Causality Relationship between Central Bank Reforms and Inflation: Evidence from Developing Countries

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Abstract
This study provides evidence on the relationship between central bank reforms and inflation dynamics in a sample of 37 developing countries. We use panel structural break test and Granger non-causality tests on annual inflation and the legal index of central bank independence (CBI), as a proxy of central bank reform, over 40 years period. The empirical results indicate a positive effect of central bank independence on inflation stabilization. Besides, we find that there exists bi-directional causality between central bank reforms and inflation. These findings suggest that central bank independence is beneficial in terms of sustained macroeconomic stabilization and should harness among developing countries. In particular, reforms should design to give central banks more autonomy in the conduct of monetary policy and financial sector regulation.

Keywords: central bank independence, inflation, developing countries, structural break, Granger causality

JEL Classifications: E31, E58

How to Cite:
Introduction

Central banks hold an essential role in achieving macroeconomic stability, particularly in influencing the level of inflation and stability of the exchange rate. Central bank independence provides the central bank with autonomy from interference in the pursuance of its monetary policy objective and financial sector stability. This condition gives the central bank the freedom to focus on the attainment of the monetary policy goal to control the inflation rate (Berger et al., 2000), reducing the output gap and unemployment in some countries. Central bank independence can classify into three aspects, such as personnel, financial, and policy independence. Personnel independence refers to the fact that the government has restricted influence on the central bank’s Governor and its Boards. On this, Neumann (1991) argues that the public might view government influence as encouraging the central bank to pursue the kinds of policies that are in the government’s interests. The second aspect is financial independence, which indicates the central bank’s ability to budget and fund its activities without resorting to printing money to finance budget deficits (Eijffinger et al., 1998). The third is that policy independence reflects on the central bank’s ability to choose an objective, instrument, and the desecration to implement the monetary policy without any political interference.

One of the primary aims of central bank independence is to address the time-inconsistency problem as a consequence of the policy that is no longer optimal, in response to the original objective (Barro & Gordon, 1983a, b and Rogoff, 1985). This time inconsistency problem, if not addressed, leads to inflation bias that occurs when a government interferes with a central bank’s operation. In this context, if the central bank knows public inflation expectation, it tends to create inflation surprise to increase seignorage and to push the employment rate. The later could result in loss of credibility for the central bank that likely to hinder its ability to manage inflation expectation through its policy instruments? In a sense, this suggests that delegating monetary policy to an independent central bank is anticipated to promote economic agent’s trust in future monetary stability. Therefore an independent central bank is better positioned to eliminate the time inconsistency problem of monetary policy (Rogoff, 1985, and Bernanke, 2010). Empirically, however, the debate has shifted to whether it is central bank reforms that drive down inflation or the inflation experience that creates desires for countries to implement central bank reforms.

In the empirical literature, this debate on the causality between central bank independence and inflation has received and continues to attract considerable attention. Grilli et al. (1991) show that there is a significant negative relationship between CBI and inflation. Even after splitting the sample periods into four decades, the negative relationship seems to persist. This result suggests that countries with a low central bank independence index experience a higher inflation rate compared to their counterpart. Cukierman et al. (1992) and Alpanda &Honig (2014) argue that a substantial effect of inflation when a low degree of central bank independence, whereas less independent central banks, produce a higher rate of inflation.

Cukierman et al. (1992) estimated the CBI index for 72 countries using actual independence as the turnover rate of the central bank governor, and divided the sample into developing and developed countries for the period between 1950 and 1989. They found a
negative relationship between CBI and inflation only for advanced economies but did not find the same result for developing economies. They show that the role of the turnover rate of governor contributes significantly to the reduction of inflation for developing countries, but not for developed countries. Also, they found a two-way causality between inflation and central bank independence. In contrast, Alesina & Summers (1993) employed simple plots to capture the relationship between CBI and inflation for 16 developed countries over the sample period 1955-1988 and establish a negative relationship between these two variables. Similarly, Jonsson (1995), Cukierman et al. (1992), and Arnone & Romelli (2013) show that there is a negative relationship between CBI and inflation. After splitting the period into three different decades, Jonsson (1995) show that CBI became the most critical aspect of reducing inflation during the high inflation period (1972-1979) for their sample.

In terms of causality, Dreher et al. (2008) estimated the existence of a causality relationship between central bank independence and inflation for 137 countries over the sample period 1970-2004 and found that causality runs from inflation to turnover rate. Similarly, Ahsan et al. (2008) conducted a Granger causality test on central bank independence and inflation on Asia Pacific countries but did not find a two-way causality between the two variables. According to Landstrom (2013), there is no evidence from correlation statistic to determine whether a higher degree of central bank independence leads to low inflation, or whether the causality lies in the opposite direction. Cole (2018) show that the beneficial effects of forwarding guidance increase if a central bank pursues price-level targeting instead of inflation targeting.

The empirical literature is, therefore, inconclusive on the relationship and direction of causality between central bank independence and inflation. The significant limitations of the existing empirical literature arise from the estimation of structural VARs and co-integration to establish the causal relationship between central bank independence and inflation without accounting for structural breaks.

This study addresses this limitation by employing the structural break test proposed by Bai & Perron (2003), which is elaborated in Westerlund (2006) to examine the history of inflation in developing countries and the direction of causality. This paper evaluates the idea that the level of central bank independence explains the inflation moderation in these economies, as a higher index of independence implies that the central bank is focused more on attaining price stability. We identify the structural break in the inflation rate, which relates to the exact moment of each central bank's change in the law that caused an increase in the level of independence. Structural changes in inflation could also attribute to central bank reform, which changes the monetary policy and central bank objective, such as adopting inflation targeting. If central bank reforms have been successful, it expects that it will result in a structural change in inflation. Griffin (2011) shows the impact of structural changes in the inflation rate. However, their binary approach is limited in explaining the shift associated with the intercept, trend, or CBI reform regime. Our approach base on the LM co-integration test that is appropriate for multiple structural breaks, both the levels and trended regression.

We employ the legal index of central bank independence to estimates the structural breaks and investigate the causal relationship between central bank independence and
inflation. Further, we evaluate whether a higher degree of central bank independence leads to lower inflation or higher inflation leads to the central bank becoming more independent in developing countries by using Dumitrescu & Hurlin (2012) panel Granger non-causality test. The panel Granger non-causality test is suitable for heterogeneous panels since it assumes that there are identical lag orders and balanced panels for all cross-countries (Kumar 2011). Given that our panel data comprise of 37 developing countries, this technique is the most appropriate to show a causal relationship between central bank independence and inflation.

The rest of this paper is structured as follows. The next section provides the econometric methodology and data used for our empirical analysis. Section 3 presents the results and discussions. Section 4 concludes and provides areas for extensions.

Method

We use a panel dataset for a sample of 37 developing countries covering the period 1972 - 2016. Our choice of developing countries in the sample is based on the availability of consistent time series data on all the variables of interest. The variable consists of annual inflation and an index for central bank reforms. Inflation measures as an annual percentage change of the consumer price index that derives from the International Financial Statistic of the international monetary (IMF) database. Central bank reform measure using the legal CBI index that originally constructed by Cukierman et al. (1992). The index ranges between 0 and 1, where higher values denote a greater degree of central bank independence while the value of 0 indicates that the central bank is not independent. Our legal CBI index data construct from aggregate weighted taken from Garriga (2016) dataset, covering the period 1972-2012. We use the same weights to extend the legal CBI index to 2016 for countries in our sample. Our extension is tractable given that there have been no changes in the central bank’s law for the sample countries over the period.

The empirical model using Westerlund (2006) that based on a test co-integration, which allows for the presence of structural breaks in the deterministic variable of a co-integrating panel regression. It is Lagrange Multiplier (LM) for the null hypothesis of co-integration that allows for the possibility of multiple structural breaks in both the level and trend of a co-integrated panel regression. Westerlund (2006) indicates that the test is general enough to allow for endogenous regressors, serial correlation, and an unknown number of breaks that may locate at different dates for different individuals. The test can be conducted through two different approaches to check for the structural breaks in the data, namely: known and unknown breakpoints test statistics.

In this paper, we employ the test for unknown breakpoints based on the extension of Bai and Perron (1998, 2003) to establish the location of the breakpoints of inflation for each country. First, we consider the following panel regressions with m breaks (m+1 regimes):

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1 The countries in the sample are: Argentina, Barbados, Bolivia, Brazil, Chile, Colombia, Costa Rica, Egypt, Ethiopia, Ghana, Guatemala, Honduras, Indonesia, Kenya, Malaysia, Mauritania, Mexico, Morocco, Nepal, Nicaragua, Nigeria, Pakistan, Paraguay, Peru, Philippines, South Africa, Sri Lanka, Suriname, Tanzania, Thailand, Trinidad and Tobago, Tunisia, Turkey, Uganda, Uruguay, Venezuela, Zambia
Where $y_t$ is the observed dependent variable for country $i$ at time $t$, $x_t(p \times 1)$ and $z_t(p \times 1)$ are vector of covariates, and the corresponding vector of coefficients are $\beta$ and $\delta_j$ for $j = 1, \ldots, m + 1$. The structural break points $(T_1, \ldots, T_m)$ are explicitly unknown. The purpose is to estimate the unknown regression coefficients together with the break points when $T$ observations on $(y_t, x_t, z_t)$ are available. The variance of the error term $\varepsilon_t$ does not need to be constant. It is important to note that break in the variance of the error term is permitted, as long as it occurs at the same date with the parameters.

The estimation method base on the least-squares principle. In this approach, each partition, the associated least squares estimates of the parameters $\beta$ and $\delta_j$ are obtained by minimizing the sum of squared residuals based on the solution of following optimization problem:

$$
\hat{T}_i = \arg\min_{T_i} \sum_{j=1}^{m+1} \sum_{t=T_{i+1}}^{T_i} (y_{i,t} - z'_{i,t}\delta_j - x'_{i,t}\beta_t)^2
$$

Where $\hat{T}_i = (\hat{T}_1, \ldots, \hat{T}_m)$ denotes the vector of estimated structural breakpoints. The parameters $\delta_{it}$ and $\beta_{it}$ are cointegration estimator based on partition $T_i = (T_{i+1}, \ldots, T_m)$ of the structural breaks. Therefore the breakpoint estimators can be considered as global minimizers of the objective function. Given that breakpoints are discrete parameters and can only take a finite number of values, they can be estimated by a grid search.

The estimation is performed in two steps. First, Bai & Perron (2003) adopt a dynamic programming algorithm to estimate unknown regression parameters and unknown breakpoints $(T_i)$, and use the result to generate the sum of squared residuals for each break. The next step involves estimating the number of breaks using the sum of squared residuals from the previous step. The two steps are repeated $m$ times to obtain a vector of estimated breakpoints for each individual country in the sample. The LM statistic is estimated using $\hat{T}_i$ to substitute $T_i$ for each individual.

The panel Granger causality provides useful information related to temporality, erogeneity, and independence (Kumar, 2011). Temporality means that the previous values of a variable $x_{i,t}$ can cause another variable $y_{i,t}$. The standard method to examine causality is the Granger non-causality test for heterogeneous panel data models. We therefore closely follow the approach proposed by Dumitrescu & Hurlin (2012) that based on the individual country Wald statistics of Granger non-causality averaged across the cross-section units.

If we consider two covariance stationary variables, denoted $x_{i,t}$ and $y_{i,t}$, observed on $T$ periods and on $N$ individuals. As suggested by Dumitrescu & Hurlin (2012), for each individual country $i = 1, \ldots, N$, at time $t = 1, \ldots, T$, we consider the following linear model for testing the relationship between our two variables of interest:

$$
y_{i,t} = \alpha_i + \sum_{k=1}^{K} \gamma^{(k)}_{i,t-k} y_{i,t-k} + \sum_{k=1}^{K} \beta^{(k)}_{i,t} x_{i,t-k} + \varepsilon_{i,t}
$$

Where $K \in \mathbb{N}^*$ denotes the maximum number of lags that is identical for all cross-section units
of the panel and also assume the panel is balanced, \( \beta_i = (\beta_1^{(1)}, \ldots, \beta_k^{(1)})' \) is the regression coefficient of the slope of \( x_{i,t-k} \). \( t \) denotes the time period, the autoregressive parameter \( \gamma_i^{(k)} \) is coefficient of \( y_{i,t-k} \). The individual effect \( \alpha_i \) is supposed to be fixed over time. Initial conditions of the individual processes \( y_{i,t} \) and \( x_{i,t} \) are given and observable. The null hypothesis for the test is given by:

\[
H_0: \beta_1 = \ldots = \beta_K = 0 \quad \forall i = 1, ..., N
\]

Which corresponds to the absence of causality among the variables in the panel. The test assumes that there can be causality for some variables but not necessarily all. Therefore the alternative is given by:

\[
H_1: \beta_1 = \ldots = \beta_K = 0 \quad \forall i = 1, ..., N
\]

\[
\beta_1 \neq 0 \quad \text{or} \quad \beta_K \neq 0 \quad \forall i = N_1 + 1, ..., N
\]

Where \( N_1 \in [0, N - 1] \) is unknown, \( 0 < N_1/N < 1 \). If \( N_1 = N \), there is no causality for any individuals in the model but if \( N_1 = 0 \), there is causality for all individuals in the model. \( N_1 \) must be strictly smaller than \( N \), otherwise there is no causality for all individuals.

According to Lopez and Weber (2017), conducting the Dumitrescu and Hurlin (2012) Granger non-causality test involves first running the \( N \) individual regression based on equation (3); second, examine F-tests of the \( K \) linear hypotheses \( \beta_1 = \ldots = \beta_K = 0 \) to get the individual country Wald statistics, \( W_i \). In the third step we compute average Wald statistic.

\[
(W) \bar{W} = \frac{1}{N} \sum_{i=1}^{N} W_{i,T}
\]

(6)

Where: \( W_i \) denotes the standard adjusted Wald statistic for the \( i^{th} \) individual country observed during period \( T \). This test is designed to detect causality at the panel-level; therefore rejecting \( H_0 \) is not sufficient for one to conclude that there is no causality for some individuals.

**Results and Discussions**

Before estimation of the structural breaks and causality between the legal index of CBI and inflation, we first examined the measure of dispersion, central tendency, and volatility of the two variables. Table 1 indicates that the average inflation during the sample period was 62.06 percent, with a very high variability of up to 515.95 over the sample period. The high variability is mostly due to the difference in the inflation experience of these developing countries. The variability ranges between a deflation of 9.81 percent and a hyperinflation of up to 11,749.64 percent. The degree of central bank independence averaged 0.47, with a relatively lower variability signified by the standard deviation of 0.188 and dispersion between 0.13 and 0.95.

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Std. Deviation</th>
<th>Min</th>
<th>Max</th>
</tr>
</thead>
<tbody>
<tr>
<td>Inflation</td>
<td>62.057</td>
<td>515.954</td>
<td>-9.809</td>
<td>11,749.64</td>
</tr>
<tr>
<td>CBI</td>
<td>0.474</td>
<td>0.188</td>
<td>0.134</td>
<td>0.951</td>
</tr>
</tbody>
</table>
The large variability in inflation over the period suggests that some countries experienced very high inflation while others hand negative inflation before the adoption of central bank reforms. In such cases, the motivation to adopt central bank independence could have driven by the desire to reign-in on inflation and to exploit the inverse relationship established in Alesina & Summers (1993), Bodea & Hicks (2015), and Martin (2015) that countries with high CBI index experience low inflation.

We use the least square approach of Bai & Perron (2003) suggested by Westerlund (2006) to estimate the number of breaks and their location. Bai & Perron (2003) can be used to detect multiple structural breaks by assuming unknown structural breakpoints. This technique requires a specific number of breaks; hence we decided to apply only one break due to a short period covered by our time series data. The breakpoints for each country and the trend of inflation illustrate in Appendix I. Table 2 suggest that central bank reforms that changed their level of independence may have been responsible for the structural break in inflation. The long-run relationship between CBI and inflation could have caused by the existence of structural breaks in the data. These breaks in the inflation may be a result of reform in the monetary policy and the central bank’s objective or a change from external factors that can influence inflation, such as the rise in oil price.

Since central bank independence has implemented in many countries throughout the 1990s and early 2000s (Acemoglu et al., 2008), it is essential to evaluate the impact on inflation. The moderate and high inflation countries are grouped into two categories, comprising of countries that experience breaks before 1990 and those that had their break after 1990. For a moderate inflation group, seven countries (Barbados, Malaysia, Pakistan, Philippines, Thailand, Trinidad and Tobago, and Tunisia) had a break before 1990 where among the high inflation group, Bolivia, Chile, Ghana, Mexico, and Uganda experienced break before 1990.

The dates for those breaks correspond with the oil price fall at the beginning of the 1980s, which suggests that many countries could have shifted their focus to maintaining price stability. The results further indicate that the best countries in the moderate inflation group (Colombia, Costa Rica, Egypt, Guatemala, Honduras, Indonesia, Kenya, Morocco, Nepal, Paraguay, South Africa, and Tanzania) experience their break after the 1990s. Similarly, we find that the high inflation countries such as Argentina, Brazil, Nicaragua, Nigeria, Peru, Turkey, Uruguay, and Zambia had their structural breaks after the 1990s. The patterns in these countries characterized by an increase in the degree of the legal index of central bank independence. However, some countries such as Ethiopia, Mauritania, Sri Lanka, Suriname, and Venezuela had year-long break time after central bank independence has implemented. This condition could partly be a limitation in our approach that relies on only one structural break test.

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2 Although it is possible to apply multiple breaks, we only employ one break due to the fact that our time period is short on account of annual time series data. We also deviate from multiple breaks since it sometimes provide conflicting dates of break points.
Table 2. Structural Break Tests

<table>
<thead>
<tr>
<th>Country</th>
<th>Break date</th>
<th>CB Reform(s)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Argentina</td>
<td>1991</td>
<td>1992</td>
</tr>
<tr>
<td>Barbados</td>
<td>1983</td>
<td>1998</td>
</tr>
<tr>
<td>Bolivia</td>
<td>1986</td>
<td>1995</td>
</tr>
<tr>
<td>Brazil</td>
<td>1995</td>
<td>1988</td>
</tr>
<tr>
<td>Chile</td>
<td>1978</td>
<td>19,751,989</td>
</tr>
<tr>
<td>Colombia</td>
<td>1992</td>
<td>1999</td>
</tr>
<tr>
<td>Costarica</td>
<td>1996</td>
<td>1995</td>
</tr>
<tr>
<td>Egypt</td>
<td>1996</td>
<td>1975, 2004</td>
</tr>
<tr>
<td>Ethiopia</td>
<td>2005</td>
<td>1994</td>
</tr>
<tr>
<td>Ghana</td>
<td>1985</td>
<td>2002</td>
</tr>
<tr>
<td>Guatemala</td>
<td>1997</td>
<td>2002</td>
</tr>
<tr>
<td>Honduras</td>
<td>1998</td>
<td>1996</td>
</tr>
<tr>
<td>Indonesia</td>
<td>1999</td>
<td>1999</td>
</tr>
<tr>
<td>Kenya</td>
<td>1995</td>
<td>1996</td>
</tr>
<tr>
<td>Malaysia</td>
<td>1983</td>
<td>1994, 2009</td>
</tr>
<tr>
<td>Mauritania</td>
<td>2007</td>
<td>2007</td>
</tr>
<tr>
<td>Mexico</td>
<td>1989</td>
<td>1993</td>
</tr>
<tr>
<td>Morocco</td>
<td>1996</td>
<td>2006</td>
</tr>
<tr>
<td>Nepal</td>
<td>2000</td>
<td>2002</td>
</tr>
<tr>
<td>Nicaragua</td>
<td>1992</td>
<td>1992, 1999</td>
</tr>
<tr>
<td>Paraguay</td>
<td>1995</td>
<td>1995</td>
</tr>
<tr>
<td>Philippines</td>
<td>1986</td>
<td>1993</td>
</tr>
<tr>
<td>Peru</td>
<td>1992</td>
<td>1992</td>
</tr>
<tr>
<td>Thailand</td>
<td>1982</td>
<td>2008</td>
</tr>
<tr>
<td>Trinidad and Tobago</td>
<td>1985</td>
<td>1995</td>
</tr>
<tr>
<td>Tunisia</td>
<td>1988</td>
<td>1988, 2006</td>
</tr>
<tr>
<td>Turkey</td>
<td>2001</td>
<td>2001</td>
</tr>
<tr>
<td>Uganda</td>
<td>1990</td>
<td>1993, 2000</td>
</tr>
<tr>
<td>Uruguay</td>
<td>1995</td>
<td>1995, 2010</td>
</tr>
<tr>
<td>Zambia</td>
<td>1997</td>
<td>1996</td>
</tr>
</tbody>
</table>

Note: The breakpoints were estimated using the Bai and Perron (2003) procedure

To understand the implications of the structural break, we divided our samples into two categories comprising of countries that had their break before the reforms, and those had their break after the reforms. The result in Table 2 reveals that 20 countries had a break before, while 17 countries had their structural break after the central bank reforms.
Among the countries with moderate inflation, 12 countries had their structural breaks before reforms, namely Barbados, Colombia, Egypt, Guatemala, Kenya, Malaysia, Nepal, Pakistan, Philippines, South Africa, Thailand, and Trinidad and Tobago. In contrast, ten developing countries, among which includes Costa Rica, Ethiopia, Honduras, Indonesia, Mauritania, Morocco, Paraguay, Sri Lanka, Tanzania, and Tunisia, experience their structural breaks after central bank reforms. In the case of the high inflation group, seven countries, including Chile, Nicaragua, Peru, Turkey, Uruguay, Venezuela, and Zambia, encountered break after reforms.

Meanwhile, eight countries, namely Argentina, Bolivia, Brazil, Ghana, Mexico, Nigeria, Suriname, and Uganda, had a break before central bank reforms. Based on this evidence, we can conclude that central bank reform is associated with inflation. However, we have to examine more deeply whether reform causes inflation, or it is the inflation experience that drives the desire for the reform.

We run the panel Granger non-causality test for heterogeneous panel data models following the approach of Dumitrescu & Hurlin (2012). We also utilize the group mean Wald test statistic, to establish the causality relationship between inflation and the legal index of central bank independence. The advantage of the Granger non-causality test is that it can be employed even in models that possibly integrated and in co-integrated systems without pre-testing for unit roots and co-integration. We set the number of lags in this test is two.

Table 3 present the findings from the estimation of the two null hypotheses. The first hypothesis is that CBI does not homogeneously cause inflation, and the other hypothesis is that inflation does not homogeneously cause CBI. The result of our estimation indicates that both hypotheses rejected. This result consistent with Dumitrescu & Hurlin (2012), that also find a bidirectional causality between CBI and inflation. The causality runs from the legal index of CBI to inflation, and the reverse from inflation to CBI is also true. Our empirical finding is also consistent with Cukierman et al. (1992), and Brumm (2011) that finds two-way causality between CBI and inflation.

Table 3. Panel Non-Causality Test

<table>
<thead>
<tr>
<th>Null Hypothesis</th>
<th>Wald Statistics</th>
<th>Z Statistics</th>
<th>Probability values</th>
</tr>
</thead>
<tbody>
<tr>
<td>CBI does not cause Inflation</td>
<td>4.515</td>
<td>6.552</td>
<td>0.000***</td>
</tr>
<tr>
<td>Inflation does not cause CBI</td>
<td>77.938</td>
<td>206.662</td>
<td>0.000***</td>
</tr>
<tr>
<td>Moderate Inflation Countries</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>CBI does not cause Inflation</td>
<td>4.746</td>
<td>5.537</td>
<td>0.000***</td>
</tr>
<tr>
<td>Inflation does not cause CBI</td>
<td>4.813</td>
<td>5.678</td>
<td>0.000***</td>
</tr>
<tr>
<td>High Inflation Countries</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>CBI does not cause Inflation</td>
<td>4.176</td>
<td>3.584</td>
<td>0.000***</td>
</tr>
<tr>
<td>Inflation does not cause CBI</td>
<td>185.189</td>
<td>317.700</td>
<td>0.000***</td>
</tr>
</tbody>
</table>

Note: Computed from sample data (1972-2016). *** denotes rejection at 1% significance level

This result suggests that inflation variability is high among central banks that have not achieved the reforms required for CBI compared to those that have achieved a high level of
independence. CBI is more effective in lowering inflation in the presence of high levels of banking sector development and institutional quality (Arnone, 2013; Agoba et al., 2017). However, Cukierman et al. (2002) also find that during the early stages of liberalization, the CBI is unrelated to inflation. Perera et al. (2013), and Posso & Tawadros (2013) show that a negative relationship between central bank financial strength and inflation. CBI increases inflation in lowest-income countries (Ftiti et al., 2017).

Our results still indicate bi-directional causality between CBI and inflation by Splitting the sample into high inflation and low, moderate inflation countries. Although the Wald statistics are very high for the causality from inflation to CBI for the high inflation countries, it is significant. This empirical finding is consistent with the results and implication of structural breaks on the inflation dynamics of developing countries discussed in the previous subsection. The subsection provided evidence, which suggests that some countries that experienced breakpoint date after reforms experienced low and stable inflation. On the other hand, countries that had inflation structural break dates before their central bank reforms could have a drive into the adoption of CBI reforms to maintain low and stable inflation.

Our results give credence to central bank independence as a crucial requirement for macroeconomic stability and suggest that CBI and inflation drive one another. The choice of the indicator used to measure the level of central bank independence may influence the results. Using indicators such as political independence, financial independence, and turnover of governors may produce different results from those established in this study.

Conclusion

This study provided evidence on the relationship and causality between central bank reforms and inflation using annual panel data from a sample of 37 developing countries over the period 1972 to 2016. The significant contribution of this paper was the adoption of structural break test based on the estimation approach of Bai & Perron (2003), and the Granger non-causality tests proposed by Dumitrescu & Hurlin (2012) to capture the breaks and establish the direction of causality on panel data. The empirical results indicate that a high degree of central bank independence leads to low and stable inflation, implying that central bank reforms are crucial for attaining macroeconomic stability. In particular, we found that some countries experienced a break date before the 1990s, which can partly be attributed to the reversal of the oil price shock and drives towards central bank policy autonomy. The rest of the countries' breaks came after the 1990s, at the time, when many countries had begun to devolve powers of policy independence to central banks. On this, our result shows that a total of 20 countries had a break before their central bank reforms, and 17 countries had a break after the reform, which suggests that for some central banks, reforms were a driver for the structural breaks. The results from the Granger non-causality tests indicate a bidirectional causality relationship between CBI and inflation. This result suggests that high inflation may have forced some countries to implement central bank reforms, and in most cases, increased central bank independence led to lower inflation.

The empirical finding is consistent with the theoretical concept that central bank
reforms lead to low and stable inflation. Establishing the direction of causality is essential in ensuring the effectiveness of the reforms. In terms of policy implications, policymakers could focus on increasing the degree of independence through a sound legal framework. Measures to strengthen the capacity of the central banks to implement monetary action should harness to foster sustained macroeconomic stability and long-run economic growth.

Our finding only utilized the legal index of central bank independence. Given that the other measures, namely index based on the turnover rate of central bank governors, and the index based on governance are essential in gauging the level of independence, it is vital to explore whether these indices also indicate that central bank reforms lower and stabilizes inflation. Another interesting question is whether the effect of CBI on inflation is permanent or transitory.

References


Appendix I. Break Point for Central Bank Independence and Annual Inflation