coefficient of the output gap has a negative sign for the annual data, which is not in accordance with the Taylor rule or its modification by Clarida et al. (2000).

In the case of the quarterly data (1971:Q1–2011:Q4), which includes variables such as the CPI inflation rate, a large-scale manufacturing production index as a proxy for GDP, and the interest rate, the coefficient of the expected inflation rate is below unity (0.021, SE 0.007), i.e., far less than in the reform period (0.049, SE 0.10). Although significant, the low value indicates a chaotic monetary policy. The coefficient of the output gap is significant and positive, which points to the sensitivity of the cyclical variable, but it also remains low and is the same (0.03, SE 0.020). The value of the smoothing parameter *p* remains low (0.0003 and 0.04, respectively), implying that there is no interest rate inertia.

4.2. Backward-Looking Monetary Policy Rules

Taylor (1993a) presents the coefficient of the backward-looking inflation rate $(1+C_{\pi})>1$ as a necessary stability condition. If the coefficient has a value of 1.5, the response of the central bank matches the historical trend. In Table 2, we use ordinary least squares (OLS) to estimate the Taylor rule because Taylor (1993a) assumes a linear relationship between the interest rate and inflation rate and deviations in GDP. Since there is an indication of autocorrelation, we apply the augmented Dickey-Fuller (ADF) and L-Jung tests to the residuals; the tests show that the residuals are consistent. The specification for the Taylor rule computed on the basis of annual time series data (1971–2011) and employing the past behavior of monetary policy shows that the SBP has not followed the Taylor rule in controlling inflation in Pakistan.

Table 2: Backward-looking monetary policy rules

Period	Inflation rate	Output gap	Constant	Adjusted R-sq. (AC) ADF test
1971–2011 (OLS)	0.240**	-0.672*	6.099***	0.449
	(0.0052)	(0.167)	(0.617)	(1.08)
				-3.745
1971:Q1–2011:Q4 (OLS)	0.066**	0.096*	1.633***	0.65
	(0.022)	(0.039)	(0.433)	(2.33)
				-15.02
1989:Q1–2011:Q4	0.392**	0.150*	5.280***	0.302
	(0.063)	(0.0965)	(0.644)	(0.678)
				-4.39

Note: Standard errors are reported in parentheses. AC = auto correlation, ADF = augmented Dickey Fuller test. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.
Source: Author's calculations.

Our OLS results confirm those of Malik and Ahmed (2010). The coefficient of the inflation rate is 0.24 with an SE of 0.0052. The magnitude of the inflation rate (0.24) is less than 1.5, implying a less responsive monetary policy. The output gap has a negative sign, which not only contradicts the Taylor rule but also indicates that the SBP decreased the interest rate during high inflation regimes or vice versa. The overall R-squared value is low at 0.45, implying that monetary policy is unsystematic. The OLS results are consistent and significant.

When we apply the Taylor rule to the quarterly data using the large-scale manufacturing index as a proxy for real GDP, the magnitude is even smaller than for the annual data, indicating an even weaker monetary policy stance. However, the sign of the coefficient of the output gap is positive. Under the Taylor rule, a positive coefficient implies there is a high possibility that inflation will increase in the future; the adjustment between the output gap and interest rate reflects the use of a pre-emptive, cyclical policy. This is considered to be a short-run objective of growth without compromising long-term price stability in the economy.

Although the output gap is positive with respect to the real interest rate on the basis of the quarterly data, it is not significant and the residuals are autocorrelated. For robust results, we use an AR (1) model to eliminate the autocorrelation. We find there is a possibility of future inflation; the SBP appears to have preferred adopting a cyclical policy, giving precedence to short-run growth over long-run price stability. The coefficient is less than 0.5, indicating a cyclical and less aggressive policy.

4.3. Reform Period (Quarterly Time Series 1989:Q1–2011:Q4)

Our results for this period show that the inflation rate coefficient is less than 1 at 0.392. This means that the interest rate did not adjust fully to inflationary pressures in the economy and the SBP did not pursue a stable low-inflation objective. In other words, it did not exercise autonomy or practice "leaning against the wind" (Tchaidze, 2001). The coefficient of the output gap at 0.150 is positive, which shows that the increase in the output gap was cyclical and likely to increase the future inflation rate. However, as the value is low and insignificant, the output gap was not aggressive.

The overall R-squared value is very low at 0.302, indicating a very weak policy stance and confirming that the SBP adopted a. The residuals are consistent as shown by the ADF test. These results are not different from those for the annual data series except for the output gap sign, which is positive but small. Table 3 summarizes the regime-wise results for the Taylor rule for the period 1989:Q1–2011:Q4 with backward-looking and forward-looking monetary policy rules.

4.3.1. The Hanfi Regime 1989:Q1–1993:Q2

In the forward-looking model, the low but positive coefficient of the output gap (0.0182, SE 0.005) points to a sensitive, cyclical monetary policy. The coefficient of the inflation rate (0.015, SE 0.021) indicates an insignificant and weak monetary policy stance, with the likelihood of a higher rate of inflation in the future. The high value of *p* (0.47) shows a considerable degree of interest rate inertia. The backward-looking model yields an inflation rate coefficient that is less than 1 (–0.023, SE 0.105), indicating the lack of adjustment between the inflation and interest rates. Although its sign is different, it remains insignificant as in the forward-looking model. The relatively high value of R-squared indicates a systematic policy.

Table 3: Summary results for Taylor rule by monetary policy regime

Regime		Inflation rate	Output gap	Constant	H–J $\chi 2$ P-value R-sq.
Hanfi	GMM	0.015	0.0182***	1.828**	3.58
1989:Q1–1993:Q2		(0.021)	(0.0051)	(0.271)	(p = 0.47)
	OLS	-0.023	0.068**	0.358	0.43
		(0.219)	(0.209)	(0.384)	
Yaqub	GMM	0.0351	0.0124**	1.966***	4.46555
1993:Q3–1999:Q2		(0.022)	(0.004)	(0.234)	(p = 0.35)
	OLS	0.250	0.021*	0.152	0.17
		(0.188)	(0.011)	(0.238)	
Husain	GMM	-0.0489*	0.012	2.01***	13.41
1999:Q3–2005:Q3		(0.034)	(0.021)	(0.196)	(p = 0.009)
	OLS	0.154	-0.0118	-0.186	0.078
		(0.1797)	(0.011)	(0.186)	
Akhtar	GMM	0.030***	0.011	1.889***	0.888
2005:Q4–2008:Q3		(0.007)	(0.012)	(0.123)	(p = 0.93)
	OLS	0.162***	-0.008	0.117	0.85
		(0.026)	(0.006)	(0.701)	
Present regime	GMM	0.0119***	(-0.009)	2.30**	4.22
2008:Q4–2011:Q4		(0.003)	(0.006)	0.044	(p = 0.38)
	OLS	0.184**	-0.0096	0.318	
		(0.102)	(0.032)	(0.295)	0.32

Note: Standard errors are reported in parentheses. The forward-looking model includes four lags for the inflation rate, the short-long spread, CPI rate, and M2 growth rate. * p < 0.05, ** p < 0.01, *** p < 0.001.
Source: Author's calculations.

4.3.2. The Yaqub[2] Regime 1993:Q3–1999:Q2

Monetary policy under Yaqub was no different from that under his predecessors where inflation was concerned: the coefficient of the inflation rate in the forward-looking model is less than 1 (0.035 with SE 0.022) and the coefficient of the output gap is positive and significant but has a low value (0.0124 with SE 0.004). The p-value in the last column is high and, again, shows interest rate smoothing. The results of the backward-looking model are insignificant but the direction of monetary policy is the same. Under the Yaqub regime, therefore, the SBP preferred price stability to short-run growth since the coefficient of the inflation rate increased from 0.02 to 0.03 but did not fully adjust prices. The value of R-squared reflects a less systematic monetary policy stance.

4.3.3. The Husain Regime 1999:Q3–2005:Q3

In the forward-looking model, the coefficient of the inflation rate is negative and insignificant (–0.0489 with SE 0.034) and the coefficient of the output gap, which measures sensitivity to the cyclical variable, is also insignificant. The interest rate smoothing value is low, showing interest rate inertia. As indicated by the low value of R-squared, the policy adopted was less active in achieving long-run price stability. The negative value for the output gap variable reflects a lack of commitment with respect to the short-run output gap in the economy.

4.3.4. The Akhtar Regime 2005:Q4–2008:Q3

Akhtar's regime was different from previous regimes in that it has the highest value of $p = 0.93$. This implies considerable interest rate inertia although the coefficient of the inflation rate is low and significant. The output gap remains insignificant. The backward-looking model yields similar results. The high R-squared term shows that the Akhtar regime's monetary policy was systematic. However, the value of the inflation rate coefficient is less than 1 (0.162), which indicates a less aggressive policy. The output gap is negative and insignificant. The SBP's policy was, therefore, less active in attaining the goal of price stability and output gap, although the high value of R-squared implies an aggressively tight monetary policy.

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[2] It is worth noting that, under President Leghari's caretaker regime in 1996, a committee was set up under Yaqub to review the State Bank of Pakistan Act, which changed the tenure of the governor from one five-year term to two three-year terms. While one five-year term allows the governor relative independence from the start of the tenure, the alternative makes the governor vulnerable insofar as he or she has to seek a second term.

4.3.5. The Present Regime 2008:Q4–2011:Q4

In this case, the value of the inflation rate is still low but significant (0.011 with SE 0.003) and the output gap is negative and insignificant (–0.009 with SE 0.006). The smoothing parameter is considerably high but lower than that under the Akhtar regime. In the backward-looking model, the inflation rate is less than 1 (0.184), implying there was no adjustment in the real interest rate and inflation rate. The output gap is negative, which confirms the use of an anti-cyclical policy. The low value of R-squared shows the policy was discretionary. Here, the monetary policy failed to control inflation and there was little adjustment between the interest rate and inflation rate. The output gap is negative, showing the lack of sensitivity to the cyclical variable.

The results of the forward-looking model indicate that monetary policy remained almost the same throughout the reform period (1989:Q1–2011:Q4). The coefficients of the inflation rate and output gap remain low. There is considerable interest rate inertia, confirming the interest rate-smoothing hypothesis, but the coefficient of the inflation rate remains low. Husain's period is characterized as chaotic, with insignificant coefficients with the wrong signs, and no interest rate smoothing.

4.4. Chow Test Results: A Structural Break

The Chow test is used to examine whether the interest rate for the inflation rate and output gap is the same before or after a specific regime or across various regimes. The test is applied to various SBP regimes to assess any differences in policy: 1993:Q3, 1999:Q3, 2005:Q4, and 2008:Q4. The results show that the coefficients are not stable across the Hanfi and Yaqub regimes, while the Hussain, Akhtar, and present regimes are structurally stable. A comparison of the backward-looking model and forward-looking model confirms the robustness of the results. In both cases, the signs and magnitudes of the parameters remain the same and consistent. The value of the inflation rate remains less than 1 for all regimes and the output gap switches from positive to negative for the present regime in both models.

On the whole, the SBP has always adopted inflationary policies but failed to adjust inflation with the real interest rate. Nominal interest rates have remained rigid in relation to the movement of the inflation rate. The low value of R-squared reflects a disordered policy stance, which can be attributed to incorrect estimates of the state of the economy.

Akhtar adopted a tighter monetary stance that was organized but less active. This confirms that the SBP's monetary policy is backward looking and is not based on any particular rule. Commitment to policy will remain low if policymakers focus solely on their short-run aspirations.

5. The Rules Path

The Taylor rule demonstrates that a 2 percent inflation or interest rate is the best path available for safeguarding the output gap and inflation rate objective. Once we have fixed either the interest rate or the inflation rate at 2 percent, we can calculate the other as follows:

$$\hat{\pi}^* = \frac{r^* - \hat{C}}{\hat{C}_\pi} \qquad \hat{r}^* = \hat{C} + \hat{C}_\pi \pi^*$$

The results in Table 4 show that the SBP has not followed any particular rule in determining the inflation rate or interest rate relationship other than a "rule of thumb." Further, there is no evidence of Taylor's "double duex" assumption of a 2 percent inflation or interest rate relationship in Pakistan. On average, the real interest rate has been very low—a reflection of the controlled financial market. The average inflation rate was 8.83 percent, which shows that price stability has not been seriously pursued by the SBP. One indication of this is that, under Hanfi and Akhtar, real interest rates were on the rise, not as a policy decision but because of the deteriorating fiscal situation in the economy.

Table 4: Implied interest rate and inflation targets

Specification	$r^* \Rightarrow$	$\hat{\pi}^*$	$\pi^* \Rightarrow$	$\hat{r}^*$	Nominal interest rate	Actual π	Real interest rate
1971–2011	22	-1.80	22	-1.86	10.63	9.29	1.34
1989:Q1–2011:Q1	22	0.37	22	12.94	11.97	9.09	2.89
Hanfi 1989:Q1–1993:Q2	22	0.311	22	-71.4	10.00	9.57	0.425
Yaqub 1993:Q3–1999:Q2	22	-1.31	22	2.45	15.90	10.05	5.85
Husain 1999:Q3–2005:Q3	22	-1.88	22	-2.58	9.68	4.78	4.90
Akhtar 2005:Q4–2008:Q3	22	1.56	22	-2.24	10.04	10.54	-0.50
Present regime 2008:Q4–2011:Q4	22	-2.06	22	-1.31	13.55	14.88	-1.34

Source: Author's calculations.

If the inflation rate is targeted around an average inflation rate of 9 percent and the potential GDP growth rate also averages around its trend (5.4 percent during 1971–2011), the macroeconomic quadratic loss function can be estimated as follows:

$$Social\ Loss = [2(Ln(\pi - \pi^*)^2) + Ln(q - q^*)^2] \quad (12)$$

where $Ln(\pi - \pi^*)^2)$ denotes $\sigma^2\pi$ and $Ln(q - q^*)^2$ denotes $\sigma^2 y$.

Table 5 shows the social loss that results from keeping the interest rate lower than the optimal rate. In developing economies such as Pakistan, policymakers tend to argue that the interest rate must be kept low in order to maximize output, but this study shows that it increases the variability of inflation and output. The objective of interest rate smoothing and the low value of R-squared implies a less aggressive policy.

Table 5: Social loss 1971–2011

Specification	$\boldsymbol{\sigma^2 y}$	$\boldsymbol{\sigma^2\pi}$	$\boldsymbol{SL = \sigma^2 y + \sigma^2\pi}$
1971–2011	4.68	3.71	13.08
1989:Q1–2011:Q1	4.35	1.22	9.93
Hanfi 1989:Q1–1993:Q2	0.78	0.37	1.95
Yaqub 1993:Q3–1999:Q2	2.36	0.38	5.09
Husain 1999:Q3–2005:Q3	3.37	0.96	7.72
Akhtar 2005:Q4–2008:Q3	10.95	0.69	22.58
Present regime 2008:Q4–2011:Q4	2.51	0.26	5.28

Source: Author's calculations.

6. Conclusion

Reforms in Pakistan's financial and monetary sector began in 1989. This paper provides empirical estimates of the rule-based monetary policy applied during the pre- and post-reform periods under different regimes at the SBP. We compare the effectiveness of this policy under backward-looking as well as forward-looking reaction functions. We find that the inflation rate produced mixed results with growth, and that fiscal dominance made it difficult to pursue a discretionary monetary policy.

Our estimates indicate one minor difference between the pre- and post-reform periods. Generally, the monetary policy was characterized by a less aggressive response and was neither forward looking nor backward looking. In the pre-reform period, however, the SBP was less likely to raise the interest rate in response to an increase in inflation. In the post-reform period, it was more likely to raise the interest rate in response to past inflation trends and expected inflation, but not to the extent of the required adjustment.

The results show that the Taylor-type rule does not explain interest rate adjustment in Pakistan over the period of study, given the inconsistencies we see in the interest rate and inflation rate. The SBP has not used a Taylor-type rule to maintain low inflation and stable output. Stable output is important in the short run to prevent future inflationary expectations. The structural break during the Hanfi and Yaqub regimes shows that their policies were different from those of their predecessors. Husain's regime generally accommodated the interest rate. Hanfi and Akhtar's regimes showed some convergence of the interest rate toward the real interest rate, although it remained below the actual real interest rate.

Overall, the SBP has not used the short-run interest rate to control long-run inflation and ensure stable output growth. Its policy stance reflects neither a forward-looking nor backward-looking model. The indeterminate relationship between the interest rate and inflation rate resulted in an output gap that has had social implications. An important implication for policy is that the SBP should use a modified rules-based approach to avoid fiscal dominance.

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Appendix

Chow forecast test: Predictions for observations from 1993

F-statistic	2.597698	Prob. F (3, 34)	0.0600
Log likelihood ratio	8.254823	Prob. chi-square (3)	0.0410
Wald statistic	7.793094	Prob. chi-square (3)	0.0505

Test predictions for observations from 1993 to 2010

	Value	Df	Probability
F-statistic	20.95344	(18, 19)	0.0000
Likelihood ratio	121.4954	18	0.0000

F-test summary

	Sum of squares	Df	Mean squares
Test SSR	107.0749	18	5.948607
Restricted SSR	112.4690	37	3.039701
Unrestricted SSR	5.394033	19	0.283896
Unrestricted SSR	5.394033	19	0.283896

LR summary test

	Value	Df
Restricted LogL	-77.43350	37
Unrestricted LogL	-16.68582	19

Unrestricted log likelihood adjusts test equation

Variable	Coefficient	Std. error	t-statistic
C	0.505661	0.259490	1.948676
D (DEF)	-0.222812	0.178928	-1.245264
D (YGAP)	0.001393	0.005277	0.263974
R-squared	0.080087	Mean dependent var.	0.227273
Adjusted R-squared	-0.016745	SD dependent var.	0.528413
SE of regression	0.532819	Akaike info criterion	1.704855
Sum squared residual	5.394033	Schwarz criterion	1.853634
Log likelihood	-15.75341	Hannan-Quinn criterion	1.739903
F-statistic	0.827068	Durbin-Watson stat.	1.114722

Quarterly data: Chow breakpoint test

Equation sample: 1989:Q1–2011:Q4

F-statistic	3.860546	Prob. F (3, 83)	0.0122
Log likelihood ratio	11.62542	Prob. chi-square (3)	0.0088
Wald statistic	11.58164	Prob. chi-square (3)	0.0090

Equation sample: 1999:Q3

F-statistic	5.983896	Prob. F (3, 83)	0.0010
Log likelihood ratio	17.42633	Prob. chi-square (3)	0.0006
Wald statistic	17.95169	Prob. chi-square (3)	0.0005

Equation sample: 2005:Q3

F-statistic	2.437183	Prob. F (3, 83)	0.0704
Log likelihood ratio	7.513802	Prob. chi-square (3)	0.0572
Wald statistic	7.311549	Prob. chi-square (3)	0.0626

Equation sample: 2008:Q4

F-statistic	0.025386	Prob. F (3, 83)	0.9945
Log likelihood ratio	0.081625	Prob. chi-square (3)	0.9939
Wald statistic	0.076157	Prob. chi-square (3)	0.9945

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Estimating Firms' Vulnerability to Short-Term Financing Shocks: The Case of Foreign Exchange Companies in Pakistan

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Abstract

Using firm-level balance sheet data for 20 of the 24 exchange companies in Pakistan for the period 2006–11, we explore the sources of firms' vulnerability to short-term financing shocks. Based on the probability estimates of a maximum likelihood binary probit model, this paper shows that the incidence and degree of vulnerability of foreign exchange companies to short-term financing shocks has risen significantly over time. If not managed opportunely, these shocks can cumulate into long-term financing shocks and even lead to corporate failure in the long run. Our regression results show that the corporate managers of these companies cannot ignore macroeconomic factors such as global changes and the macroeconomic environment (inflation and GDP growth) in addition to firm-specific factors (growth opportunities, firm size, permanent earnings, earnings volatility, and working capital management) when managing their firms' vulnerability to short-term financing shocks.

Keywords: Foreign exchange companies, vulnerability, financing shocks, Pakistan.

JEL classification: G33, G30 L80, L89.

1. Introduction

Generally, companies' financial reporting entails documenting three types of cash flows: operating, investing, and financing flows. Operating cash flows arise on account of a company's normal business operations and are crucial because they indicate a firm's capacity to generate sufficient positive cash flows to maintain and expand its operations. In case of corporate failure, firms may require external financing, consequently raising the probability of rolling over their current liabilities and becoming exposed to funding shocks.

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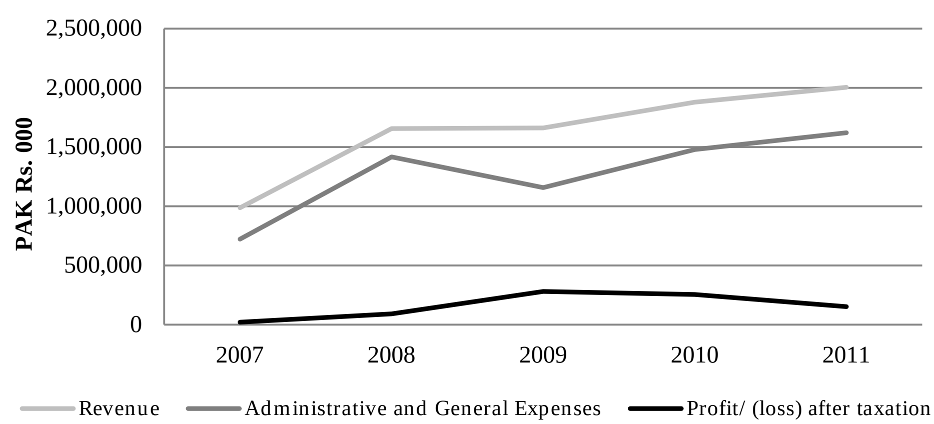
These funding shocks, if not managed well in time, can erode the firm's capital to a potentially dangerous extent (Tudela & Young, 2003b) and lead to corporate bankruptcy or failure. However, at the root of these corporate failures is the inability of firms to overcome or manage their response to short-term funding shocks. Understanding the sources of vulnerability to such shocks is, therefore, critical for corporate managers as well as for investors seeking credit exposure (see Chan-Lau & Gravelle, 2005) to such vulnerable companies. Initially, these funding shocks may weaken a firm's liquidity, but they can also affect the solvency of the corporate sector, potentially destabilizing the economy as a whole if the sector is important. Understanding the sources of vulnerability of foreign exchange companies to short-term financing shocks is crtical, given the foreign exchange exposure of the financial and real sectors in the Pakistan's current exchange rate crisis.

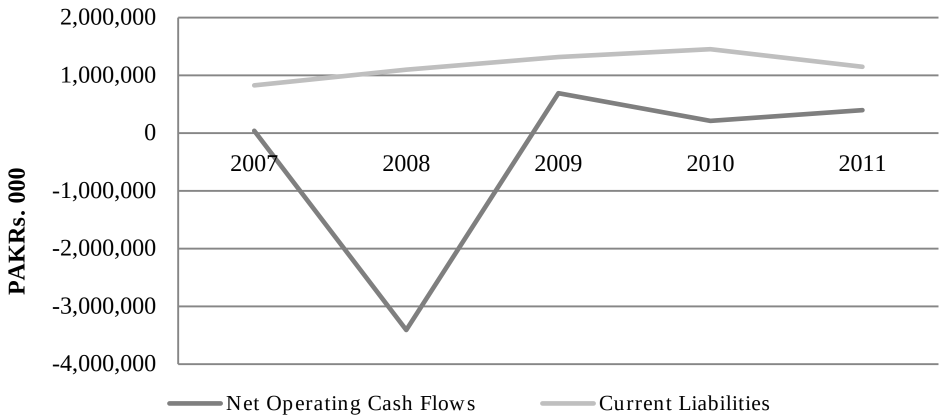
We observe a relatively significant and high degree of volatility in these firms' net operating cash flows compared to their revenue, administrative, general expense, and profit-after-tax flows during the period 2006–11 (Figures 1 and 2). Net operating cash flows are consistently below current liabilities (Figure 2) and provide evidence of the vulnerability of foreign exchange companies to short-term financing shocks in Pakistan. This evidence forms the basis for this study. Currently, there is scant literature on the vulnerability of exchange firms to short-term funding shocks and no attempt has been made to explore the sources of this vulnerability in the context of Pakistan. Accordingly, this study aims to fill this gap in the literature.

Figure 1: Selected indicators of foreign exchange industry



Note: All values are annual aggregates.
Source: State Bank of Pakistan.

Figure 2: Vulnerability of foreign exchange industry to short-term financial shocks



Note: All values are annual aggregates.

Source: State Bank of Pakistan.

The paper is organized as follows: Section 2 reviews the literature; Section 3 describes the data sources and variables employed, and discusses the study's research design and methodology; Section 4 presents our results; and Section 5 concludes the study.

2. Review of the Literature

Hong and Wu (2013) estimate a discrete-time hazard model of bank failure using data on US commercial banks for the period 1985–2004. They identify idiosyncratic and systemic funding liquidity risks as a major predictor of bank failures in 2008 and 2009.[1] Friend and Levonian (2013) conclude that using a market-based measure of capital (or leverage) allows one to predict bank failures farther in advance, thus providing more time to respond and reduce the cost of such failures.

Predictions of corporate banking failures are well documented in the literature. Among the models that attempt to develop an understanding of the factors that lead to bank failure are discriminant analysis (Stuhr & Van Wicklen, 1974; Pettway & Sinkey, 1980), factor analysis, probit and logit regression models (West, 1985; Gajewski, 1989; Thomson, 1991; Reynolds, Fowles, Gander, Kunaporntham, & Ratanakomut, 2002), event-history analysis (Looney, Wansley, & Lane, 1989), market data analysis (Demirgüç-

[1] Out-of-sample forecast.

Kunt, 1989), the proportional hazards model (Whalen, 1991), and belief networks (Sarkar & Ghosh, 1998).

Thomson (1991) uses a logit model to estimate the probability of bank failure, which, he argues, is a function of solvency, capital adequacy, asset quality, management quality, earnings performance, and the relative liquidity of the firm's portfolio. Reynolds et al. (2002) estimate the probability of failure in the context of Thailand's financial companies by applying a probit model to a panel of 91 financial firms. They relate the financial crisis to massive borrowing by banks and other financial institutions abroad and to risky housing loans.

Martin (1977) uses a logit model to estimate bank failure predictions based on data for the period 1970–76. He concludes that higher returns on assets and the capital-to-assets ratio reduce the likelihood of bank failure while a higher proportion of commercial and industrial lending raises the probability of bank default. Hanweck (1977) uses a probit model to predict bank failure based on data for the period 1973–75. He reports that the higher growth of total assets, returns on assets, and capital-to-assets ratios reduce the probability of bank failure while greater financial leverage and larger firm size raise the probability of bank default.

Similarly, Pantalone and Platt (1987) use a logit model to estimate bank failure predictions for the period 1983–85. They find that higher returns on assets and the capital-to-assets ratio reduce the likelihood of bank failure while a higher growth rate for residential lending, the proportion of commercial and industrial lending, and overall financial leverage raise the probability of bank default.

Merton (1974) introduced the idea of quantitatively modeling credit risk to show how the probability of default for an individual firm can be deduced from its market valuation. Beaver (1966) and Altman (1968) show that financial variables can be used to predict firms' liquidation. Ohlson (1980) estimates the likelihood of bankruptcy for nonfinancial firms in the US, using both logit and probit models. His results point to the statistically significant and positive impact of financial structure (total debt-to-total-liabilities ratio) and the negative impact of firm performance (return on total assets), size, and current liquidity (current ratio, working-capital-to-total-assets ratio, and the ratio of operating cash flows to current liabilities) on the probability of failure.

Gilbert, Krishnagopal, and Schwartz (1990) use the multinomial logit technique to estimate probabilities of default and then use these probabilities to classify their sample of nonfinancial firms into two groups, i.e., bankrupt and nonbankrupt. Poston, Harmon, and Gramlich (1994) similarly classify firms into three groups, i.e., turnarounds, business failures, and survivors.

A vulnerable corporate sector can transmit and/or magnify real or financial shocks to the extent that they may weaken overall macroeconomic resilience (González-Miranda, 2012). For policymakers and investors seeking credit exposure, it is therefore important to quantify the probability of such defaults (Chan-Lau & Gravelle, 2005). González-Miranda examines the corporate sector vulnerabilities of individual nonfinancial firms in a situation where their financing is brought to a halt. The study applies a probit model to a panel of 18 countries for the period 2000–11 and finds that higher leverage and maturity exposure raise a firm's probability of exposure to a funding shock, while larger firms with access to buffers are less vulnerable. Greater exchange rate flexibility, however, helps mitigate corporate vulnerability to financing shocks by encouraging hedging (see Cowan, De Gregorio, Micco, & Neilson, 2008; Patnaik & Shah, 2010; Kamil, 2012). Tirapat and Nittayagasetwat (1999) hold that macroeconomic conditions are also critical indicators of a potential financial crisis; investigating a sample of Thai listed firms, they show that the higher a firm's sensitivity to inflation, the greater will be its exposure to financial distress.

It is interesting to note that, while predictions of nonfinancial corporate default, bank failure, and overall financial distress have been widely debated in the literature, the vulnerability of firms to short-term funding shocks (which are likely to cumulate and can lead to organization failure) has received little attention, especially in the context of Pakistan. This study aims to fill this gap in the literature.

3. Methodology

This section describes the methodology, variables, and data sources used in the study.

3.1. Research Design

We use a probit model which serves well as a classification tool (see Langley & Sage, 1994). Drawing on Hanweck (1977), Martin (1977), West (1985), Pantalone and Platt (1987), Gajewski (1989), Thomson (1991), and

Reynolds et al. (2002), we take vulnerability to short-term financing shocks $[VFS_{i,t}]$ as the dependent variable, i.e., $VFS_{i,t}$ is equal to 1 when current liabilities $[CL_{i,t}]$ exceed the net operating cash flows $[NOCF_{i,t}]$ of firm i in year t and 0 when current liabilities $[CL_{i,t}]$ are either less than or equal to net operating cash flows $[NOCF_{i,t}]$.

We use this form of the dependent variable $[VFS_{i,t}]$ to estimate a maximum likelihood (ML) binary probit model because the dependent variable can then take on a binary form with respect to the presence or absence of financing shocks. This helps us in estimating the probability of a firm being subject to financing shocks. We model the probability $[Pr]$ of vulnerability to a financing shock of $[VFS_{i,t}]$ as follows:

$$\Pr\big(VFS_{i,t} = 1 \mid \mid [(X_{i,t}\beta\big), (Z_t\gamma)] = 1 - F(-X_{i,t}^{'}\,\beta - Z_t^{'}\,\gamma) \qquad (1)$$

where $X_{i,t}$ is a vector of firm-specific explanatory variables that vary across firms as well as over time; $Z_t$is a vector of explanatory variables that vary only over time; and F is a continuous, strictly increasing function that takes on a real value and returns a value ranging from 0 to 1. We assume that the index specification is linear in the parameters so that it takes the form $(X_{i,t}^{'}\beta)$ and $(Z_t^{'}\gamma)$ respectively. The choice of function determines the type of binary model. It follows that:

$$\Pr(VFS_{i,t} = 0 \mid \mid [(X_{i,t}\beta),\ (Z_t\gamma)] = F(-X_{i,t}^{'}\,\beta - Z_{t\,t}^{'}\gamma) \qquad (2)$$

Based on this specification, we can estimate the parameters of this model using the ML method. The likelihood function is written as:

$$l(\beta)=\sum_{i=1}^{n} VFS_{i,t} log\,[1 - F(-X_{i,t}^{'}\,\beta - Z_t^{'}\gamma)] + (1 - VFS_{i,t}\,) log[F\,(-X_{i,t}^{'}\,\beta - Z_t^{'}\gamma)] \qquad (3)$$

We code the values of $VFS_{i,t}$ as follows:

$$VFS_{i,t} = \begin{cases} 1\ if\ CL_{i,t} > NOCF_{i,t} \\ 0\ if\ CL_{i,t} \le NOCF_{i,t} \end{cases} \qquad (4)$$

This implies that the binary model of firms' vulnerability to financing shocks takes the form

$$VFS_{i,t} = 1 - F\big(-X_{i,t}^{'}\,\beta - \ Z_t^{'}\gamma\big) + \varepsilon_i \qquad (5)$$

where $\varepsilon_i$is a residual representing the deviation of the binary $VFS_{i,t}$ from its conditional mean.

3.2. Choice and Description of Variables

We use a dummy variable for vulnerability to financing shocks [$VFS_{i,t}$], which is equal to 1 if net operating cash flows are less than current liabilities and 0 otherwise.

Permanent earnings are likely to reduce $VFS_{i,t}$ while earnings volatility will increase it. We use the return on assets as a measure of profitability (see Martin, 1977; Ohlson, 1980) calculated as follows:

$$ROA_{i,t} = \frac{NPAT_{i,t}}{TA_{i,t}} * 100 \tag{6}$$

If we view the current earnings $\left(ROA_{i,t}\right)$ of firm i at time t as the sum of permanent ($PROA_{i,t}$) and earnings volatility ().), this yields

$$ROA_{i,t} = PROA_{i,t} + RROA_{i,t} \tag{7}$$

We use the following simple technique (see Hussain, 2013) to isolate permanent earnings ($PROA_{i,t}$) from current earnings($ROA_{i,t}$):

Step 1: We regress current earnings $ROA_{i,t}$ on current earnings ($ROA_{i,t-1}$) lagged by one year ($ROA_{i,t-1}$) in the following form:

$$ROA_{i,t} = \alpha + \beta * ROA_{i,t-1} + \mu_{i,t} \tag{8}$$

where $\mu_{i,t}$ represents transitory earnings $RROA_{i,t}$

Step 2: We create a series of residuals ($\mu_{i,t}$) based on the results of equation (8) above to capture earnings volatility or the risk factor ($RROA_{i,t}$).

Step 3: We subtract the residual series [($RROA_{i,t}$)] obtained in Step 2 from the series of current earnings ($ROA_{i,t}$) to obtain the permanent component of earnings ($PROA_{i,t}$).

González-Miranda (2012) reports the negative impact of relative firm size (RFS) on the likelihood of funding shocks. Therefore, we expect a negative coefficient with size and measure it as follows:

$$RFS_{i,t} = \frac{TA_{i,t}}{\sum_{i=1}^{n} TA_{i,t}} * 100 \tag{9}$$

where $TA_{i,t}$ denotes the book value of the total assets of firm i at time t while $\sum_{i=1}^{n} TA_{i,t}$ denotes the book value of the total assets of the industry comprising n firms.

Better working capital management is likely to reduce funding shocks or, alternatively, to raise vulnerability (Hong & Wu, 2013). We choose the current ratio $[CR_{i,t}]$ as a proxy for working capital management, which is calculated as follows:

$$CR_{i,t} = \frac{CA_{i,t}}{CL_{i,t}} \tag{10}$$

where $CA_{i,t}$ and $CL_{i,t}$ represent the book value of current assets and current liabilities of firm i in year t.

Growing firms are likely to be more vulnerable to financing shocks on account of their larger funding needs for growth. Therefore, we expect a negative coefficient with growth opportunities [log(TA)]. We use the logarithm of the book value of assets as a proxy for growth opportunities.

In addition to firm-specific variables, we also include three macroeconomic variables (see Tirapat & Nittayagasetwat, 1999): inflation and GDP growth to capture macroeconomic effects and the nominal effective exchange rate (NEER) to capture the effect of global changes. The NEER serves as a good proxy for global changes because cash flows are likely to be highly influenced, given that most of the current assets and liabilities of foreign exchanges companies are denominated in foreign currencies.

3.3. Dataset

We use secondary data from the State Bank of Pakistan's balance sheet analysis of the financial sector. Our sample covers 20 of the 24 exchange companies operating in Pakistan for which a complete data series for the period 2006–11 is available. Four companies were dropped from the sample on account of incomplete or inconsistent data series. The data on macroeconomic indicators is derived from the State Bank of Pakistan's *Handbook of statistics on Pakistan economy 2010* and the *Statistical bulletin* for 2012.

4. Results and Discussion

Table 1 presents summary statistics that reveal some interesting facts. The NEER and current ratio of exchange companies (CR) are subject to a very high degree of volatility. When firms are exposed to funding shocks due to the volatility of exchange rates, they will fight to manage these short-term shocks by adjusting their working capital. The high degree of inflation volatility also highlights the extent of uncertainty in the economy as a whole, which, in turn, is likely to increase the exposure of exchange firms to short-term funding shocks.

Table 1: Summary statistics

	VFS	LOG (TA(-1))	RFS	PROA	RROA (-1)	D (CR)	NEER	INF (-1)	D (GDPG)
Mean	0.684	12.424	4.584	0.020	0.002	1.000	58.306	13.084	-0.961
Median	1.000	12.309	3.500	0.018	0.002	-0.040	58.777	12.000	-2.000
Maximum	1.000	13.701	14.500	0.098	0.212	242.310	70.298	20.770	1.400
Minimum	0.000	11.522	1.600	-0.115	-0.247	-376.940	50.025	7.770	-3.100
SD	0.468	0.495	2.550	0.027	0.062	62.797	7.684	4.810	1.749
Skewness	-0.789	0.673	1.828	-1.323	-0.531	-1.546	0.619	0.696	0.137
Kurtosis	1.623	3.401	6.452	11.785	9.530	21.968	1.906	2.082	1.489
Jarque-Bera	14.44	6.490	83.22	277.05	144.09	1215.80	8.98	9.16	7.76
Probability	0.001	0.039	0.000	0.000	0.000	0.000	0.011	0.010	0.021
Observations	79	79	79	79	79	79	79	79	79

Source: Author's calculations.

Categorical descriptive statistics for the explanatory variables are presented in Table A1 in the Appendix. The dependent variable frequencies given in Table A2 indicate the presence of short-term financing shocks in 68 percent of the observations in the sample. This evidence fully supports the aims of this study.

Table 2 presents the regression results of the ML-binary probit model. The estimates show that lagged growth opportunities and changes in liquidity significantly (at 10 and 5 percent, respectively) reduce the probability of firms' exposure to short-term financing shocks; larger firms, however, are more likely to face such shocks. A one-percent improvement in growth opportunities reduces vulnerability by almost 1.70 percent. Earnings volatility raises vulnerability (though insignificantly) while permanent earnings significantly (at 10 percent) reduce vulnerability to short-term funding shocks. Changes in the global and local macroeconomic environment significantly (at 5 percent) increase firms' vulnerability while inflation significantly (at 5 percent) reduces the probability of this vulnerability.

Table 2: Regression results

Method: ML-binary probit (quadratic hill climbing)				
Dependent variable: VFS				
Sample period: 2006–11				
Variable	**Coefficient**	**SE**	**z-stat.**	**Prob.**
C: constant	15.6726	10.6060	1.478	0.1395
LOG (TA (-1)): growth opportunities	-1.6988	0.8702	-1.952	0.0509
RFS: relative firm size	0.4320	0.1799	2.401	0.0163
PROA: permanent earnings	-39.9871	22.5336	-1.775	0.0760
RROA (-1): earnings volatility	14.8642	9.5469	1.557	0.1195
D (CR): liquidity	-0.0116	0.0053	-2.182	0.0291
NEER: global changes	0.1615	0.0522	3.095	0.0020
INF (-1): inflation	-0.2436	0.1065	-2.287	0.0222
D (GDPG): macroeconomic environment	1.3279	0.4287	3.097	0.0020
McFadden R-squared	0.2800	SE of regression		0.4148
LR statistic	27.6119	Log likelihood		-35.503
Prob. (LR statistic)	0.0006	Avg. log likelihood		-0.4494
		Avg. log likelihood		-0.4494
Observations with dep. = 0	25	Total obs.		79
Observations with dep. = 1	54			

***Source*:** Author's calculations.

Using the ML-binary probit model, we have estimated the vulnerability (probability) of firms to financing shocks for various years. Our estimates reveal that 16 out of 20 firms have a 5 percent or higher probability of being vulnerable to short-term financing shocks in 2011, compared to only 1 out of 20 firms in 2008 (Table A3 in the Appendix). This clearly shows that the incidence and degree of vulnerability of foreign exchange firms to short-term financing shocks has risen over time.

The results of the expectation-prediction evaluation test for the model's binary specification (Table A4 in the Appendix) show that the estimated equation yields prediction-expectation values that are 100 percent and 78.45 percent, respectively, correct for the presence of short-term financing shocks (dep. = 1) at a success cutoff rate of 5 percent or higher.

5. Conclusion and Policy Implications

Our regression results have shown that growth opportunities, permanent earnings, working capital management, and changes in inflation reduce the probability of financing shocks while firm size, and global (NEER) and macroeconomic changes have a positive and significant impact on the vulnerability of foreign exchange companies to short-term financing shocks in Pakistan. In view of these results, corporate managers of exchange rate companies cannot afford to ignore either macroeconomic and global factors or firm-specific factors in managing the vulnerability of their firms to short-term financing shocks.

About 80 percent of the firms in our sample have a 5 percent probability or higher of being vulnerable to short-term financing shocks in 2011, compared to only 1 percent in 2008. This provides evidence of a significant rise in the incidence and degree of vulnerability of foreign exchange companies to financing shocks over time.

If not managed well in time, these short-term financing shocks can cumulate into long-term shocks and lead to corporate failure of exchange rate companies in the long run through financial and real sector effects. This, in turn, can have serious impacts on an uncertain economy. Therefore, understanding the sources of vulnerability of this sector to short-term financing shocks is critical for policy advisors, investors seeking credit exposure to such vulnerable companies, and corporate managers.

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Appendix

Table A1: Categorical descriptive statistics for explanatory variables

	Mean		
Variable	**Dep. = 0**	**Dep. = 1**	**All**
C	1.000000	1.000000	1.000000
LOG (TA (-1))	12.43354	12.42026	12.42446
RFS	4.052000	4.829630	4.583544
PROA	0.023412	0.017771	0.019556
RROA (-1)	0.002351	0.001565	0.001813
D (CR)	25.48960	-10.33833	0.999620
NEER	56.34568	59.21406	58.30634
INF (-1)	13.16000	13.04944	13.08443
D (GDPG)	-0.844000	-1.014815	-0.960759
	Standard deviation		
Variable	**Dep. = 0**	**Dep. = 1**	**All**
C	0.000000	0.000000	0.000000
LOG (TA (-1))	0.342306	0.554679	0.495125
RFS	1.426861	2.907221	2.549883
PROA	0.029635	0.025801	0.027010
RROA (-1)	0.063207	0.062131	0.062068
D (CR)	70.63707	55.94860	62.79688
NEER	6.427661	8.095948	7.684488
INF (-1)	4.108864	5.138098	4.809976
D (GDPG)	1.539773	1.849366	1.749243
Observations	25	54	79

Table A2: Dependent variable frequencies

Dep. value	Count	Percent	Cumulative count	Percent
0	25	31.00	25	31.65
1	54	68.00	79	100.00

Table A3: Estimates of firms' vulnerability to financing shocks

	Probability (%)*	
Firm [a]	**2011**	**2008**
MEPL	17.2	2.8
NECL	6.4	8.9
PCECPL	6.1	2.5
NEIPL	5.9	2.4
SEC	5.8	3.0
RIECPL	5.7	2.4
WSECPL	5.6	2.7
AIMEPL	5.5	2.9
PIECPL	5.4	2.9
RECL	5.3	2.8
AECPL	5.3	2.9
DEECPL	5.3	3.5
GECPL	5.3	2.7
PECPL	5.2	4.4
RECPL	5.2	1.8
FECPL	5.0	3.2
HCEPL	4.5	3.2
HHECPL	4.4	2.6
AECPL	4.3	2.7
HQIEP	4.2	2.1

Note: * = probability (%) of being vulnerable to financing shock. [a] Acronyms have been used for the firms.

Table A4: Expectation-prediction evaluation for binary specification, Success cutoff: C = 0.05

	Estimated equation			Constant probability		
	Dep. = 0	**Dep. = 1**	**Total**	**Dep. = 0**	**Dep. = 1**	**Total**
P (dep. = 1) <= C	2.00	0.00	2.00	0.00	0.00	0.00
P (dep. = 1) > C	23.00	54.00	77.00	25.00	54.00	79.00
Total	25.00	54.00	79.00	25.00	54.00	79.00
Correct	2.00	54.00	56.00	0.00	54.00	54.00
Correct (%)	8.00	100.00	70.89	0.00	100.00	68.35
Incorrect (%)	92.00	0.00	29.11	100.00	0.00	31.65
Total gain*	8.00	0.00	2.53			
Percent gain**	8.00	NA	8.00			
	Estimated equation			**Constant probability**		
	Dep. = 0	**Dep. = 1**	**Total**	**Dep. = 0**	**Dep. = 1**	**Total**
E (# of dep. = 0)	13.46	11.64	25.10	7.91	17.09	25.00
E (# of dep. = 1)	11.54	42.36	53.90	17.09	36.91	54.00
Total	25.00	54.00	79.00	25.00	54.00	79.00
Correct	13.46	42.36	55.82	7.91	36.91	44.82
Correct (%)	53.83	78.45	70.66	31.65	68.35	56.74
Incorrect (%)	46.17	21.55	29.34	68.35	31.65	43.26
Total gain*	22.19	10.09	13.92			
Percent gain**	32.46	31.89	32.17			

Note: * = change in "% correct" from default (constant probability) specification. ** = percent of incorrect (default) predictions corrected by equation.

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