**Does Inflation Bias Stabilize Real Growth? Evidence from Pakistan**

**Zafar Hayat** a,b,\*, **Faruk Balli** b,c, **Muhammad Rehman**d

a *Research Department, State Bank of Pakistan, Pakistan;* b *School of Economics and Finance, Massey University, New Zealand;* c *Department of International Trade and Marketing, Gediz University, Izmir, Turkey;* d *Monetary Policy Department, State Bank of Pakistan, Pakistan*

.....................................................................................................................................................................................

**Abstract**

The motive of a typical discretionary central banker to accommodate excess inflation (inflation bias) is either to stabilize real growth or to spur it beyond the natural rate. To what extent inflation bias helps to materialize this intention warrants empirical investigation. A more direct empirical probe into this issue, however, requires observable inflation bias indicators, which we model through desirable and threshold inflation rates, and society’s preferences for these. While examining the effects of inflation bias for the typical case of the discretionary monetary policy strategy of Pakistan, we found that contrary to the desired boost (stabilization) in real growth, inflation bias produced counterproductive results. Inflation bias was not merely ineffective in inducing real growth in the long term but significantly destabilized it. The higher the inflation bias, the higher the intensity of its destabilizing effect and vice versa. These results are robust to different inflation bias indicators and subsample analysis.

***JEL Codes:*** *E31, E52, E58.*

***Key Words:*** *Inflation bias indicators, Desirable and threshold inflation, Inflation bias–growth nexus, ARDL, Pakistan.*

\*Corresponding author’s email: zafarhayat78@gmail.com; Cell: +92 346 1957588.

**Acknowledgements:** We are grateful to Dr. James Obben, Dr. Shamim Shakur, Dr. Sajjad Zaheer, Dr. Jameel Ahmed, Dr. Zulfiqar Haider and Fayyaz Hussain for useful discussion on parts of the research. The views expressed in this research are those of the authors and do not represent the views of the State Bank of Pakistan in any way.

1. **Introduction**

There is wide agreement that vesting unconstrained discretion with a central bank for achieving the dual objectives of inflation and real growth results in excess inflation, an outcome commonly known as inflation bias in the literature. In theory, such a central banker tends to be flexible on the inflation objective, either to spur real growth beyond its potential (Kydland and Prescott, 1977; Barro and Gordon, 1983a) or to help it not falter (Cukierman, 2000). Since there is no universal definition of inflation bias, its use and interpretation in the literature have mostly remained contextual. The central theme, however, is the end product of inflation beyond some unknown but desirable level preferred by society. For example, Garman and Richards (1989) posited that inflation bias is the difference between observed inflation and society’s preferred inflation. Gartner (2000) viewed it as the tendency of central banks with representational preferences (preferences for employment and inflation) to generate inefficiently high inflation rates without gaining the benefit of output beyond its potential. Ruge-Murcia (2004) presented it as the systematic difference between equilibrium and optimal inflation, whereas Romer (2006) conceptualized it as the tendency of monetary policy to produce higher (than optimal) rates of inflation over extended periods.

Empirical research has established the evidence of inflation bias rather indirectly. Stylized models have been used to focus on a particular explanation of inflation bias rather than the outcome per se. For instance, Garman and Richard (1989) used voters’ preferences, Romer (1993) focused on the relationship between openness and inflation, Ireland (1999) examined the cointegrating relationship between inflation and unemployment, Cukierman and Gerlach (2003) estimated the relationship between output volatility and inflation, Ruge-Mercia (2004) explored the relationship between inflation and conditional variance of unemployment, and Berlemann (2005) used the symmetry in the employment–inflation trade-off. Recently, Hayat et al. (2016) quantified the discretionary behavior of a typical discretionary central banker, and the causal persistent behaviors in inflation and real growth variables, and found that discretion is significantly biased against inflation without stimulating any offsetting real growth gains.By and large, a common feature of the empirical work is the use of inflation as a proxy for inflation bias, hence trivializing the distinction between the two. This implicit assumption of the synonymy of inflation bias and inflation in empirical analysis is rather strong but seems to exist because of the absence of directly observable quantitative indicators of inflation bias as put forth by Surico (2008, p. 35), namely that “measuring and disentangling the inflation bias remains a challenging topic for future research.”

To bridge this gap, we model inflation bias in a way that allows its empirical approximation over time. We build our inflation bias models on the basis of desirable and threshold inflation rates, along with society’s preferences. We posit that the main problem in generating inflation bias indicators not only hinges on the identification and estimation of society’s desired as well as maximum acceptable levels of inflation rates, but also on society’s preferences.

Therefore, building on the notions of desirability and the acceptability of certain inflation rates by society and their preferences, we develop a framework for generating numerical time series indicators of inflation bias. In turn, these inflation bias indicators facilitate a direct empirical investigation to assess whether inflation bias significantly help to boost or stabilize real growth. If it does, then the central bank has achieved its aims; otherwise, a rethink of the discretionary monetary policy may be warranted. Since inflation bias is the outcome of typical cases of discretion—defined as a central bank with the dual objectives of inflation and growth, and attempting to attain a higher than potential level of the latter. This no longer seems to be relevant in advanced countries like the USA (Blinder, 1998); the phenomenon, however, may commonly be found in developing countries like Pakistan.[[1]](#footnote-1) For example, Pakistan’s central bank explicitly targets the dual objectives of inflation and real growth, and, in general, the growth targets are set beyond the natural rate (Figure 1).[[2]](#footnote-2)

As the inflation bias theory assumes long-term, the parameters of the ‘inflation bias’ indicators have been obtained through the autoregressive distributed lag (*ARDL*) bounds testing and estimation approach. The main result of the paper shows that instead of being ineffective or exerting a stabilizing effect, ‘inflation bias’ significantly destabilizes real growth. This result is robust across the generated ‘inflation bias’ indicators and across samples. The results also indicate that the higher the degree of inflation bias, the more intense and adverse its effects on the real growth. These findings are particularly illuminating when viewed in the context of developing countries, as these nations normally tend to be more biased towards inflation to acheive a relatively higher real growth.

The remainder of the paper is structured as follows. Section 2 outlines the methodological framework and Section 3 analyzes the stationarity properties of the variables. Section 4 presents and discusses the results, and conducts the robustness analysis while Section 5 concludes the paper.

1. **Methodological framework**

In this section, we discuss the development and implementation of our methodological framework in three steps. In the first step, we uniquely model inflation bias on the basis of desirable and threshold inflation rates, and their respective preference parameters. In the second step, we rationalize and elicit the distinction between desirable and threshold inflation rates and society’s preferences and develop a framework for estimating these preference parameters. Lastly, we model the final generated inflation bias indicators to explore its long-term trade-off with real growth.

* 1. *Framework for generation of inflation bias indicators*

As highlighted earlier, inflation bias is not universally defined and no guiding criterion is available that can be used to generate its time series numerical indicators. Nevertheless, for working purposes, consistent with the essence of the notion, we define the $i$th inflation bias indicator $IB\_{t}^{i}$ as follows:

$IB\_{t}^{i}=\left\{\begin{array}{c}IB\_{t}^{d\_{i}}= \left(π\_{t}-π^{d\_{i}}\right)\*ω^{d\_{i}} for i=1, …., n-1 \\IB\_{t}^{th}= \left(π\_{t}-π^{th}\right)\*ω^{th} for i= n \end{array}\right\} $, *(1)*

where $π\_{t}$ is the observed inflation rate over time $t$, and we suppose that there are $n-1$ possible $i$th ‘desirable’ inflation rates $π^{d\_{i}}$ and one ‘threshold’ inflation rate $π^{th}$ such that $π^{d\_{1}}< π^{d\_{2}}< π^{d\_{3}}….<π^{d\_{n-1}}< π^{th}.$ Since society tends to prefer low inflation rates, the preference of society thus would tend to vary (decrease) as the inflation increases from low levels to high levels (i.e., it would tend to prefer $π^{d\_{1}}$ over $π^{d\_{2}}$ over...,$ π^{th}$).[[3]](#footnote-3) The terms $\left(π\_{t}-π^{d\_{i}}\right)$ and $\left(π\_{t}-π^{th}\right)$ reflect the extent of the departure of observed inflation over time from the desirable and threshold inflation rates respectively, and $ω^{d\_{i}}$ and $ω^{th}$ represent the levels of society’s preference for departures from the desirable and threshold inflation rates respectively.

On average, as the observed inflation departs from low levels, the less it is preferred by society until it reaches the threshold inflation rate (see step 2 for the distinction between desirable and threshold inflation rates). For example, $\left(π\_{t}-π^{d\_{1}}\right)$ would be more preferred by the society than $\left(π\_{t}-π^{d\_{2}}\right)$ and so forth until $\left(π\_{t}-π^{th}\right)$, because the average inflation over time exceeding $π^{d\_{1}}$ would tend to be less than the average inflation exceeding $π^{d\_{2}}$ and so forth until $π^{th}$. Thus in terms of the preference parameter, society’s preferences for different inflation rates may better be expressed as $ω^{d\_{1}}$ > $ω^{d\_{2}}$ > $ω^{d\_{n-1}}$ > $ω^{th}$.

In terms of inflation bias as is defined in Equation 1, this would imply that the more the observed inflation departs from its most desirable level, the more the central bank is biased against inflation, and therefore the least it is liked by the society. For example, in our case, within the desirable levels of inflation $π^{d\_{i}}$, which are estimated/identified using Equation 2a in supplementary material, turns out to be 1% to 3% (Appendix 1). Since the society’s preference $ω^{d\_{i}}$ corresponding to each rate of inflation differs i.e. in this case 1% inflation is preferred as compared to 3%, as is indicated by the respective coefficients on $ω^{d\_{i}}$, the generated $IB$’s also differ accordingly. Conceptually, the distinction among the three desirable inflation rates from 1% to 3% is that the former is preferred over the latter by the society. Thus the more the observed inflation departs from the most desirable level—which in this case is 1%—the more the central bank is biased against inflation and the least it is preferred by the society. Precisely, $IB$3 indicates more bias on part of the central bank relative to $IB$1. Similarly, the fourth inflation bias indicator $IB$4 generated on the basis of $π^{th}$—which is 5% with the least preference value $ω^{th}$ (Appendix 1)—is least preferred as compared to the $IB$1, $IB$2 and $IB$3 and reflects the maximum terminal level of the bias on the part of the central bank that the society may not desire but may accept as per the case we have developed.

A negative omega implies that as the inflation shifts from one level to another, the effect on the real growth is negative (adverse). In our case this happens for the inflation rates exceeding 5% level i.e. ≥6%–26%, where 26% is the maximum observed inflation rate in Pakistan’s economic history (Appendix 1). The negative omegas thus essentially imply that the corresponding inflation rates are undesirable and therefore not accepted to the society for two main reasons. First, at these levels the corresponding inflation rates are relatively high and therefore not liked by the society and second, these inflation rates hinders the real growth, which again is a characteristic not liked by the society. Thus negative omegas cause twofold harms to the society. Since by definition (Equation 1), any inflation rate in excess to a certain desirable $π^{d\_{i}}$ or threshold level $π^{th}$ level is inflation bias, negative omegas may not be used here for the purpose because it does not befit the very definition. Precisely such inflation rates are neither preferred nor accepted by the society. Therefore, for inflation ≥6% up to 26%, which are consistently negatively related to the real growth, the idea of desirability/acceptability of inflation vanishes.

Technically, our weighing scheme for the difference of desirable and threshold inflation rates from observed inflation is important for more than the obvious logical reasons because a simple (unweighted) difference may potentially pose problems. For example, an unweighted difference would be rather mechanical, which would render individual regression estimates for different inflation bias indicators $IB\_{t}^{i}$ meaningless. In such a case, the differences between different inflation bias indicators when regressed on the dependent variable would only be captured by the intercept term instead of the parameter estimates.

Further, it is also important to highlight that just in case in a particular point in time $t$, $\left(π\_{t}-π^{d\_{i}}\right)$ or $\left(π\_{t}-π^{th}\right)$ < 0, a conceptual problem would arise as, by definition, each value of the inflation bias indicator(s) at any point in time $t$ should be greater than 0 (i.e. $IB\_{t}^{i}\geq 0 $), as a negative value for any inflation bias indicator at a certain point in time $t$ (i.e. $IB\_{t}^{i}<0 $) would instead mean a deflation bias at that point in time. Again, at any point in time $t$, $IB\_{t}^{i}=0$ would imply the absence of inflation bias. Therefore, for working purposes and in order to be consistent in our definitions, in the case of $IB\_{t}^{i}<0$, negative values, if any, need to be restricted to ‘0’, hence we assume absence of inflation bias in that particular period $t$.[[4]](#footnote-4)

1. *Rationalization and estimation of desirable and threshold inflation rates, and their respective preference parameters*

Before rationalizing and estimating society’s preference parameters $ω^{d\_{i}}$ and $ω^{th}$, it is imperative to make a clear distinction between the desirable $π^{d\_{i}}$ and threshold $π^{th}$ inflation rates to be able to quantify Equation 1. By the threshold inflation rate, in the context of a long-term inflation–growth nexus, we mean a level of inflation that reflects a state beyond which the effects of inflation on real growth become negative (see Khan and Senhadji, 2001). While being consistent with Garman and Richards (1989), desirability, on the other hand, means that from society’s point of view, any change in inflation may be ‘desirable’ if it leads the economy towards an optimum (i.e. it significantly enhances growth).

In contrast to the nonexistence of empirical literature on estimating desirable inflation rates, a large body of empirical literature has explored the threshold effects of inflation on growth (nonlinear relationship) both for panels of advanced and developing countries, and for individual countries (Sarel, 1996; Khan and Senhadji, 2001; Drukker et al. 2005; Burdekin et al. 2004; Kannan and Joshi, 1998; Mohanty et al. 2011; Nawaz and Iqbal, 2010; Munir and Mansur, 2009; Mubarik, 2005; Hayat and Kalirajan, 2009). It is, however, important to note that largely their focus is to establish the point of inflexion in the data, whereas ours is to estimate and identify the ‘good’ part of inflation ($π^{d\_{i}}$) and the maximum acceptable rate of inflation $(π^{th}$) embedded in the range of inflation historically experienced by a particular country—in this case, Pakistan—to be used as benchmarks to identify and disentangle different levels of excess (bad) inflation.

To further clarify the distinction between desirable and threshold inflation rates, considering a hypothetical example in the context of the inflation–growth nexus, we may argue that if there is only one threshold that occurs (say, at 7% inflation), it may be expected that inflation rates from 1% to 7% would have a positive impact on real growth. Although all such inflation rates may be positively related to real growth, their statistical significance, however, may vary: some of them may be statistically significant and others may not. All the statistically significant inflation rates below the ‘threshold’ level may be deemed ‘desirable’ because they significantly induce the economy to grow, which, in turn, may roughly approximate an improvement in the wellbeing of society. The threshold inflation may not be the best choice but may represent the cut-off preference point from society’s perspective, as beyond this rate, inflation negatively affects real growth. Inflation occurring beyond the threshold level would instead plunge society into an economic state where it loses both in terms of facing relatively higher average inflation rates and as a result of losing real growth that would have been otherwise attainable.

As is clear both from theory and empirical studies (e.g. Barro and Gordon, 1983a,b; Motley, 1998; Dotsey, 2008), a low and stable inflation rate is more likely to be relatively more growth-enhancing. We expect the associated parameters for low inflation rates to be positive. Logically, we also expect a decrease in the magnitude of the positive effects of inflation on real growth as inflation goes up until it reaches the threshold level. In order to be consistent with the aforementioned literature on the nonlinear relationship (threshold effects) between inflation and growth, any inflation beyond the threshold is expected to not only negatively affect real growth but its magnitude would tend to increase with an increase in inflation from one level to another.

If this phenomenon is empirically true, any parameter that may possibly capture the positive effect of the shifts in inflation on real growth from one level to another may allow us to track (approximate) the level of preference of society. In our Equation 1, $ω^{d\_{i}}$ and $ω^{th}$ essentially track the positive variations and their magnitudes in terms of real growth caused by the shifts in inflation rates (say, by one unit) from one level to another. Thus for any positive inflation rate below $π^{thr}$, the larger the magnitude of the preference parameter, the more it is preferred by society because we have the prior intuition that society should assign more weight to low inflation and a smaller weight to a relatively higher inflation. It is, however, pertinent to mention that in the case where inflation falls into a range higher than the threshold (i.e. $π$ > $π^{thr}$), this would be rejected rather than preferred by society, as it would not only be higher but would also deteriorate real growth, hence causing twofold harm to society. Thus, by definition, for any level of inflation exceeding the threshold, the preference issue becomes irrelevant because such inflation rates are not preferred (accepted) by society at all i.e. from the society’s perspective, it has crossed any rationally definable/justifiable preference limits.

In order to obtain the potential preference parameters of interest ($ω^{d\_{i}}$ and $ω^{th}$), we construct the term $DUM\_{t}^{s}$ to be regressed on real growth for different inflation rates ranging from 1% to 26%.[[5]](#footnote-5) The term $DUM\_{t}^{s}$ is defined as:

$$DUM\_{t}^{s}=\left\{\begin{array}{c}\left(π\_{t}-π^{\*\_{s}}\right) when π\_{t}\geq π^{\*\_{s}} \\0 otherwise\end{array}\right.$$

$π^{\*\_{s}}= \frac{s}{100}, s=1,….,26$.

In order to obtain $ω^{d\_{i}}$ and $ω^{th}$ from 26 different versions of $DUM\_{t}^{s}$, consistent with a range of popular growth studies (e.g. Barro, 1990, 1991, 1995; Romer, 1989, 1990; Barro and Sala-i-Martin, 1992, 1995; Levine and Renelt, 1992; Sarel, 1996; Khan and Senhadji, 2001), we specified a dynamic *ARDL* real growth model.[[6]](#footnote-6) The error correction version of the *ARDL* baseline real growth model given as:

$∆\hat{GDP}\_{t}=α\_{0}^{s} +\sum\_{f=1}^{p\_{1,s}}α\_{f}^{GDP,s}∆\hat{GDP}\_{t-f}+\sum\_{j=0}^{p\_{2,s}}α\_{j}^{π,s}∆π\_{t-j}+\sum\_{k=0}^{p\_{3,s}}α\_{k}^{POP,s}∆ \hat{POP}\_{t-k}+\sum\_{l=0}^{p\_{4,s}}α\_{l}^{INV,s}∆\hat{ INV}\_{t-l} +\sum\_{m=0}^{p\_{5,s}}α\_{m}^{FDI,s}∆\left(\frac{FDI\_{t-m}}{GDP\_{t-m}}\right)+\sum\_{n=0}^{p\_{6,s}}α\_{n}^{DUM^{s}}∆DUM\_{t-n}^{s}+β\_{0}^{s}\hat{GDP}\_{t-1}+β\_{1}^{s}π\_{t-1}+β\_{2}^{s}\hat{POP}\_{t-1}+β\_{3}^{s}\hat{INV}\_{t-1}+β\_{4}^{s}\left(\frac{FDI\_{t-1}}{GDP\_{t-1}}\right)+β\_{5}^{s}DUM\_{t-1}^{s}+ϵ\_{t}^{s} for s=1,….,26, $(2)

where $\hat{GDP}\_{t}$, $\hat{ INV\_{t}}$ and $\hat{ POP\_{t}}$ represent the growth rates of real gross domestic product (*GDP*), gross fixed capital formation (*INV*) and population (*POP*) at time $t$ respectively. The variable $\frac{FDI\_{t}}{GDP\_{t}}$ is the ratio of foreign direct investment (*FDI*) to the real *GDP* and $ϵ\_{t}$ is the Gaussian white noise process. It may be noted that if a long-term relationship exists, then the long-term relationship is normalized and expressed as follows:

$\hat{GDP}\_{t}=ϕ\_{1}^{s}π\_{t}+ϕ\_{2}^{s}\hat{POP}\_{t}+ϕ\_{3}^{s}\hat{INV}\_{t}+ϕ\_{4}^{s}\left(\frac{FDI\_{t-1}}{GDP\_{t-1}}\right)+ϕ\_{5}^{s}DUM\_{t}^{s}+η\_{t}^{s}$, (2a)

where, $ϕ\_{z}^{s}=-\frac{β\_{z}^{s}}{β\_{0}^{s}}$ , for $z$ = 1,2…,5.

Here, it is important to mention that although research has identified a range of growth determinants (Levine and Renelt, 1991 provides a summary of such variables), not all of them have been found to be robust, except investment (see Levine and Renelt, 1992), which is there in our model in addition to inflation, population and *FDI*. All the four variables in our model may reasonably account for the monetary sector, real sector, labor force and international capital sectors, and the use of *ARDL* allows us to account for optimal dynamics of the dependent and independent variables. Furthermore, we tried numerous other potentially important variables (for which data were available) for possible inclusion in our baseline real growth model such as indicators of human capital (primary and secondary school enrolments); ratios of government debt, exports, imports and *M2* to *GDP*; openness; exchange rate and trade balance. However, we dropped these variables subsequently because either (i) they were insignificant, (ii) they were non-cointegrated or (iii) the estimated models (while retaining these indicators) could not pass either of the key diagnostic tests for normality, serial correlation, functional form and heteroscedasticity, or stability tests such as the cumulative sum of squares of residuals (*CUSUM*) and cumulative sum of squares of recursive residuals (*CUSUMQ*).

The coefficients $ω^{d\_{i}}$ and $ω^{th}$ on the term $DUM\_{t}^{s}$ obtained through *Equation 2a* capture the difference in the effects of inflation on real growth when it shifts from one level to another; the extent of its significance may be determined by the respective *P*-values. Since $ω^{d\_{i}}$ and $ω^{th}$ indicate how different inflation levels exceeding a particular rate in terms of their effect on real growth are, they may reasonably serve as preference parameters. Since their magnitudes and signs would tend to vary for different arbitrary inflation rates from 1% to 26%, they allow us to account for the differences in the growth effects as inflation goes up because this is how $DUM\_{t}^{s}$ has been defined and constructed.

*2.3 Inflation bias and real growth nexus*

To reiterate, the ultimate goal of generating the inflation bias indicators $IB\_{t}^{i}$ is to explore whether they have a long-term positive cointegrating relationship with real growth to possibly justify the existence of inflation bias. To this end, we first generated the time series of $IB\_{t}^{i}$ through Equation 1 for different values of $π^{d\_{i}}$ and $π^{th}$. The respective $ω^{d\_{i}}$ and $ω^{th}$ values are obtained from Equation 2 by using $DUM\_{t}^{s}$ for inflation values from 1% to 26%.

Only inflation in the range from 1% to 3% turns out to be desirable, as inflation in this range exerts a positive and statistically significant effect on real growth. Therefore, three inflation bias indicators were generated on this basis from Equation 1. The fourth inflation bias indicator was generated for $π^{th}$=5% because 5% turned out to be threshold rate of inflation; in other words, beyond this inflation rate, the effects of inflation on real growth became negative (see Appendix 1 for these results, which are discussed in detail in the next section). These indicators were then introduced into the real growth model now specified as Equation 3 by substituting $π\_{t}$ one by one with each inflation bias indicator to obtain their long-term parameter estimates. The error correction version of the inflation bias–real growth *ARDL* model thus takes the form:

$$∆\hat{GDP}\_{t}=γ\_{0}^{i} +\sum\_{f=1}^{p\_{1,i}}γ\_{f}^{GDP,i}∆\hat{GDP}\_{t-f}+\sum\_{j=0}^{p\_{2,i}}γ\_{j}^{IB,i}∆IB\_{t-j}^{i}+ \sum\_{k=0}^{p\_{3,i}}γ\_{k}^{POP,i}∆\hat{ POP}\_{t-k} +\sum\_{l=0}^{p\_{4,i}}γ\_{l}^{INV,i}∆\hat{ INV}\_{t-l}+\sum\_{0}^{p\_{5,i}}γ\_{m}^{FDI,i}∆\left(\frac{FDI\_{t-m}}{GDP\_{t-m}}\right)+φ\_{0}^{i}\hat{GDP}\_{t-1}+φ\_{1}^{i}IB\_{t}^{i}\_{t-1}+φ\_{2}^{i}\hat{POP}\_{t-1}+φ\_{3}^{i}\hat{INV}\_{t-1}+φ\_{4}^{i}\left(\frac{FDI\_{t-1}}{GDP\_{t-1}}\right)+ ϑ\_{t}^{i}, for i=1, …., n. (3)$$

$$ $$

Four distinct versions of this model were estimated separately using all the four $IB\_{t}^{i}$ in turn; the long-term coefficients were obtained using a method similar to that of Equation 2a (see Section 4 for the results and discussion).

1. **Data sources and stationarity properties**

The data were obtained from the World Bank Development Indicators, covering a period from 1961 to 2010.[[7]](#footnote-7) Since our interest is in obtaining the long-term nonspurious coefficients while taking the dynamics into account, the stationarity properties of the variables were examined to reinforce whether we had selected the right estimation approach. As reported in Table 1, the stationarity tests of the variables show that the regressors are both a mixture of I(0) and I(1) orders of integration. The use of the *ARDL* testing and estimation approach of Pesaran *et al.* (2001) therefore seems the right choice, which is designed to circumvent the underlying issue of different orders of integration in the variables.

|  |
| --- |
| ***Table 1: Stationarity properties of the variables*** |
|   | *ADF* | *PP* |
| Variables | Level | First difference | Level | First difference |
| $$π$$ | [0.16] | [0.00]\*\*\* | [0.26] | [0.00]\*\*\* |
|

|  |
| --- |
| *IB1* |

 | [0.13] | [0.00]\*\*\* | [0.18] | [0.00]\*\*\* |
| *IB2* | [0.09]\* | [0.00]\*\*\* | [0.13] | [0.00]\*\*\* |
| *IB3* | [0.06]\* | [0.00]\*\*\* | [0.07]\* | [0.00]\*\*\* |
| *IB4* | [0.02]\*\* | [0.00]\*\*\* | [0.02]\*\* | [0.00]\*\*\* |
| $$\hat{POP}$$ | [0.33] |  [0.07]\* | [0.45] | [0.03]\*\* |
| $$\hat{GDP}$$ | [0.24] | [0.00]\*\*\* | [0.05]\* | [0.00]\*\*\* |
| $$\hat{INV}$$ | [0.00]\*\*\* |  | [0.00]\*\*\* |  |
|  $FDI/GDP$ | [1.00] | [0.00]\*\*\* | [0.77] | [0.05]\* |
| This table reports the *P-v*alues of the Augmented Dicky–Fuller (*ADF*) and the Phillips–Perron (*PP*) tests in brackets. \*\*\*, \*\* and \* indicate that the series are stationary at the 1%, 5% and 10% level of significance, respectively. |

1. **Results**
	1. *Baseline growth model: cointegration, diagnostics, stability and robustness*

Because in practice, the ‘true’ orders of the *ARDL* ($p,q$) model are rarely known *a priori*, we selected our baseline real growth model by using the *SBC* and imposing a maximum lag of 3 years to allow a sufficient transmission time. This is a relatively consistent model selection criterion in small samples that leads to the selection of the most parsimonious model with the lowest number of freely estimated parameters (Enders, 1995; Pesaran and Pesaran, 1997). The model is selected from thousands of models are estimated (iterated) in the background.

Since it is not possible to determine *a priori* if $∆π$, $∆\hat{ POP}$, $∆\hat{ INV}$ and $∆\frac{FDI}{GDP}$ are the ‘long-run forcing’ variables for real $\hat{GDP}$ growth, as suggested by the third assumption of Pesaran *et al.* (2001), we tested for this by excluding the current values of the explanatory variables while setting the null hypothesis of the nonexistence of a long-run relationship as $H\_{0}:β\_{1}= β\_{2} = β\_{3} = β\_{4} =0$ against $H\_{1}:β\_{1}\ne β\_{2} \ne β\_{3} \ne β\_{4}\ne 0$.

As the computed *F*-statistics (7.34) exceed the upper bound of the critical value band (4.68) at the 1% level, we cannot reject the null hypothesis of no long-run relationship among $\hat{GDP,} π$,$ \hat{ POP}$,$ \hat{ INV}$ and $ \frac{FDI}{GDP}$. The long-term relationship exists at the 1% level even if we use the upper critical value bound (5.87) computed by Narayan (2005) for small sample sizes. The model passes the key diagnostic tests for autocorrelation, specification, normality and heteroskedasticity, as well as the stability tests (see Appendices 1 and 2).[[8]](#footnote-8)

In order to control for the effects of the well-known inflation shocks (1973, 1974 and 1975) in the aftermath of Pakistan’s war with India in 1971 and the impact of international oil shocks in 1973, a dummy variable was introduced into the model but it was dropped due to its insignificance. The *P*-values of the joint test of zero restrictions on the coefficients of the deleted variables were 0.74, 0.74 and 0.77 for the Lagrange multiplier, likelihood ratio and *F*-statistics, respectively. As a robustness check on the baseline model, we bifurcated the sample and re-estimated the results, which were found to be highly consistent with the results of the main baseline growth model.[[9]](#footnote-9)

The results further build our confidence that the baseline growth model can be used for estimating $DUM\_{t}^{s}$ as the estimated long-term coefficients exhibit the correct signs and significance (see Table 2). The effects of inflation on real growth are significantly adverse, a result that is consistent with the established viewpoint and empirical evidence (e.g. Kydland and Prescott, 1977; Barro and Gordon, 1983a; De Gregario, 1992, 1993; Barro, 1995; Wilson, 2006). Investment, on the other hand, significantly enhances real growth; again, a result that is well established in the literature (see Levine and Renelt, 1992 for the robustness of the investment–growth relationship). The long-run effects of population and *FDI* on real growth, however, did not show statistical significance but yielded the correct signs. Despite their statistical insignificance, these variables were retained in the model, as their deletion is not supported by the joint test of zero restrictions on the coefficients of the deleted variables. For example, the respective *P*-values of the Lagrange multiplier, the likelihood ratio and the *F*-tests for the joint deletion of population and *FDI* are 0.02, 0.01 and 0.03, respectively.

*4.2 Desirable and threshold inflation rates and preference parameters*

As the generation of the series of inflation bias indicators from Equation 1 requires the identification and estimation of desirable and threshold rates of inflation and the respective preference parameters of society, $DUM\_{t}^{s}$ has to be tested via the baseline real growth model (Equation 2) for arbitrary inflation rates from 1% to 26% to explore the possible inflation rates that may potentially enhance growth. Essentially, 26 versions of Equation 2were estimated using 26 different specifications of $DUM\_{t}^{s}$. The results of the baseline growth model and its 26 variations using $DUM\_{t}^{s}$, their respective cointegration and the diagnostic tests are reported in Appendix 1. The usual interpretation (e.g. for Model 1 in Appendix 1) would be that inflation is negatively associated with real growth (–4.63) but less so if inflation exceeds the 1% level (4.46). Overall, the effect of inflation exceeding the 1% level remains negative [–4.63 + 4.46 = –0.17].

The coefficients of $DUM\_{t}^{s}$ for different rates of inflation from 1% to 26% paint an interesting picture (Figure 2). The threshold occurs at the 5% level of inflation because when inflation shifts from 5% to 6% and above, its effects on real growth becomes negative, a result that is consistent with the literature on nonlinearity in the effects of inflation on real growth (e.g. Fischer, 1993; Sarel, 1996; and Khan and Senhadji, 2001). This 5% inflation may be deemed as the maximum acceptable/tolerable level of inflation from society’s perspective because inflation at this level, although relatively high, is positively related to real growth, hence at least dampens the negative effect of inflation on real growth. Since any inflation rate beyond the 5% level amplifies the implied negative effect on real growth with an increasing magnitude, inflation in the range from 6% to 26% is therefore unacceptable by society. For example, the implied negative effect for inflation exceeding 6% is [–0.19 + (-0.05) = –0.24] and that for inflation exceeding, say, 21% is [–0.18 + (-0.53) = –2.39].

Consistent with our prior intuition, although the coefficients on $DUM\_{t}^{s}$ from 1% to 5% are positive, they are statistically significant only for inflation rates from 1% to 3% (see the respective coefficients and their statistical significance on $DUM\_{t}^{s}$ in Appendix 1, Model 1 to Model 5). This indicates that inflation in 1–3% range significantly boosts real growth. This boost occurs with a decreasing magnitude as the inflation rises from 1% to 3%. Essentially, as far as the preference of society is concerned, we should prefer $\left(π\_{t}-π\_{1}\right)$ over $\left(π\_{t}-π\_{5}\right)$ because the former has more of the ‘good’ part embedded within it than the latter term (i.e. low inflation and a statistically significant positive effect on real growth), and is quantitatively larger than the latter term.

For example, the coefficient on the respective $DUM\_{t}^{s}$ of the term $\left(π\_{t}-π\_{1}\right)$ is not only higher (4.46) but is also statistically significant compared to 0.23 on the respective term with a 5% threshold (see Appendix 1, Model 1 and Model 5). Using this logical reasoning, we therefore weigh the term $\left(π\_{t}-π\_{t}^{i}\right)$ in Equation 1 by the respective coefficients on $DUM^{d\_{i}}$ and $DUM^{th}$ as society’s preference parameters accordingly to generate various indicators of inflation bias from Equation 1 using 1% to 3% as desirable, and 5% as the threshold or maximum acceptable level.

* 1. *The nexus between inflation bias and real growth*

 In order to test for the possible existence of a cointegrating relationship, the null and alternative hypotheses were defined as $H\_{0}: φ\_{1}= φ\_{2} = φ\_{3} = φ\_{4}=φ\_{5}=0$ against $H\_{1}:φ\_{1}\ne φ\_{2} \ne φ\_{3} \ne φ\_{4}\ne φ\_{5}\ne 0$ for all four versions of Equation 3 for the four inflation bias indicators separately. The *F*- statistic for all the models was greater than the asymptotic critical bound values of Pesaran and Pesaran (2009) and Narayan (2005) at the 1% level (see Table 2). This indicates the existence of a cointegrating relationship and thus the long-term coefficients were obtained subsequently.

|  |
| --- |
| ***Table 2: ARDL Bounds Test Results***  |
|  |  |  |  | Lag Order |  | Pesaran and Pesaran (2009)\*  |  | Narayan (2005)\*  |  | Cointegration outcome |
| *Model* | *Computed F-statistics* |   | *SBC criterion* |  | *Lower bound at 1%*  | *Upper bound at 1%* |  | *Lower bound at 1%*  | *Upper bound at 1%* |  | *F-Statistics > CV bounds at:* |
| *Model 1* | 7.42 |  | 0,2,1,0,1 |  | 3.52 | 4.78 |  | 3.95 | 5.58 |  | 1% |
| *Model 2* | 7.41 |  | 0,2,1,0,1 |  | 3.52 | 4.78 |  | 3.95 | 5.58 |  | 1% |
| *Model 3* | 8.39 |  | 0,0,0,0,0 |  | 3.52 | 4.78 |  | 3.95 | 5.58 |  | 1% |
| *Model 4* | 8.29 |  | 0,0,0,0,0 |  | 3.52 | 4.78 |  | 3.95 | 5.58 |  | 1% |
| \* *Critical value bounds at k=5 with an unrestricted intercept and no trend.* |

All the models are stable when tested by the *CUSUM* and the *CUSUMQ* (Appendix 3). The *P*-values of the diagnostic tests are presented along with the main results in Table 3. The first two models pass all the key diagnostic tests as the null hypotheses of no serial correlation, no misspecification, normality of residuals and homoscedasticity could not be rejected. It may be noted that Model 3 and Model 4, which contain *IB3* and *IB4*, do not pass the specification test and their $R$2 is significantly lower than the models with *IB1* and *IB2*.

The estimated long-term coefficients of all the inflation bias indicators show that inflation bias is significantly detrimental to real growth (Table 3). Although this result supports the influential theoretical contributions of Kydland and Prescott (1977) and Barro and Gordon (1983a,b), which stated that inflation bias does not stimulate higher than potential growth in the long-run, it also reveals that rather than being ineffective, inflation bias essentially destabilizes real growth. Out of the set of four inflation bias indicators, *IB1* and *IB2* provide a better explanation in terms of the fit of the data compared to *IB3* and *IB4* in the full sample.[[10]](#footnote-10) Furthermore, both *IB1* and *IB2* are significantly detrimental to real growth at the 1% level, whereas *IB3* and *IB4* remain insignificant.

Ignoring significance, the size of the coefficients of the inflation bias indicators largely suggest that the lower the inflation bias, the less it deteriorates real growth, whereas the higher the inflation bias, the more it is detrimental to real growth. Due to the nonexistence of empirical research on the same lines in developing countries like Pakistan, we cannot refer to the findings of other studies to directly support ours; however these results are highly consistent with theory and the monetary policy practices of advanced countries.[[11]](#footnote-11) For example, by and large, the advanced countries’ central banks set their inflation targets around 2% (Romer and Romer, 2002).[[12]](#footnote-12) This rate is consistent not only with ‘price stability’ but also allows a sufficient cushion to trivialize zero lower bound in a world of small shocks (Blanchard *et al.* 2010). Surico (2008) estimated a bias of 1% in the case of the USA for the pre-1979 policy regime and observed that inflation bias disappears when the inflation target is close to 2%.

The findings of this paper seem to have nontrivial implications for Pakistan’s monetary policy as, historically, the average inflation bias has critically been high at 8.87%, which is roughly eight times higher than the 1% level in the USA in the pre-1979 period.[[13]](#footnote-13) Therefore, for the central bank of Pakistan, accepting relatively higher inflation rates with the hope of achieving higher real growth rates has not been an effective strategy. Being accommodative in terms of excess inflation has resulted in a significant loss in otherwise attainable real growth. Both the outcomes of excess inflation and lost real growth are indicative of the policy’s nonperformance when gauged against the twofold statutory objectives of the country’s monetary policy. Society will not appreciate a policy of this kind, which, instead of improving their welfare by ensuring low and stable inflation and a steady sustainable economic growth, worsens their situation.

The monetary authorities in Pakistan therefore may need to reorient and design their approach towards how they conduct monetary policy, consistent with the best monetary policy practices, by diverting their focus towards attaining low and stable inflation, preferably around the 2% level. Although this is demanding and challenging, focusing on attaining these inflation levels, on one hand, would mitigate the losses from higher inflation and, on the other hand, may pave the way for reaping the fruits of price stability, namely low and stable inflation and sustainable economic growth. Here, it is important to mention that price stability may be deemed as one of the important factors largely under the control of monetary authorities, which will have to be supplemented by other factors that determine real growth beyond its jurisdiction.

Price stability is believed to be beneficial for sustainable economic growth. Keeping the inflation rate below or close to 2% would allow the central bank to contain the volatility of inflation. The costs of inflation volatility on the economy are diverse and therefore should be taken into account by Pakistan’s monetary policy authorities in their decision making process. Friedman (1977) conjectured that the primary cause of business cycles is unpredictable inflation (see also Engle, 2003). Unpredictability in turn breeds uncertainty, which affects the investors’ behavior and hence lowers investment, economic growth and employment generation. Hossain (2014) noted that inflation volatility adversely affects real growth via its spillover effects on interest rates and exchange rates. All these channels might have been effective in contributing to the adverse effects of excess inflation on real growth. Since inflation bias has implications for the monetary policy credibility of Pakistan, a focus on price stability may help them improve the effectiveness of monetary policy in containing inflation bias, which, in turn, may acheive real growth gains in future.

In other words, this would imply that the central bank of Pakistan will have to pursue its objective of achieving sustainable economic growth by ensuring low and stable inflation rather than actively pursuing higher than potential real growth rates of the economy, which are not possible because of the time inconsistency problem associated with this monetary policy.

|  |
| --- |
| ***Table 3: Long-term parameter estimates of inflation bias indicators (1961–2010)*** |
| Models /Variables | Variables | Fit of the models and the diagnostic tests |
| $$IB1$$ | $$IB2$$ | $$IB3$$ | $$IB4$$ | $$\hat{ POP}$$ | $$\hat{ INV}$$ | $$\frac{FDI}{GDP}$$ | $$γ$$ | $$R^{2}$$ | AUTO | SPEC | NORM | HETR |
|  | -0.05\*\* |  |  |  | 0.94 | 0.16\*\*\* | 23.75 | 4.02\* |  |  |  |  |  |
| Model 1 ($IB1$) | (0.02) |  |  |  | (0.74) | (0.04) | (27.10) | (2.11) | 0.46 | [0.93] | [0.20] | [0.15] | [0.85] |
|  | [0.01] |  |  |  | [0.21] | [0.00] | [0.38] | [0.06] |  |  |  |  |  |
|  |  | -0.12\*\* |  |  | 0.92 | 0.16\*\*\* | 23.60 | 3.84\* |  |  |  |  |  |
| Model 2 ($IB2$) |  | (0.04) |  |  | (0.74) | (0.04) | (27.20) | (2.11) | 0.45 | [0.92] | [0.18] | [0.16] | [0.87] |
|  |  | [0.01] |  |  | [0.22] | [0.00] | [0.39] | [0.07] |  |  |  |  |  |
|  |  |  | -0.02 |  | 0.62 | 0.13\*\*\* | -12.12 | 3.44 |  |  |  |  |  |
| Model 3 ($IB3$) |  |  | (0.05) |  | (0.77) | (0.04) | (27.10) | (2.21) | 0.24 | [0.40] | [0.01] | [0.68] | [0.67] |
|  |  |  | [0.64] |  | [0.42] | [0.00] | [0.65] | [0.12] |  |  |  |  |  |
|  |  |  |  | -0.15 | 0.61 | 0.13\*\*\* | 12.09 | 3.43 |  |  |  |  |  |
| Model 4 ($IB4$) |  |  |  | (0.30) | (0.77) | (0.04) | (26.84) | (2.21) | 0.24 | [0.40] | [0.01] | [0.68] | [0.69] |
|  |   |   |   | [0.61] | [0.43] | [0.00] | [0.65] | [0.12] |   |   |   |   |   |
| This table reports the cointegrating relationship of the real GDP and the inflation bias indicators. The *P*-values of the diagnostic tests are presented sequentially with AUTO denoting the Langrange multiplier test for autocorrelation. SPEC represents a general test for omitted variables and a functional form test (the Ramsey regression equation specification error test (RESET)) using the square of the fitted values. NORM indicates the test for normality based on a test of the skewness and kurtosis of the residuals. HETR represents the heteroscedasticity test based on the regression of squared residuals on the squared fitted values. The P-values reported for diagnostic tests are based on the *F*-test except NORM, which uses the Lagrange multiplier version. All the *P*-values are given in the brackets and the values in parentheses are the standard errors. The significance level of the coefficients at 1%, 5% and 10% are indicated by \*\*\*, \*\* and \*, respectively. |

*4.3**Robustness*

As a robustness check, the testing and estimation procedure was repeated covering the monetary policy activism phase in Pakistan from 1971–2010. In this period, the average *M2* growth remained at 15.45% compared to 11.33% during 1961–1970. The monetary activism phase was characterized by high average inflation (9.39%) and relatively lower average real growth at 4.90%, which is in sharp contrast to the 3.51% inflation and 7.24% real growth in the moderate period from 1960–1970.[[14]](#footnote-14) For estimation purposes, the initial two years of 1971 and 1972 were excluded from the analysis to eliminate the potential effect of Pakistan’s war with India in 1971. This war badly affected real growth rates in Pakistan as, on average, a growth rate of 0.64% was witnessed for 1971 and 1972. Since the country also experienced an all-time high average inflation rate of ~24% for 1973–1975 because of international oil price shocks and domestic floods, a dummy variable was introduced to control for these effects, which was dropped subsequently because of its insignificance.[[15]](#footnote-15)

The results of the activist phase also confirm a significant long-term negative inflation bias-growth nexus for all the indicators at 1% level (Table 4).[[16]](#footnote-16) In fact, the fit of the data for all the models has improved and they pass important diagnostic tests. Interestingly, for the activist monetary policy period, the inflation bias indicators (*IB3* and *IB4*) are also significant and their effect is quantitatively larger than the effect of *IB1* and *IB2*. This implies that the more the inflation departs from desirable levels, the more are the adverse effects of inflation bias on real growth and vice versa. In other words, the higher the average inflation bias, the more it destabilizes real growth and vice versa (Figure 3). For example, the average inflation bias computed from the observed inflation $π>2\%$ is 8.87% and that for $π>5\%$ is 10.27%. Moreover, in the case of *IB1*, a 1% increase in inflation bias reduces real growth by 0.05%, whereas in the case of *IB4,* the corresponding detrimental effect is 1.21%.

|  |
| --- |
| ***Table 4: Long-term parameter estimates of the inflation bias indicators (1973–2010)*** |
| Models /Variables | Variables | Fit of the models and the diagnostic tests |
| $$IB1$$ | $$IB2$$ | $$IB3$$ | $$IB4$$ | $$\hat{ POP}$$ | $$\hat{ INV}$$ | $$\frac{FDI}{GDP}$$ | $$γ$$ | $$R^{2}$$ | AUTO | SPEC | NORM | HETR |
| Model 1 ($IB1$) | -0.06\*\*\* |  |  |  | 0.99 | 0.14\*\*\* | 23.00 | 4.21 |  |  |  |  |  |
|  | (0.01) |  |  |  | (0.64) | (0.05) | (24.26) | (2.10) | 0.50 | [0.36] | [0.66] | [0.61] | [0.65] |
|  | [0.00] |  |  |  | [0.13] | [0.00] | [0.35] | [0.05] |  |  |  |  |  |
| Model 2 ($IB2$) |  | -0.13\*\*\* |  |  | 0.99 | 0.14\*\*\* | 22.99 | 3.95 |  |  |  |  |  |
|  |  | (0.04) |  |  | (0.64) | (0.05) | (24.26) | (2.08) | 0.50 | [0.35] | [0.66] | [0.61] | [0.65] |
|  |  | [0.00] |  |  | [0.13] | [0.00] | [0.35] | [0.06] |  |  |  |  |  |
| Model 3 ($IB3$) |  |  | -0.20\*\*\* |  | 0.99 | 0.14\*\*\* | 22.95 | 3.69 |  |  |  |  |  |
|  |  |  | (0.06) |  | (0.64) | (0.05) | (24.25) | (2.06) | 0.49 | [0.35] | [0.66] | [0.61] | [0.65] |
|  |  |  | [0.00] |  | [0.13] | [0.00] | [0.35] | [0.08] |  |  |  |  |  |
| Model 4 ($IB4$) |  |  |  | -1.21\*\*\* | 0.80 | 0.14\*\* | 18.22 | 3.87 |  |  |  |  |  |
|  |  |  |  | (0.39) | (0.65) | (0.05) | (24.41) | (2.08) | 0.48 | [0.48] | [0.72] | [0.58] | [0.67] |
|   |   |   |   | [0.00] | [0.22] | [0.01] | [0.46] | [0.07] |   |   |   |   |   |
| This table reports the cointegrating relationship of the real GDP and the inflation bias indicators. The *P*-values of the diagnostic tests are presented sequentially with AUTO denoting the Langrange multiplier test for autocorrelation. SPEC represents a general test for omitted variables and a functional form test (the Ramsey regression equation specification error test (RESET)) using the square of the fitted values. NORM indicates the test for normality based on a test of the skewness and kurtosis of the residuals. HETR represents the heteroscedasticity test based on the regression of squared residuals on the squared fitted values. The P-values reported for diagnostic tests are based on the *F*-test except NORM, which uses the Lagrange multiplier version. All the *P*-values are given in the brackets and the values in parentheses are the standard errors. The significance level of the coefficients at 1%, 5% and 10% are indicated by \*\*\*, \*\* and \*, respectively. |

1. **Conclusion**

This paper contributed by proposing a methodological framework for modeling inflation bias to be able to generate its time series indicators to empirically assess its effectiveness in terms of stabilizing growth in the context of a developing country. This investigation is important because even today, most developing countries like Pakistan exercise their discretion to stimulate real growth at the cost of having higher inflation the way the advanced countries used to in the great inflation era of the 1970s. The question whether inflation bias really does the job lies at the core of this research. While stressing the need for having the conceptual distinction between inflation and inflation bias *per se* in empirical investigations, the paper proposed a framework that allows the generation of inflation bias indicators on the basis of desirable and threshold inflation rates, and society’s preferences for these. Four time series indicators of inflation bias with varying degrees of bias were therefore generated and subsequently modeled with real growth to explore their long-term nexus. The results indicate that discretionary monetary policy has resulted in a non-negligible degree of inflation bias, which has had increasingly detrimental effects on real growth: the more inflation departs from low levels, the more it becomes inimical. Largely, this result is robust across the generated inflation bias indicators and the subsample analysis. Contrary to the conventional idea that relatively higher inflation rates are better for developing countries, these findings suggest that price-stability-consistent monetary policy practices should be adopted by the central bank of Pakistan. Rather than actively pursuing higher than potential real growth rates via monetary policy, which induces inflation bias, the State Bank of Pakistan will have to reorient its monetary policy design to achieve its real growth objective by means of low and stable inflation, preferably at 3% or below. Since discretionary monetary policy practices may predominantly be found in developing countries, which are normally eager to attain relatively speedier economic growth, country-specific research along similar lines may be instrumental in exploring if, like Pakistan, they might also be (although unknowingly) sailing in a similar boat steered in a direction that may not lead to the destination of prosperity.

**References**

Abo-Zaid, S., and Tuzemen, D. (2012). Inflation targeting: a three decade perspective. *Journal of Policy Modeling. 34,* 621-645.

Barro, R. J. (1990). Government spending in a simple model of endogenous growth. *Journal of Political Economy, 98*(2), 103–125.

Barro, R. J. (1991). Economic growth in a cross section of countries. *Quarterly Journal of Economics, 106*, 407–444.

Barro, R. J. (1995). *Inflation and economic growth*. [NBER Working Paper No. 5326].

Barro, R. J., and Gordon, D. B. (1983a). A positive theory of monetary policy in a natural rate model. *Journal of Political Economy, 91*, 589–610.

Barro, R. J., and Gordon, D. B. (1983b). Rules, discretion and reputation in a model of monetary policy. *Journal of Monetary Economics, 12*, 101–121.

Barro, R. J., and Sala-i-Martin, X. (1992). Convergence. *Journal of Political Economy, 100*, 223–251.

Barro, R. J., and Sala-i-Martin, X. (1995). *Economic Growth*. New York: McGraw-Hill.

Bec, F., Salem, M., and Collard, F. (2002). Asymmetries in monetary policy reaction function, evidence for the U.S., French and German central banks. In B. Mizrach (Ed.), *Studies in Nonlinear Dynamics and Econometrics, Vol. 6(2)*.

Berlemann, M. (2005). Time inconsistency of monetary policy: empirical evidence from polls. *Public Choice, 125*, 1–15.

Blanchard, O., DellAriccia, G., and Mauro, P. (2010). Rethinking macroeconomic policy. *Journal of Money, Credit and Banking*, 42, 199–215.

Blinder, A. S. (1998). *Central banking in theory and practice*: MIT Press, Cambridge.

Burdekin, R. C. K., Denzau, A. T., Keil, M. W., Sitthiyot, T., & Willett, T. D. (2004). When does inflation hurt economic growth? different nonlinearities for different countries. *Journal of Macroeconomics*, 26(2004), 519–532.

Concalves, C. E. S., and Salles, J. M. (2008). Inflation targeting in emerging economies: what do the data say? *Journal of Development Economics, 85(1)*, 312–318.

Cukierman, A. (2000). *The inflation bias result revisited*. [Berglas School of Economics, Tel Aviv University].

Cukierman, A., & Gerlach, S. (2003). The inflation bias revisited: theory and some international evidence. [The Manchester School]. 71, 541-565.

De Gregorio, J. (1992). The effects of inflation on economic growth. European Economic Review, 36(2-3), 417-424.

De Gregorio, J. (1993). Inflation, taxation and long-run growth. *Journal of Monetary Economics, 31*(271-298).

Dotsey, M. (2008). Commitment versus discretion in monetary policy. [Federal Reserve Bank of Philadelphia, Business Review]. (Q4), 1–8.

Drukker, D., Gomis-Porqueras, P., & Hernandez-Verme, P. (2005). Threshold effects in the relationship between inflation and growth: a new panel-data approach. [Proceedings of the 11th International Conference on Panel Data].

Enders, W. (1995). Applied econometric time series: John Wiley and Sons, USA.

Engle, R. (2003). Risk and volatility: econometric models and financial practice. Nobel Lecture, December 8, 2003.

Fischer, S. (1993). The role of macroeconomic factors in growth. *Journal of Monetary Economics, 32*(3), 485-511.

Friedman, M. (1977). Nobel Lecture: inflation and unemployment. *Journal of Political Economy*. 85, 451–472.

Garman, D. M., and Richards, D. J. (1989). Policy rules, inflationary bias and cyclical stability. *Journal of Money, Credit and Banking*, 21, 409–421.

Gartner, M. (2000). Political macroeconomics: a survey of recent developments. Journal of Economic Surveys, 14, 527-561.

Hanif, M.N. (2014). Monetary policy experience of Pakistan. [MPRA Paper No. 60855].

Hayat, Z., Balli, F., Obben, J., Shakur, S. (2016). An empirical assessment of monetary discretion: the case of Pakistan. *Journal of Policy Modeling. 38(5)*, 954–970.

Hayat, Z. U. C., & Kalirajan, K. P. (2009). Is there a threshold level of inflation for Bangladesh? *The Journal of Applied Economic Research, 3(1)*, 1-20.

Hossain, A.A. (2014). Inflation and inflation volatility in Australia. *Economic Papers*. *33(2),* 163–185.

Ireland, P. N. (1999). Does the time-inconsistency problem explain the behavior of inflation in the United States? *Journal of Monetary Economics*, 44, 279–291.

Kannan, R., & Joshi, H. (1998). Growth-inflation trade-off: Empirical estimation of threshold rate of inflation for India. [Economic and Political Weekly 33(42/43)]. 2724-2728.

Khan, M. S., and Senhadji, A. S. (2001). Threshold effects in the relationship between inflation and growth. *IMF Staff Papers*. 48, 1–21.

Kydland, F. E., and Prescott, E. C. (1977). Rules rather than discretion: the inconsistency of optimal plans. *Journal of Political Economy* 85, 473–492.

Levine, R., and Renelt, D. (1991). *Cross country studies of growth and policy: Some methodological, conceptual and statistical problems*. [World Bank Working Paper Series No. 608].

Levine, R., and Renelt, D. (1992). A sensitivity analysis of cross-country growth regressions. *American Economic Review*, 82, 942–963.

Lin, S., & Ye, H. (2009). Does inflation targeting make a difference in developing countries? *Journal of Development Economics, 89*, 118–123.

Mohanty, D., Chakraborty, A., Das, A., & John, J. (2011). Inflation threshold in India: an empirical investigation. [RBI Working Paper Series 18 (2011)].

Motley, B. (1998). Growth and inflation: a cross-country study. *Economic Review – Federal Reserve*

*Bank of San Francisco, (1)*, 15.

Mubarik, Y. (2005). Inflation and growth: an estimate of the threshold level of inflation in Pakistan. *SBP- Research Bulletin*. 1, 35–44.

Munir, Q., & Mansur, K. (2009). Non-linearity between inflation rate and GDP growth in Malaysia. *Economics Bulletin*. *29*(3), 1555-1569.

Narayan, P. K. (2005). The saving and investment nexus for China: evidence from cointegration tests. *Applied Economics*, 37(17), 1979–1990.

Iqbal, N., & Nawaz, S. (2010). Investment, inflation and economic growth nexus. *Pakistan Development Review*.

Pesaran, H. M., and Pesaran, B. (1997). *Working with Microfit 4.0: interactive econometric analysis*. Oxford, UK: Oxford University Press.

Pesaran, M. H., and Shin, Y. (1999). An auto regressive distributed lag modelling approach to cointegration analysis. In S. Strom (Ed.) *Econometrics and Economic Theory in the 20th Century: The Ragner Frisch Centennial Symposium*. Cambridge, UK: Cambridge University Press.

Pesaran, B., and Pesaran, M.H. (2009). Time series econometrics using Microfit 5.0. Oxford University Press

Pesaran, M. H., Shin, Y., and Smith, R. J. (2001). Bounds testing approaches to the analysis of level relationships. J*ournal of Applied Econometrics*, 16, 289–326.

Romer, P. M. (1989). *Human Capital and Growth: Theory and Evidence*. [NBER Working Paper No.3173].

Romer, P. M. (1990b). Endogenous technological change. *Journal of Political Economy*, 98(2), 71–102.

Romer, D. (1993). Openness and inflation: theory and evidence. *Quarterly Journal of Economics*, 98, 869–903.

Romer, D., & Romer, C. (2002). The evolution of economic understanding and postwar stabilization policy. In: *Federal Reserve Bank of Kansas City Symposium on Rethinking Stabilization Policy*. 11–78.

Romer, D. (2006). *Advanced Macroeconomics (3rd ed.)*. McGraw-Hill/Irwin Publishers, New York.

Ruge-Murcia, FJ. (2004). The inflation bias when the central bank targets the natural rate of unemployment. *European Economic Review*, 48, 91–107.

Sarel, M. (1996). Nonlinear effects of inflation on economic growth. *IMF Staff Papers*. 43(1).

Surico, P. (2008). Measuring the time inconsistency of US monetary policy. *Economica*, 75, 22–38.

Wilson, B. K. (2006). The links between inflation, inflation uncertainty and output growth: new time series evidence from Japan. *Journal of Macroeconomics, 28*(2006), 609-620.

|  |
| --- |
| ***Appendix 1: Long-term parameter estimates of the baseline growth model and simulation results*** |
| Models/Variables | Variables |   | Fit of the models and the diagnostic tests |
| $$π$$ | $$\hat{ POP}$$ | $$\hat{ INV}$$ | $$\frac{FDI}{GDP}$$ | $$DUM$$ | $$α$$ | $$R^{2}$$ | ARDL | FSTS | COIN | AUTO | SPEC | NORM | HETR |
| Baseline Model | -0.24\*\* | 0.95 | 0.16\*\*\* | 23.78 |  | 4.20 |  |  |  |  |  |  |  |  |
|  | (0.08) | (0.74) | (0.04) | (27.78) |  | (2.11) | 0.46 | 0,2,1,0,1 | 7.34 | 1% | [0.96] | [0.22] | [0.16] | [0.85] |
|  | [0.01] | [0.21] | [0.00] | [0.39] |  | [0.05] |  |  |  |  |  |  |  |  |
| Model 1 (INF=1) | -4.63\* | 1.32\* | 0.17\*\*\* | 28.05 | 4.46\* | 6.75 |  |  |  |  |  |  |  |  |
|  | (2.61) | (0.73) | (0.04) | (27.55) | (2.62) | (3.09) | 0.44 | 0,2,0,0,1 | 6.42 | 1% | [0.81] | [0.01] | [0.95] | [0.34] |
|  | [0.08] | [0.07] | [0.00] | [0.31] | [0.09] | [0.04] |  |  |  |  |  |  |  |  |
| Model 2 (INF=2) | -2.19\* | 1.32\* | 0.17\*\*\* | 28.05 | 2.02\* | 6.33 |  |  |  |  |  |  |  |  |
|  | (1.18) | (0.72) | (0.04) | (27.55) | (1.18) | (2.92) | 0.44 | 0,2,0,0,1 | 6.37 | 1% | [0.81] | [0.01] | [0.95] | [0.34] |
|  | [0.07] | [0.07] | [0.00] | [0.31] | [0.09] | [0.04] |  |  |  |  |  |  |  |  |
| Model 3 (INF=3) | -1.48\* | 1.33\* | 0.17\*\*\* | 28.31 | 1.31\* | 6.17 |  |  |  |  |  |  |  |  |
|  | (0.76) | (0.72) | (0.04) | (27.57) | (0.76) | (2.85) | 0.44 | 0,2,0,0,1 | 6.35 | 1% | [0.83] | [0.01] | [0.96] | [0.34] |
|  | [0.06] | [0.07] | [0.00] | [0.31] | [0.09] | [0.04] |  |  |  |  |  |  |  |  |
| Model 4 (INF=4) | -0.79 | 1.10 | 0.17\*\*\* | 28.47 | 0.58 | 5.73 |  |  |  |  |  |  |  |  |
|  | (0.53) | (0.75) | (0.04) | (27.41) | (0.55) | (2.57) | 0.47 | 0,2,1,0,1 | 6.17 | 1% | [0.83] | [0.09] | [0.34] | [0.56] |
|  | [0.14] | [0.15] | [0.00] | [0.31] | [0.30] | [0.03] |  |  |  |  |  |  |  |  |
| Model 5 (INF=5) | -0.45 | 1.05 | 0.16\*\*\* | 26.70 | 0.23 | 4.80 |  |  |  |  |  |  |  |  |
|  | (0.39) | (0.76) | (0.04) | (27.41) | (0.40) | (2.39) | 0.46 | 0,2,1,0,1 | 6.27 | 1% | [0.89] | [0.16] | [0.26] | [0.73] |
|  | [0.25] | [0.18] | [0.00] | [0.34] | [0.58] | [0.05] |  |  |  |  |  |  |  |  |
| Model 6 (INF=6) | -0.19 | 0.92 | 0.15\*\*\* | 22.98 | -0.05 | 4.06 |  |  |  |  |  |  |  |  |
|  | (0.28) | (0.76) | (0.04) | (27.41) | (0.31) | (2.58) | 0.46 | 0,2,1,0,1 | 6.13 | 1% | [0.96] | [0.24] | [0.15] | [0.86] |
|   | [0.50] | [0.23] | [0.00] | [0.41] | [0.87] | [0.08] |   |   |   |   |   |   |   |   |
| This table reports the cointegrating relationship of real GDP and its potential determinants in a multivariate setting. $DUM$ is the interactive dummy ($D\_{t}. \left(π\_{t}^{0}-π\_{a}\right))$. ARDL represents the lag order of the variables as selected by the SBC. FSTS shows the computed *F*-statistics used to test for the existence of a cointegrating relationship. The upper critical value bound of Pesaran *et al.* (2001) for *k*=6 at the 1% level is 4.43. COIN indicates the cointegration at a certain level of confidence. The *P*-values of the diagnostic tests are presented sequentially, with AUTO denoting the Lagrange multiplier test for autocorrelation. SPEC represents a general test for omitted variables and a functional form test (the Ramsey regression equation specification error test (RESET)) using the square of the fitted values. NORM indicates the test for normality based on a test of the skewness and kurtosis of the residuals. HETR represents the heteroscedasticity test based on the regression of squared residuals on the squared fitted values. The *P*-values reported for diagnostic tests are based on the *F*-test, except for NORM, which uses the Lagrange multiplier version. All the *P*-values are given in brackets and the values in parentheses are the standard errors. The error correction term $ECM(-1)$ may not be obtained because the SBC did not select the lag-dependent variable as optimal. Technically, ARDL models tend to reduce to dynamic distributed lag models if the model selection criterion does not identify any lag of the regressand as optimal. The significance of the coefficients at the 1%, 5% and 10% levels are indicated by \*\*\*, \*\* and \*, respectively.  |
| ***Appendix 1 Continued ……….. simulation results*** |
| Models/Variables | Variables |   | Fit of the models and the diagnostic tests |
| $$π$$ | $$\hat{ POP}$$ | $$\hat{ INV}$$ | $$\frac{FDI}{GDP}$$ | $$DUM$$ | $$α$$ | $$R^{2}$$ | ARDL | FSTS | COIN | AUTO | SPEC | NORM | HETR |
| Model 7 (INF=7) | -0.10 | 0.88 | 0.15\*\*\* | 21.68 | -0.18 | 3.68 |  |  |  |  |  |  |  |  |
|  | (0.22) | (0.74) | (0.04) | (27.43) | (0.49) | (2.25) | 0.47 | 0,2,1,0,1 | 6.06 | 1% | [0.94] | [0.22] | [0.14] | [0.88] |
|  | [0.66] | [0.24] | [0.00] | [0.43] | [0.49] | [0.25] |  |  |  |  |  |  |  |  |
| Model 8 (INF=8) | -0.11 | 0.88 | 0.15\*\*\* | 21.47 | -0.18 | 3.67 |  |  |  |  |  |  |  |  |
|  | (0.18) | (0.74) | (0.04) | (27.32) | (0.22) | (2.21) | 0.47 | 0,2,1,0,1 | 6.05 | 1% | [0.91] | [0.20] | [0.13] | [0.90] |
|  | [0.56] | [0.24] | [0.00] | [0.43] | [0.41] | [0.10] |  |  |  |  |  |  |  |  |
| Model 9 (INF=9) | -0.11 | 0.88 | 0.15\*\*\* | 21.15 | -0.18 | 3.70 |  |  |  |  |  |  |  |  |
|  | (0.17) | (0.74) | (0.04) | (27.32) | (0.22) | (2.21) | 0.47 | 0,2,1,0,1 | 6.11 | 1% | [0.91] | [0.17] | [0.13] | [0.87] |
|  | [0.45] | [0.24] | [0.00] | [0.44] | [0.34] | [0.09] |  |  |  |  |  |  |  |  |
| Model 10 (INF=10) | -0.13 | 0.88 | 0.15\*\*\* | 20.98 | -0.19 | 3.78 |  |  |  |  |  |  |  |  |
|  | (0.13) | (0.73) | (0.04) | (27.25) | (0.19) | (2.16) | 0.47 | 0,2,1,0,1 | 6.14 | 1% | [0.92] | [0.16] | [0.13] | [0.83] |
|  | [0.32] | [0.24] | [0.00] | [0.44] | [0.34] | [0.08] |  |  |  |  |  |  |  |  |
| Model 11 (INF=11) | -0.14 | 0.87 | 0.15\*\*\* | 19.82 | -0.19 | 3.84 |  |  |  |  |  |  |  |  |
|  | (0.12) | (0.73) | (0.04) | (27.31) | (0.18) | (2.13) | 0.47 | 0,2,1,0,1 | 6.13 | 1% | [0.95] | [0.15] | [0.13] | [0.78] |
|  | [0.26] | [0.24] | [0.00] | [0.47] | [0.30] | [0.08] |  |  |  |  |  |  |  |  |
| Model 12 (INF=12) | -0.14 | 0.85 | 0.15\*\*\* | 18.06 | -0.21 | 3.90 |  |  |  |  |  |  |  |  |
|  | (0.11) | (0.73) | (0.04) | (27.40) | (0.18) | (2.11) | 0.48 | 0,2,1,0,1 | 6.11 | 1% | [0.97] | [0.14] | [0.13] | [0.75] |
|  | [0.21] | [0.25] | [0.00] | [0.51] | [0.25] | [0.07] |  |  |  |  |  |  |  |  |
| Model 13 (INF=13) | -0.15 | 0.85 | 0.15\*\*\* | 15.63 | -0.23 | 3.92 |  |  |  |  |  |  |  |  |
|  | (0.11) | (0.73) | (0.04) | (27.83) | (0.19) | (2.11) | 0.48 | 0,2,1,0,1 | 6.11 | 1% | [0.98] | [0.15] | [0.13] | [0.75] |
|   | [0.20] | [0.25] | [0.00] | [0.57] | [0.25] | [0.07] |   |   |   |   |   |   |   |   |
| This table reports the cointegrating relationship of real GDP and its potential determinants in a multivariate setting. $DUM$ is the interactive dummy ($D\_{t}. \left(π\_{t}^{0}-π\_{a}\right))$. ARDL represents the lag order of the variables as selected by the SBC. FSTS shows the computed *F*-statistics used to test for the existence of a cointegrating relationship. The upper critical value bound of Pesaran *et al.* (2001) for *k*=6 at the 1% level is 4.43. COIN indicates the cointegration at a certain level of confidence. The *P*-values of the diagnostic tests are presented sequentially, with AUTO denoting the Lagrange multiplier test for autocorrelation. SPEC represents a general test for omitted variables and a functional form test (the Ramsey regression equation specification error test (RESET)) using the square of the fitted values. NORM indicates the test for normality based on a test of the skewness and kurtosis of the residuals. HETR represents the heteroscedasticity test based on the regression of squared residuals on the squared fitted values. The *P*-values reported for diagnostic tests are based on the *F*-test, except for NORM, which uses the Lagrange multiplier version. All the *P*-values are given in brackets and the values in parentheses are the standard errors. The error correction term $ECM(-1)$ may not be obtained because the SBC did not select the lag-dependent variable as optimal. Technically, ARDL models tend to reduce to dynamic distributed lag models if the model selection criterion does not identify any lag of the regressand as optimal. The significance of the coefficients at the 1%, 5% and 10% levels are indicated by \*\*\*, \*\* and \*, respectively.   |

|  |
| --- |
| ***Appendix 1 Continued ……….. simulation results*** |
| Models/Variables | Variables |   | Fit of the models and the diagnostic tests |
| $$π$$ | $$\hat{ POP}$$ | $$\hat{ INV}$$ | $$\frac{FDI}{GDP}$$ | $$DUMM$$ | $$α$$ | $$R^{2}$$ | ARDL | FSTS | COIN | AUTO | SPEC | NORM | HETR |
| Model 14 (INF=14) | -0.14 | 0.83 | 0.15\*\*\* | 12.74 | -0.26 | 3.95 |  |  |  |  |  |  |  |  |
|  | (0.11) | (0.73) | (0.04) | (28.41) | (0.21) | (2.11) | 0.48 | 0,2,1,0,1 | 6.12 | 1% | [0.99] | [0.15] | [0.12] | [0.75] |
|  | [0.20] | [0.26] | [0.00] | [0.65] | [0.23] | [0.06] |  |  |  |  |  |  |  |  |
| Model 15 (INF=15) | -0.14 | 0.83 | 0.15\*\*\* | 12.23 | -0.30 | 3.96 |  |  |  |  |  |  |  |  |
|  | (0.11) | (0.73) | (0.04) | (28.47) | (0.24) | (2.10) | 0.48 | 0,2,1,0,1 | 6.13 | 1% | [0.99] | [0.16] | [0.11] | [0.76] |
|  | [0.20] | [0.26] | [0.00] | [0.67] | [0.22] | [0.06] |  |  |  |  |  |  |  |  |
| Model 16 (INF=16) | -0.17 | 0.83 | 0.15\*\*\* | 11.59 | -0.24 | 4.19 |  |  |  |  |  |  |  |  |
|  | (0.10) | (0.75) | (0.04) | (30.67) | (0.27) | (2.11) | 0.47 | 0,2,1,0,1 | 6.02 | 1% | [0.87] | [0.18] | [0.12] | [0.79] |
|  | [0.12] | [0.27] | [0.00] | [0.70] | [0.39] | [0.05] |  |  |  |  |  |  |  |  |
| Model 17 (INF=17) | -0.17 | 0.82 | 0.15\*\*\* | 11.05 | -0.29 | 4.19 |  |  |  |  |  |  |  |  |
|  | (0.11) | (0.75) | (0.04) | (30.44) | (0.31) | (2.11) | 0.47 | 0,2,1,0,1 | 6.07 | 1% | [0.88] | [0.18] | [0.11] | [0.80] |
|  | [0.11] | [0.28] | [0.00] | [0.71] | [0.36] | [0.05] |  |  |  |  |  |  |  |  |
| Model 18 (INF=18) | -0.17 | 0.82 | 0.15\*\*\* | 11.04 | -0.28 | 4.19 |  |  |  |  |  |  |  |  |
|  | (0.10) | (0.75) | (0.04) | (30.44) | (0.31) | (2.11) | 0.47 | 0,2,1,0,1 | 6.12 | 1% | [0.90] | [0.18] | [0.09] | [0.80] |
|  | [0.11] | [0.28] | [0.00] | [0.72] | [0.36] | [0.05] |  |  |  |  |  |  |  |  |
| Model 19 (INF=19) | -0.17\* | 0.82 | 0.15\*\*\* | 10.62 | -0.35 | 4.19 |  |  |  |  |  |  |  |  |
|  | (0.10) | (0.74) | (0.04) | (30.11) | (0.28) | (2.11) | 0.48 | 0,2,1,0,1 | 6.15 | 1% | [0.93] | [0.19] | [0.08] | [0.82] |
|  | [0.11] | [0.27] | [0.00] | [0.72] | [0.32] | [0.35] |  |  |  |  |  |  |  |  |
| Model 20 (INF=20) | -0.17\* | 0.82 | 0.15\*\*\* | 10.58 | -0.43 | 4.19 |  |  |  |  |  |  |  |  |
|  | (0.10) | (0.74) | (0.04) | (29.64) | (0.40) | (2.10) | 0.48 | 0,2,1,0,1 | 6.15 | 1% | [0.96] | [0.20] | [0.06] | [0.83] |
|   | [0.09] | [0.27] | [0.00] | [0.72] | [0.28] | [0.05] |   |   |   |   |   |   |   |   |
| This table reports the cointegrating relationship of real GDP and its potential determinants in a multivariate setting. $DUM$ is the interactive dummy ($D\_{t}. \left(π\_{t}^{0}-π\_{a}\right))$. ARDL represents the lag order of the variables as selected by the SBC. FSTS shows the computed *F*-statistics used to test for the existence of a cointegrating relationship. The upper critical value bound of Pesaran *et al.* (2001) for *k*=6 at the 1% level is 4.43. COIN indicates the cointegration at a certain level of confidence. The *P*-values of the diagnostic tests are presented sequentially, with AUTO denoting the Lagrange multiplier test for autocorrelation. SPEC represents a general test for omitted variables and a functional form test (the Ramsey regression equation specification error test (RESET)) using the square of the fitted values. NORM indicates the test for normality based on a test of the skewness and kurtosis of the residuals. HETR represents the heteroscedasticity test based on the regression of squared residuals on the squared fitted values. The *P*-values reported for diagnostic tests are based on the *F*-test, except for NORM, which uses the Lagrange multiplier version. All the *P*-values are given in brackets and the values in parentheses are the standard errors. The error correction term $ECM(-1)$ may not be obtained because the SBC did not select the lag-dependent variable as optimal. Technically, ARDL models tend to reduce to dynamic distributed lag models if the model selection criterion does not identify any lag of the regressand as optimal. The significance of the coefficients at the 1%, 5% and 10% levels are indicated by \*\*\*, \*\* and \*, respectively.  |

|  |
| --- |
| ***Appendix 1 Continued ……….. simulation results*** |
| Models/Variables | Variables |   | Fit of the models and the diagnostic tests |
| $$π$$ | $$\hat{ POP}$$ | $$\hat{ INV}$$ | $$\frac{FDI}{GDP}$$ | $$DUMM$$ | $$α$$ | $$R^{2}$$ | ARDL | FSTS | COIN | AUTO | SPEC | NORM | HETR |
| Model 21 (INF=21) | -0.18\* | 0.84 | 0.15\*\*\* | 11.35 | -0.53 | 4.19 |  |  |  |  |  |  |  |  |
|  | (0.09) | (0.73) | (0.04) | (28.98) | (0.45) | (2.10) | 0.48 | 0,2,1,0,1 | 6.11 | 1% | [0.99] | [0.22] | [0.05] | [0.86] |
|  | [0.07] | [0.26] | [0.00] | [0.69] | [0.25] | [0.05] |  |  |  |  |  |  |  |  |
| Model 22 (INF=22) | -0.19\*\* | 0.86 | 0.15\*\*\* | 13.19 | -0.63 | 4.18 |  |  |  |  |  |  |  |  |
|  | (0.09) | (0.73) | (0.04) | (28.23) | (0.51) | (2.09) | 0.48 | 0,2,1,0,1 | 6.05 | 1% | [0.94] | [0.21] | [0.05] | [0.87] |
|  | [0.04] | [0.24] | [0.00] | [0.64] | [0.2] | [0.05] |  |  |  |  |  |  |  |  |
| Model 23 (INF=23) | -0.18\*\* | 0.86 | 0.15\*\*\* | 13.15 | -0.80 | 4.19 |  |  |  |  |  |  |  |  |
|  | (0.08) | (0.73) | (0.04) | (28.05) | (0.61) | (2.09) | 0.48 | 0,2,1,0,1 | 5.91 | 1% | [0.85] | [0.22] | [0.04] | [0.90] |
|  | [0.04] | [0.24] | [0.00] | [0.64] | [0.20] | [0.05] |  |  |  |  |  |  |  |  |
| Model 24 (INF=24) | -0.19\*\* | 0.87 | 0.15\*\*\* | 14.09 | -1.01 | 4.19 |  |  |  |  |  |  |  |  |
|  | (0.08) | (0.73) | (0.04) | (27.75) | (0.75) | (2.09) | 0.48 | 0,2,1,0,1 | 5.90 | 1% | [0.84] | [0.22] | [0.04] | [0.90] |
|  | [0.02] | [0.24] | [0.00] | [0.61] | [0.19] | [0.05] |  |  |  |  |  |  |  |  |
| Model 25 (INF=25) | -0.19\*\* | 0.87 | 0.14\*\*\* | 14.23 | -1.37 | 4.19 |  |  |  |  |  |  |  |  |
|  | (0.08) | (0.73) | (0.04) | (27.73) | (1.02) | (2.09) | 0.48 | 0,2,1,0,1 | 5.90 | 1% | [0.84] | [0.22] | [0.04] | [0.90] |
|  | [0.02] | [0.24] | [0.00] | [0.61] | [0.19] | [0.05] |  |  |  |  |  |  |  |  |
| Model 26 (INF=26) | -0.19\*\* | 0.87 | 0.15\*\*\* | 14.23 | -2.20 | 4.19 |  |  |  |  |  |  |  |  |
|  | (0.08) | (0.73) | (0.04) | (27.73) | (1.64) | (2.09) | 0.48 | 0,2,1,0,1 | 5.90 | 1% | [0.84] | [0.22] | [0.04] | [0.90] |
|   | [0.02] | [0.24] | [0.00] | [0.61] | [0.19] | [0.05] |   |   |   |   |   |   |   |   |
| This table reports the cointegrating relationship of real GDP and its potential determinants in a multivariate setting. $DUM$ is the interactive dummy ($D\_{t}. \left(π\_{t}^{0}-π\_{a}\right))$. ARDL represents the lag order of the variables as selected by the SBC. FSTS shows the computed *F*-statistics used to test for the existence of a cointegrating relationship. The upper critical value bound of Pesaran *et al.* (2001) for *k*=6 at the 1% level is 4.43. COIN indicates the cointegration at a certain level of confidence. The *P*-values of the diagnostic tests are presented sequentially, with AUTO denoting the Lagrange multiplier test for autocorrelation. SPEC represents a general test for omitted variables and a functional form test (the Ramsey regression equation specification error test (RESET)) using the square of the fitted values. NORM indicates the test for normality based on a test of the skewness and kurtosis of the residuals. HETR represents the heteroscedasticity test based on the regression of squared residuals on the squared fitted values. The *P*-values reported for diagnostic tests are based on the *F*-test, except for NORM, which uses the Lagrange multiplier version. All the *P*-values are given in brackets and the values in parentheses are the standard errors. The error correction term $ECM(-1)$ may not be obtained because the SBC did not select the lag-dependent variable as optimal. Technically, ARDL models tend to reduce to dynamic distributed lag models if the model selection criterion does not identify any lag of the regressand as optimal. The significance of the coefficients at the 1%, 5% and 10% levels are indicated by \*\*\*, \*\* and \*, respectively.   |

***Appendix 2: Stability tests (baseline growth model)***



***Appendix 3: Stability tests for the inflation bias–growth nexus models (1961–2010)***





***Appendix 4: Stability tests for the inflation bias–growth nexus models (1973–2010)***



1. This may predominantly be the case where the central banks have made relatively less progress in acknowledging and adopting price stability-consistent monetary policy practices. For example, the average inflation rates in developing countries that did not adopt inflation targeting remained higher than those in countries that adopted it (Concalves and Salles, 2008; Lin and Ye, 2009; Abo-zaid and Tuzemen, 2012). [↑](#footnote-ref-1)
2. Bec *et al.* (2002) noted that inflation bias arises from two features of monetary policy behavior: first, the twofold objectives of inflation and output; second, targeting output beyond the potential level of the economy. [↑](#footnote-ref-2)
3. This might not necessarily be true in case of advanced countries, nevertheless in case of developing countries like Pakistan, this assumption is realistic in the sense that a big chunk of the population (40%) is living below the poverty line and another (20%) is close to it, and they are therefore more vulnerable to high inflation rates. At-least this segment of the overall population tend to prefer low inflation given the fact that Pakistan is not a welfare state and that any possible short-term growth spurts that may accrue from high inflation may not be inclusive but instead would benefit a particular small segment in the society. Therefore in general our assumption is plausible. [↑](#footnote-ref-3)
4. The number of values restricted to zero is 2, 4, 5 and 14 out of 50 for the four inflation bias indicators, respectively. [↑](#footnote-ref-4)
5. 26% is the maximum inflation rate ever experienced in Pakistan. It should be noted that the inflation rates were rounded off to the nearest percentage point {1%, 2%, 3…26%} because assuming continuity is otherwise overwhelmingly challenging and has no direct relevance to policy. [↑](#footnote-ref-5)
6. We estimated cointegrating relationships as this is the most appropriate way to avoid spurious results (in time series data) through the *ARDL* bounds testing and estimation strategy of Pesaran *et al.* (2001) while using the asymptotic critical bound values of Pesaran and Pesaran (2009) and Narayan (2005). This testing and estimation approach is preferred over conventional cointegration approaches as it is suitable for regressors integrated of order I(0), I(1) or both. The estimators of *ARDL* are superconsistent for long-run coefficients and perform well in small samples without losing long-run information. *ARDL* uses a two-step strategy, which works well even in the presence of endogenous regressors, irrespective of the order of integration of the explanatory variables (Pesaran and Pesaran, 1997; Pesaran and Shin, 1999). In the first step, the existence of a cointegrating relationship is established through an *F*-test. Since the asymptotic distribution of this *F*-test is non-standard, Pesaran *et al.* (2001) computed and tabulated its critical values for different orders of integration for the number of regressors with and without an intercept. If cointegration is established in the first step, in the second step, the long- and short-run coefficients are obtained. [↑](#footnote-ref-6)
7. Although data for few more years is available, we use 2010 as cutoff because of the structural policy shift as the central bank of Pakistan went to an interest rate corridor system from targeting monetary aggregates in order to stabilize interest rates and hence inflation (Hanif, 2014). [↑](#footnote-ref-7)
8. The cumulative sum of squares of the residuals ($CUSUM)$ and the cumulative sum of squares of the recursive residuals ($CUSUMQ)$ suggest that the long-run estimates are derived from stable regression functions (Appendices 1–2), implying that the regression coefficients do not exhibit systematic changes and sudden departures from constancy. [↑](#footnote-ref-8)
9. In order to ascertain the robustness of the baseline growth model, a check was conducted by regressing the subsample period from 1973 to 2010. Since the overall sample (50 observations) was not large enough to split into two equal parts, only the activist monetary policy period from 1971 to 2010 was considered. In this period, the average inflation, M2 and real growth rates were 9.39%, 15.45% and 4.9% compared to 3.51%, 11.33% and 7.24% in 1961–1970, respectively, which clearly indicates monetary activism. The initial two years (1971 and 1972) were dropped from the estimation to eliminate the potential effects of Pakistan’s war with India in 1971. This war badly affected the real growth rates in Pakistan as, on average, a growth rate of 0.64% was witnessed for 1971 and 1972. The computed *F*-statistic for the short sample is 8.26, which is much greater than the asymptotic upper critical value bound (5.06) at the 1% level, which confirms the existence of a long-term cointegrating relationship. These results are not reported to save space but may be obtained from the authors upon request. Furthermore, consistent with the approach of Levine and Renelt (1992), our specified model is robust in the sense that relatively fragile variables have been dropped and only the variables that could not be excluded on the basis of the variable deletion test have been retained in the model. It is also important to mention that the results of the model are not robust to alternative specifications, which would be too ambitious to expect because numerous variables may not necessarily be cointegrated, and may interact both with the dependent and independent variables to produce different results from our model. [↑](#footnote-ref-9)
10. As we would see in the next subsection, this problem of fit of the model goes away, when the well know golden period of the 1960s of Pakistan economy is excluded where average inflation was low and average growth was high. [↑](#footnote-ref-10)
11. Since the results of our study have considerable implications for conducting monetary policy in a developing country context, further country specific-research by other developing countries with similar discretionary monetary policy practices may be instrumental in streamlining the priorities of their central banks in terms of inflation and growth objectives. [↑](#footnote-ref-11)
12. For the monetary policy practice of keeping inflation close to or below the 2% level, the FOMC’s and ECB’s definitions may be found on their official websites. [↑](#footnote-ref-12)
13. The average is obtained for the 1961–2010 period after excluding the observed inflation rates that are equal to or below the 2% level. There could be supply side factors that may have contributed to part of the high average inflation bias, thus exaggerating the bias part; however, disentangling their impact from the overall bias is beyond the scope of this study. [↑](#footnote-ref-13)
14. For a robustness check to ascertain the long-term inverse relationship between inflation bias and real growth in a conventional way, bifurcating the sample seemed inappropriate because the observations in the two samples would be too small to allow for dynamics and to obtain any reliable long-term coefficients. [↑](#footnote-ref-14)
15. The joint test of zero restrictions on the coefficient of the dummy variable also supports the idea that it should be dropped from all the models. For example, the *P*-values of the LM test for the dummies in the models with *IB1*, *IB2*, *IB3* and *IB4* are 0.624, 0.624, 0.621 and 0.805, respectively. [↑](#footnote-ref-15)
16. To test for the existence of cointegration in the subperiod, the null and alternative hypotheses were formulated in the same fashion as those for the full sample. The SBC model selection criterion was used to select optimal lags while imposing a maximum lag of 3 years. The computed *F*-statistics for the four regressions containing *IB1*, *IB2*, *IB3* and *IB4* are 8.24, 8.46, 8.22 and 7.71, respectively. Since all these *F*-statistics are greater than the corresponding asymptotic critical values at the 1% level both for Pesaran and Pesaran (2009) and Narayan (2005), the long-term parameter estimates were obtained. All these models are stable, as can be seen from the stability tests in Appendix 4. [↑](#footnote-ref-16)