9. THE LONG-RUN AND SHORT-RUN EXCHANGE RATE PASS-THROUGH DURING THE PERIOD OF ECONOMIC REFORMS IN NIGERIA: IS IT COMPLETE OR INCOMPLETE?

Mehmet BALCILAR[1]
Ojonugwa USMAN[2]
Muhammad Sani MUSA[3]

Abstract

This study investigates the long-run and short-run exchange rate pass-through (ERPT) to inflation in Nigeria during the period of economic reforms by incorporating structural breaks over the period 2000Q1-2017Q4. We applied the minimum Lagrange multiplier unit root test with a structural break, the Bayer-Hanck combined cointegration test and the estimation procedure based on the autoregressive distributed lag (ARDL) model. The empirical results suggest that the long-run and short-run ERPT during this period is incomplete with the long-run pass-through having a stronger effect. This confirms the monetary model of short-run price stickiness. The results further divulge that output growth stimulates inflation while oil price dampens inflation in the long run and short run respectively. The effects of trade openness on inflation both in the long run and short run are positive, which cast doubt on the validity of Romer's (1993) hypothesis. The deviation from the long-run equilibrium level is corrected by 47.6% speed of adjustment in every quarter. These findings, therefore, provide insights for the policymakers to understand the form and scale of ERPT and use its estimated size as a tool to attain low level of inflation and price stability in Nigeria.

Keywords: ERPT; inflation; combined cointegration; ARDL model; Romer's hypothesis; Nigeria

JEL Classification: E31, F31, F37

[1] *Eastern Mediterranean University, Northern Cyprus, via Mersin 10, Turkey. E-mail: mehmet@mbalcilar.net*

[2] *Eastern Mediterranean University, Northern Cyprus, via Mersin 10, Turkey and School of Business Education, Federal College of Education (Technical), Potiskum – Nigeria. Corresponding author. E-mail: usmanojonugwa@gmail.com.*

[3] *Central Administration, Office of the Registrar, Kaduna Polytechnic, Kaduna State, Nigeria. E-mail: msmbiu2007@gmail.com*

1. Introduction

The quest to achieve structural transformations of the economy since Nigeria's political independence in 1960 has remained the key interest of the government and the policymakers. Despite the oscillations in the price of crude oil over the years, the economy of Nigeria has continued to depend on crude oil exports as the main source of foreign exchange revenues. The overdependence on crude oil from all intents and purposes have adverse effects on the macroeconomic stability. As correctly noted by Lariau *et al.* (2016), whenever crude oil price drops significantly, Nigeria's currency tends to face the pressure of depreciation, which invariably transmits into domestic inflation through the direct and indirect channels of exchange rate pass-through (ERPT).[4] For example, Poloamina *et al.* (2009) note that between 1985 and 1993, the depreciation of naira was by about 71% on average, annually based on the official exchange rate, and by 114% in the parallel exchange rate market. In 2009, the nominal exchange rate depreciated by 25.61%. The depreciation in the nominal exchange rate in 2015 was 21.37% while in 2016 was roughly 31.72% until the government introduced a free-floating system in June 2016, after which naira became relatively stable. For the case of inflation, Poloamina *et al.* (2009) and Oyinlola and Babatunde (2009) document that it increased annually by about 78% between 1985 and 1987. In 1988, the annual change in inflation rose to about 300% as compared to the previous year. The country experienced the peak of domestic inflation of 79.9% in 1995. This declined to 6.9% in 2000, 17.8% in 2005, 5.4% in 2007 and 11.40% in May 2019.

In response to the incidence of inflation in Nigeria, the Central Bank of Nigeria (CBN) had to embark on various economic reforms since the inception of a democratic government in 1999; hence, the monetary policy framework switched from the exchange rate targeting to inflation targeting[5] (Bamidele, 2007; Aliyu and Englama, 2009). Understandably, the idea of inflation targeting in Nigeria was conceived in 2000 via a paper presented by Uchendu (2000). This paper drew the attention of the CBN to the efficacy of inflation targeting as an instrument of macroeconomic management. Consequently, by 2007 the CBN Act was enacted and the operational autonomy and other issues related to the independence of the bank were given full consideration. According to Savensson (1999), within the framework of inflation targeting, monetary policy goals are stated with well-defined tasks, which are clear and unambiguous.

At the advent of the Structural Adjustment Program (SAP) in the mid-1986, Nigeria had to shift its economic policies and reforms towards a free market economy. This was more intensified after the return of democracy in 1999. While empirical evidence on the pass-through of exchange rate has burgeoned in Nigeria, there is a contentious question as to whether the shift in the system of exchange rate and trade policy towards market-based economy exerts pressure on the pass-through channels during the period of economic reforms. This period is characterized by not only market-based economic reforms but also

---

[4] *The direct channel of ERPT occurs when the effect of exchange rate movements on import prices is transmitted to the prices of producer and consumer goods. Whereas, the indirect channel of ERPT occurs due to the high demand for substitute goods which causes exports to be more competitive.*

[5] *Inflation targeting is a monetary policy framework whereby the CBN announces the target of inflation over a given period of time to the public.*

by the inflation-targeting regime of the CBN.[6] Theoretically, as the economy becomes more globally connected, one would expect it to be vulnerable to the effects of exchange rate fluctuations. For this reason, this study contributes to the literature on the following grounds: first, we estimate the dynamic pass-through of the exchange rate to inflation by controlling for output growth, oil price, and degree of openness during the period of economic reforms in Nigeria. The period under investigation is portrayed by a relatively low and stable level of inflation. Second, we apply a minimum Lagrange Multiplier (LM) unit root test with a structural break proposed by Lee and Strazicich (2013). This test does not only accommodate information regarding a single unknown break but solves the problem of inaccurate estimation of the break-point when there is break under the null and alternative hypotheses, as found in Zivot and Andrews (1992) unit root test with a structural break. Third, we test for the existence of a long-run relationship by using a combined cointegration test recently developed by Bayer and Hanck (2013). This cointegration test helps to obtain a combination of various individual cointegration test results and provide a more conclusive finding, devoid of any conflicting results that may lead to an arbitrary decision. Fourth, we accommodate structural breaks in the cointegration test by applying the bounds testing method proposed by Pesaran *et al.* (2001) as a robustness test. Fifth, the dynamic long-run and short-run error correction parameters are obtained based on a flexible autoregressive distributed lag (ARDL) model in the presence of structural breaks.

The rest of the paper is structured as follows: Section 2 reviews both the theoretical and empirical literature. Section 3 discusses the data and methodology of this paper. Section 4 provides the empirical results and discussion, while Section 5 presents the concluding remarks and policy implications.

2. Theoretical Framework and Empirical Literature

The theoretical framework for the ERPT evolves from the theory of the purchasing power parity (PPP), an offshoot of law of one price (LOP) which states clearly that at equilibrium, the market prices of tradable goods and services are the same in different countries if their prices are measured in the same unit of currency. The main force behind this theory is the perfectly competitive arbitrage activities, which propel the exchange rate to adjust seemingly to the equilibrium level where the PPP holds. Therefore, following Rogoff (1996), the theory of the PPP*,* with the assumptions of no existence of transportation costs, tariffs, imperfect competition, and other trade barriers is specified as:

$$P_{i,t} = ERT_t \times P_{i,t}^* \qquad (1)$$

where: $P_{i,t}$ is the domestic price for good i in period t, $P_{i,t}^*$ is the foreign price for good i in period t, and ERT_t is the nominal exchange rate in period t. Notably, equation (1) is the

---

[6] *Economic reforms introduced after the return of democracy include resuscitation of privatization and commercialization policy in 1999; National Eradication Program (NAPEP) in 2001, National Economic Empowerment and Development Strategy (NEEDS) in 2004, 7-Point Agenda in 2007, etc. In the foreign exchange market, the main policies include Dutch Auction System (RDAS) in 2006 that metamorphose into Retail Dutch Auction System (RDAS) and Wholesale Dutch Auction System (WDAS) in 2009; Interbank/Bureau-de-change market in 2015 and free-floating system in June 2016.*

absolute form of the PPP, which is the generalization of the law of one price. It demonstrates that given the same unit of currency, a basket of goods will cost the same in any country. Thus;

$$ERT_t = \frac{P_{i,t}}{P_{i,t}^*} \qquad (2)$$

where: $P_{i,t}$, $P_{i,t}^*$ and ER_t remain as previously defined. The exchange rate between two currencies is equal to the ratio of price levels in these countries. However, as suggested by the empirical literature, the ERPT is partial and incomplete, particularly in the short run. This means that the PPP and/or LOP does not hold either in absolute version or relative version.[7] This breakdown of the PPP or LOP is based on the nominal price stickiness arising from the weak competitive arbitrage activities, which remains the central debates between new Keynesian and new classical economists (Rogoff 1996).

Given that Nigeria is the largest net oil-exporting country in Africa, we assume that oil price has a significant impact on the consumer price inflation. This assumption implies that an increase in oil price may lead to a decrease in import bills (cost of inputs), which stimulates production, employment, and reduce cost-push inflation. Similarly, this could improve the balance of payments and demand-pull inflation.

While there is growing interest in analyzing the extent of the ERPT at country-specific and panel settings particularly in the open economies, the results of several studies have provided mixed pass-through elasticities in the short run and the long run. Clearly, the focus of the early studies on ERPT is centered on estimating the size of pass-through to domestic prices. The results abound in the literature and show that the ERPT is incomplete (see Menon, 1994, 1995; Kenny and McGettigan, 1996; Corsetti and Pesenti, 2005; Gagnon and Ihrig, 2004). In recent times, several studies have argued that the size of the pass-through has significantly declined, especially in the industrialized countries. For example, in a study by Otani *et al.* (2003), it was revealed that within the Japanese economy there was a huge decline in the pass-through to imports. This decline is mainly accounted for by the global decline in inflation and the promotion of intra-firm imports. In the same vein, Marazzi *et al.* (2005) report a significant decline in the ERPT to aggregate imports for the United States. Most notably, the result shows that the pass-through falls from 65% in the 1980s to somewhat 12% by the end of a decade in 2004. This result, therefore, echoes the earlier finding by Mann (1986). More so, Campa and Goldberg (2005) in their study document a similar finding that the 1990s decline in the ERPT of the Organization for Economic Cooperation and Development (OECD) countries is largely caused by the changing commodity composition of trade rather than the inflationary environment.

In a cross-sectional setting, Pollard and Coughlin (2004) use data for 30 industries to estimate the pass-through of the exchange rate to US import prices. The result, however, provides strong evidence that the reaction of these industries to changes in the exchange rate is incomplete. Cunningham *et al.* (2017) examine the time variation in the exchange rate pass-through to import prices in 24 advanced countries over the period 1995–2001. The finding suggests that the pass-through is heterogeneous and incomplete across the countries. Using quarterly data between 2000 and 2014 for seven economies in Southeast Europe, Kurtovic *et al.* (2018) demonstrate that the pass-through of the exchange rate to

---

[7] *Relative PPP suggests that a change in the price level is related to a change in the exchange rate. Hence, an economy with a relatively higher inflation rate tends to experience a depreciation of the currency.*

import prices is incomplete. This finding, in totality, refutes the claim that the size of the pass-through has been reduced over time. The result of the study further reveals that the pass-through in the transitional countries is higher than in the developed countries. Moreover, in the Hungarian economy, Hajnal *et al.* (2015) investigate the ERPT into consumer prices. The result carefully discloses that the size of the ERPT varies over time. The size before the crisis stands at 0.3%, while after the crisis the size falls between 0.1 and 0.2%.

Furthermore, one may find a significant amount of literature that relates the decline in ERPT to low and stable inflation environment, particularly in the industrialized countries (see Taylor, 2000; Bailliu and Fujii, 2004; Campa and Goldberg, 2005; Choudhri and Hakura, 2006; Junttila and Korhonen, 2012). McCarthy (2007) uses impulse responses and variance decomposition within the framework of vector autoregression (VAR) to estimate the ERPT for the economies of industrialized countries. The results reveal that the disinflationary effect identified by the study is attributed to external factors, which occurred during the past couple of years. This finding also holds during the post-1982 period. Choudhri and Hakura (2006) using a dataset for 71 countries over the period 1979-2000 find that the pass-through is related to the average inflation rate across these countries. On the contrary, Campa and Linda (2002) investigate the ERPT to import prices for 25 OECD countries over the period 1975 to 1999 with quarterly data. The empirical results indicate that the periods of high rate of inflation and high volatility of the exchange rate have a weak correlation with high ERPT. Most recently, Ben Cheikh and Louhichi (2016) use a large panel of 63 countries for the period 1992 to 2012 in order to revisit the association of inflation environment and ERPT. Their finding indicates strong evidence of a positive relationship between inflation environment and the ERPT. This finding validates the Taylor's (2000) hypothesis that the responsiveness of inflation to exchange rate changes depends positively on the inflation rate. More so, Aleem and Lahiani (2014), Dube (2016) and Kabundi and Mlachila (2018) show that low pass-through of the exchange rate is traceable to the effectiveness of monetary policy and the introduction of inflation targeting regimes.

In Nigeria's case, it is clear that the available empirical literature provides different magnitudes of the long-run and short-run ERPT. However, most of the studies show that the pass-through is lower than unit especially in the short to medium term. For example, Oladipo (2007) uses a markup approach to estimate the pass-through of the exchange rate to imports for Nigeria between 1970Q1 to 2004Q4. His major finding supports the incomplete pass-through across sectors. This implies that the foreign exporters perhaps pass-through only part of the increase in their costs of production resulting from a change in the exchange rate to import prices. Aliyu *et al.* (2009) and Oyinlola and Egwaikhide (2011) examine the extent of the ERPT to import and consumer prices in Nigeria based on the Johansen cointegration and vector error correction model (VECM). The results indicate the existence of a long-run relationship between exchange rate and domestic prices, but the pass-through elasticity is incomplete in the two studies. Essien (2005) and Adeyemi and Samuel (2013) equally find incomplete pass-through of the exchange rate in Nigeria with the pass-through increasing towards the long-run. This finding is contrary to Zubair *et al.* (2013), who estimated the long-run pass-through to be as low as 0.2 in Nigeria. Contradicting the above studies, Omisakin (2009) reports that there is no evidence to support the pass-through of the exchange rate to inflation and growth in Nigeria, both in the short run and long run. This finding is partly supported by Poloamina *et al.* (2009), who only document no substantial evidence of ERPT in the long run, while in the short run the ERPT is complete, *i.e.* a 1% depreciation in the exchange rate leads to a 1% increase in import prices. Furthermore, recently, Usman and Musa (2018) revisited the pass-through of the exchange rate in Nigeria using a cointegrated

VAR approach. The results clearly reveal that exchange rate, import price index and openness index are the major determinants of inflation in Nigeria. Similarly, Balcilar *et al.* (2019) revisit exchange rate pass-through in Nigeria and South Africa considering the effect of structural breaks. The empirical results suggest a complete long-run pass-through in Nigeria and incomplete long-run pass-through in South Africa. This difference in the pass-through confirms the importance of inflation targeting monetary policy in stabilizing the domestic prices.

The inverse relationship between inflation and openness of the economy has gained empirical support after Romer's (1993) hypothesis. The argument in support of Romer's hypothesis is that exporting firms always find it difficult to fully pass-through the exchange rate shocks to their export prices in the presence of an intensified worldwide competitive pressure emanating from trade openness. In supporting this hypothesis, Jin (2000) finds that the degree of openness impacts inflation negatively through its effect on output resulting from efficiency in resource allocations, improvement in capacity utilization and increase in investments. Studies like Hanif and Batool (2006), Lane and Milesi-Ferretti (2007), Wynne and Kersting (2007), Granato *et al.* (2007), and Mukhtar (2012) all provide evidence in support of Romer's hypothesis. On the contrary, Aliyu *et al.* (2009), Samimi *et al.* (2012), and Bada *et al.* (2016) maintain that the inverse relationship between trade openness and inflation hypothesised by Romer (1993) is mostly refuted, especially in the developing and emerging market economies, where competition is weak and markets are imperfect.

Relative to previous studies on the ERPT, this is the first study, to the best of our knowledge, that measures the pass-through of exchange rate to inflation during the period of economic reforms in Nigeria, by incorporating structural breaks. The finding of this study is expected to serve as a tool for deciding appropriate monetary policy rules to achieve low inflation levels and price stability in Nigeria.

3. Data and Methodology

3.1 Data

This study uses economic variables, which include consumer price index (CPI), exchange rate, output, oil price, and trade openness. The quarterly data on the nominal effective exchange rate (NERT) and consumer price index are obtained from the International Financial Statistics database, in order to measure exchange rate and inflation. The nominal effective exchange rate is used because it is broader with more variations, which may lead to robust results. The NERT is defined as the trade-weighted value of a basket of foreign currencies per unit of naira; thus, an increase means appreciation of naira. The output is measured as the gross domestic product (GDP) (constant 2010 US$), the oil price (OILP) corresponds to the international Brent crude oil spot price in US$ per barrel, while trade openness (TOPN) corresponds to trade volume (imports plus exports) as a percentage of GDP. The data for output and trade openness are not found on a quarterly basis. Hence, following Balcilar *et al.* (2019), we collected the data from the World Bank via World Development Indicator annually and converted to quarterly series using the quadratic interpolation method.[8] The data on Brent crude oil spot price oil price is obtained from the Datastream. One should notice that the output measure and CPI are seasonally adjusted.

---

[8] *Quadratic interpolation method is one of the approaches of converting series from low frequency to high frequency.*

Apart from trade openness, all the variables are expressed in their natural logarithm forms.[9] Finally, the period for this study spans from 2000Q1 to 2017Q4. This period coincides with the era of intensive economic reforms, particularly in the exchange rate and trade policies following the return of democracy on May 29, 1999. It also coincides with the inflation targeting monetary policy of the CBN, which witnesses a relatively low and stable inflation rate in Nigeria.

3.2 The Lee-Strazicich Unit Root Test

Many unit root tests were used in the literature on applied economics and finance. Notable among them, we include the augmented Dickey-Fuller (ADF) test proposed by Dickey and Fuller, (1979), the Phillips-Perron (PP) test proposed by Phillips and Perron (1988), the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test developed by Kwiatkowski *et al.* (1992), the Ng-Perron test proposed by Ng and Perron, (2001) and the generalized least squares Dickey-Fuller (DF-GLS) test developed by Elliot *et al.* (1996). As argued in the recent literature, these tests may provide spurious results if structural breaks are found in the series. The fact that structural breaks are usually a common feature of time series data, we account for structural break in this study by applying a minimum LM unit root test with one break proposed by Lee and Strazicich (2013). This test does not only accommodate information regarding a single unknown break, but addresses the problem of inaccurate estimation of breakpoint when there is break under the null and alternative hypotheses. Therefore, the test is superior to Zivot and Andrews (1992) unit root test with a structural break, which apparently depends on the size of the break under the null and suffers spurious rejections that increase as the break size increases.

The minimum LM unit root test by Lee and Strazicich (2013) is based on the following data generating process (DGP):

$$y_t = \delta' Z_t + X_t, \quad X_t = \beta X_{t-1} + \epsilon_t \tag{3}$$

where: Z_t represents explanatory variables and β=1 represents the null hypothesis of a unit root. $\delta' = [\delta_1, \delta_2, \delta_3]'$ are the regression coefficients. If $Z_t = [1, t]'$, the DGP is analogous to the no break LM unit root test developed by Schmidt and Philips (1992).

In testing for unit root with a break, Lee and Strazicich (2013) proposed three models – Model A, Model B, and Model C. Model A also known as 'crash' is described by $Z_t = [1, t, D_t]'$, where D_t=1 for $t \geq T_B + 1$ and 0 if otherwise. T_B denotes the possible breakpoint in the series This model allows for a one-time change in intercept under the alternative hypothesis. Model C described by $Z_t = [1, t, DT_t]'$, allows for a shift in intercept and a change in trend slope under the alternative hypothesis, where: $DT_t = t - T_B$ for $t \geq T_B + 1$ and 0 if otherwise. Model B, which is basically 'changing growth', is described by $Z_t = [1, t, DT_t^*]'$, where: $DT_t^* = t$ for $t \geq T_B + 1$ and 0 if otherwise. As noted by Lee and Strazicich (2013), Model B is excluded in the estimation process, since the economic and financial time series are mostly described by Model A or Model C. Therefore, to obtain the LM unit root statistic, the following regression equation is considered:

$$\Delta y_t = \delta' \Delta Z_t + \gamma \tilde{S}_{t-1} + \epsilon_t \tag{4}$$

---

[9] *Trade openness is already in percentage. Generally, log of a variable measured in percentage is not preferred in empirical studies when other variables are in log levels.*

Here, $\tilde{S}_t = y_t - \tilde{\psi}_x - Z_t\tilde{\delta}, \quad t = 2, \dots, T;$ $\tilde{\delta}$ represents the estimated regression coefficient. The $[1, B_t]'$ in Model A remarkably describes ΔZ_t which is rather included in Eq. 4 instead of Z_t, while $[1, \beta_t, D_t]'$ in Model C describes ΔZ_t, where B_t is equal to ΔD_t and D_t equals to ΔDT_t. A change in intercept and trend as well as one period change in drift under null alternative hypothesis all correspond to B_t and D_t. Also, the $\Delta \tilde{S}_{t-j}$, $j = 1, 2, \dots, K$, are included in Eq. 4 as additional regressors to correct for serial correlations.

To determine the optimal number of lags, K, Lee and Strazicich (2013) suggest the general to specific approach as described in Perron (1989) by starting from maximum number of lagged first-differenced terms $K = 8$ and continue to drop the maximum lagged term if the model is not statistically different from zero at 1%, 5%, or 10% significance level. The null hypothesis $H_0: \gamma = 0$ is that a unit root exists with a structural break while the alternative hypothesis is $H_0: \gamma < 0$ which implies that no unit root exists with a structural break.

3.3 The Bayer and Hanck Cointegration Test

To check for cointegrating relationship among the variables in this study, we apply Bayer-Hanck combined cointegration approach. The cointegration tests of the Engle and Granger (1987) residual-based, Boswijk (1994) error correction model-based F-test, Johansen (1995) system-based, and Banerjee *et al*. (1998) error correction model-based t-test might produce conflicting results, leading to an arbitrary decision with respect to the appropriate individual test to apply. To avoid this scenario, Bayer and Hanck (2013) proposed a suitable cointegration test that combines four major individual tests by Engle and Granger (1987), Boswijk (1994), Johansen (1995), and Banerjee *et al*. (1998) and obtain uniformly most powerful cointegration results, which are more robust, efficient, and reliable as compared to individual tests. Therefore, to combine the computed *p*-values of these individual cointegration tests, the famous Fisher test described in Fisher (1932) is used. The test is applied based on the following equations:

$$EG - JOH = -2\big[\ln(P_{EG}) + \ln(P_{JOH})\big], \tag{5}$$
$$EG - JOH - B0 - BDM = -2\big[\ln(P_{EG}) + \ln(P_{JOH}) + \ln(P_{BO}) + \ln(P_{BDM})\big], \tag{6}$$

where: P_{EG}, P_{JOH}, P_{BO}, and P_{BDM} are the probability values (*p*-values) of the cointegration tests of Engle and Granger (1987), Johansen (1995), Boswijk (1994) and Banerjee *et al*. (1998), respectively. The decision of the test as to whether there is a valid cointegration between the variables is taken based on the Fisher statistic. Particularly, the null hypothesis of no cointegration between the variables is rejected if the critical value of Bayer and Hanck (2013) is lower than the estimated Fisher statistic.

3.4 The ARDL Bounds Testing Approach

This study applies a flexible ARDL model proposed by Pesaran *et al*. (2001) to estimate the long-run and short-run pass-through of the exchange rate to consumer price inflation in Nigeria. This model does not impose any restrictive assumption that the time series must be integrated of the same order; hence, it creates greater flexibility. Also, the model performs better in testing for the long-run relationship in the small sample as compared to other methods. Therefore, the ARDL model based on the Unrestricted Error Correction Model (UECM) is expressed as:

$$lncpi_t = \rho_0 + \sum_{i=1}^{q} \rho_i \, lncpi_{t-i} + \sum_{i=0}^{p_1} \lambda_{1,i} \, lnert_{t-i} + \sum_{i=0}^{p_2} \lambda_{2,i} \, lngdp_{t-i}$$
$$+ \sum_{i=0}^{p_3} \lambda_{3,i} \, lnoilp_{t-i} + \sum_{i=0}^{p_4} \lambda_{4,i} \, topn_{t-i} + \varepsilon_t \tag{7}$$

where: $lncpi$, $lnert$, $lngdp$, $lnoilp$ and $topn$ represent the log of CPI, exchange rate, output measured by the real GDP and oil price, while $topn$ represents trade openness measured by the ratio of the trade volume (exports plus imports) to GDP. ε_t represents zero-mean white noise process with variance σ^2, $\varepsilon_t \sim iid(0, \sigma^2)$. If there is evidence of a valid long-run relationship between $lncpi$, $lnert$, $lngdp$, $lnoilp$ and $topn$, they will maintain a level relationship specified with long-run coefficients, which can be expressed on the basis of the error-correction model (ECM). To this end, we obtain the long-run parameters by estimating the following equation:

$$lncpi_t = \gamma_0 + \gamma_1 lnert_t + \gamma_2 lngdp_t + \gamma_3 lnoilp_t + \gamma_4 topn_t + \sum_{i=1}^{q} \varphi_i \, \Delta lncpi_{t-i}$$
$$+ \sum_{i=1}^{p_1-1} \sigma_{1,i} \, \Delta lnert_{t-i} + \sum_{i=1}^{p_2-1} \sigma_{2,i} \, \Delta lngdp_{t-i} + \sum_{i=1}^{p_3-1} \sigma_{3,i} \, \Delta lnoilp_{t-i} + \tag{8}$$
$$\sum_{i=1}^{p_4-1} \sigma_{4,i} \, \Delta topn_{t-i} + \varepsilon_t$$

In Eq. (8), Δ represents the first difference operator, which is expressed as $\Delta x_t = x_t - x_{t-1}$. The long-run coefficients can be obtained as $\alpha_i = \gamma_i / (1 - \sum_{j=1}^{q} \varphi_j)$, $i = 1,2,3,4$. One may notice that since all variables are expressed in their natural logarithms, the long-run coefficients, therefore, represent the long-run elasticities.[10] Moreover, the error-correction term (ECT) may be defined as: $ect_t = lncpi_t - \alpha_1 lnert_t - \alpha_2 lngdp_t - \alpha_3 lnoilpr_t - \alpha_4 topn_t$. The α_1, α_2, α_3, and α_4 coefficients denote the long-run estimates of the exchange rate effect on consumer price inflation, also known as ERPT, long-run output effect on consumer price inflation, long-run oil price effect on consumer price inflation, and long-run trade openness effect on consumer price inflation. Furthermore, as noted by the doctrine of purchasing power parity (PPP) the pass-through of the exchange rate to prices is presumably complete or full especially in the long run. This implies invariably that $\alpha_1 = -1$. However, the dependent variable, *i.e.* the change in CPI may not adjust to the long run equilibrium path instantaneously in the case of any change in the explanatory variables. The following error-correction model captures the adjustment speed from the short-run to the long-run equilibrium level:

---

[10] *The variable of trade openness is not expressed in its natural logarithm because it is already expressed as a percentage of GDP.*

$$\Delta lncpi_t = \tau_0 + \sum_{i=1}^{q} \theta_i \, \Delta lncpi_{t-i} + \sum_{i=0}^{p_1} \psi_{1,i} \, \Delta lnert_{t-i} + \sum_{i=0}^{p_2} \psi_{2,i} \, \Delta lngdp_{t-i}$$

$$+ \sum_{i=0}^{p_3} \psi_{3,i} \, \Delta lnoilpr_{t-1} + \sum_{i=0}^{p_4} \psi_{4,i} \, \Delta topn_{t-i} + \lambda ect_{t-1} + \varepsilon_t \tag{9}$$

where: ect_t is the error-correction term, obtained from the Eq. (9) and λ is the coefficient that measures the speed of adjustment to the equilibrium. The θ_i, $\psi_{1,i}$, $\psi_{2,i}$, $\psi_{3,i}$ and $\psi_{4,i}$ coefficients are the short-run coefficients representing the consumer price inflation inertia, short-run exchange rate effect on consumer price inflation also known as the ERPT, short-run output effect on consumer price inflation, short-run oil price effect on consumer price inflation, and short-run effect of trade openness on consumer price inflation.

To determine whether the pass-through of exchange rate to consumer price inflation is complete (or full) or incomplete (partial), we apply a WALD test (coefficient restrictions) with a null hypothesis $\alpha_1 = -1$ in the long run and $\psi_{1,0} = -1$ in the short run (corresponding to one period short-run pass-through). The null hypothesis of complete (full) pass-through of exchange rate is rejected if the calculated WALD test value exceeds the critical value notably determined by *F*-statistic or *t*-statistic.

4. Empirical Results and Discussion

This section presents the empirical results and discussion of the findings following the above-discussed methodology.
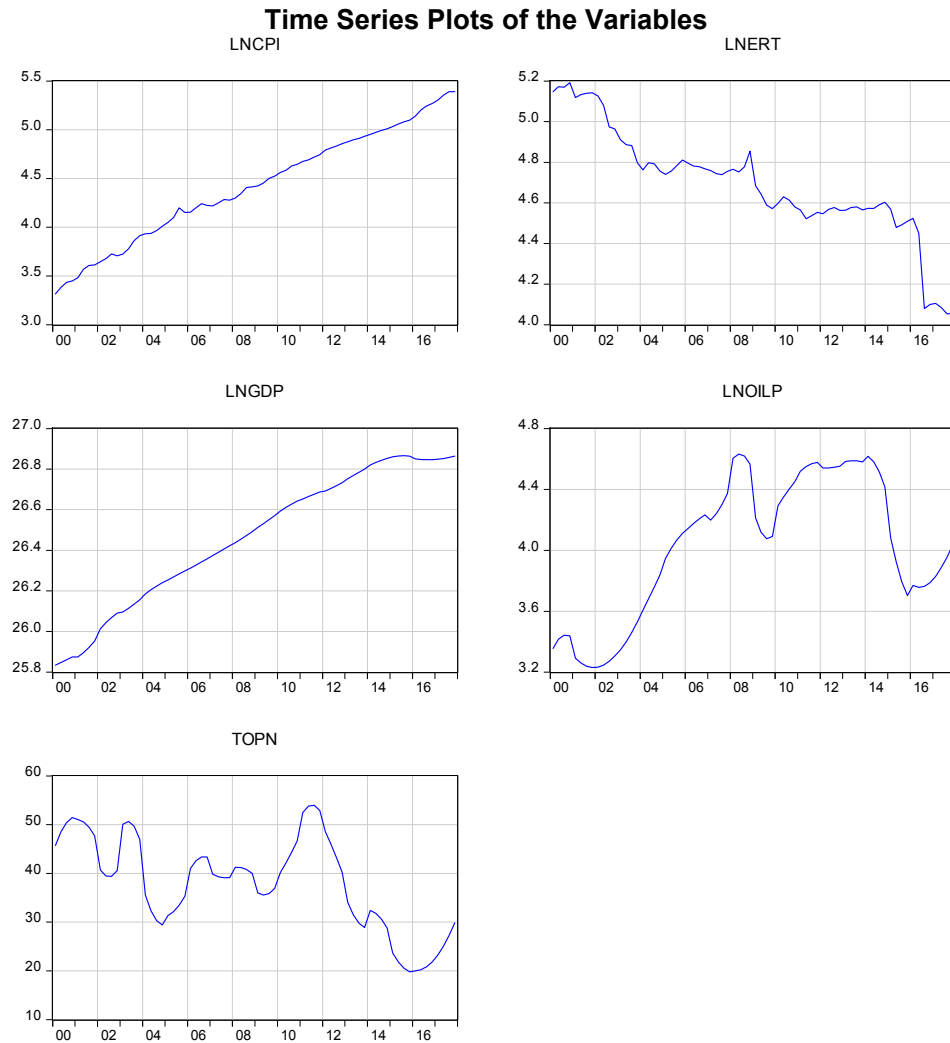
Table 1

Descriptive Statistics

	LNCPI	LNERT	LNGDP	LNOILP	TOPN
Mean	4.418225	4.691192	26.46414	4.026507	37.96349
Median	4.418930	4.711190	26.50195	4.086755	39.44313
Max.	5.391699	5.190819	26.86567	4.632468	53.93815
Min.	3.315570	4.054680	25.83446	3.230381	19.81374
Std. Dev.	0.573431	0.272069	0.328524	0.464518	9.629041
Skew.	-0.130117	-0.315829	-0.385549	-0.315224	-0.231354
Kurt.	1.957285	3.387154	1.904861	1.786164	2.118634
Jarque-Bera	3.464929	1.646643	5.381763	5.612592	2.972713
Prob.	0.176848	0.438971	0.067821	0.060428	0.226195
Sum	318.1122	337.7658	1905.418	289.9085	2733.371
Sum Sq. Dev.	23.34641	5.255533	7.662891	15.32014	6583.008
Observations	72	72	72	72	72

Note: The sample period covers 200Q1 to 2017Q4, resulting in a total of 72 observations.

Table 1 presents the descriptive statistics of the quarterly data for the variables used in this study. As Table 1 shows, TOPN has the largest mean with 37.96249, and LNOILP has the smallest mean across the variables over the period from 200Q1 to 2017Q4. The largest volatility in the series is found to that of TOPN, while the rest of the variables are less volatile. While all the variables exhibit negative skewness with estimates approximately zero in absolute terms, the kurtosis is positive across the variables, leading to non-rejection of the

Jarque-Bera probability at 5% significance level. The implication of this result is that the variables are all normally distributed. In addition, Figure 1 shows the time series plots of the variables, which exhibit breaks in the exchange rate, oil price, and trade openness, while the CPI and the GDP exhibit upward trends over the study period.

Figure 1

Time Series Plots of the Variables

The results of the minimum LM unit root test with one break developed by Lee and Strazicich (2013) is displayed in Table 2. As already mentioned in the methodology section, this test does not only accommodate information regarding unknown break, but addresses the problem of inaccurate estimation of breakpoint when there is break under the null and

alternative hypotheses. The results based on Model A (*i.e.* model with change in intercept) indicate that in the presence of a structural break, there strong evidence of unit root at level. After we take the first difference of these variables, they all become stationary. This means that all the variables are integrated of order one, I(1). However, the breakpoint estimates identified are 2003:Q2 for CPI, 2016:Q1 for the exchange rate, 2002:Q2 for GDP, 2009:Q2 for oil price and 2003:Q2 for trade openness. These breakpoints are attributed to structural changes occurred in the Nigerian economy. Particularly, the breakpoint in 2003:Q2 for CPI and trade openness is caused by the heavy wave of 2003 general elections, which increased the quantity of money in circulation and, consequently, increased the general price level. The break in 2016:Q1 for exchange rate is attributed to the crash in oil prices, which mounted inflationary pressure on the economy and, hence, loss of international competitiveness. The breakpoint in 2002:Q2 for GDP is due to the announcement of new policy thrust in the agricultural sector in late 2001 by the Federal Government. This increased the contribution of agriculture to GDP to 41% and two-third of employment in 2002. Moreover, the breakpoint in 2009:Q2 for oil price is occasioned by the global financial crisis of 2008, which shrank the demand for crude oil and, consequently, reduced its price.

Table 2

The Lee-Strazicich Unit Root Test

Variables		L-S Test at Level			L-S Test at First Difference			Critical Values		
		LM Statistic	Break Point		LM Statistic	Break Point		1%	5%	10%
Model A: Change in intercept										
LNCPI		-1.444 (5)	2003:Q2		-3.236* (3)	2002:Q2		-4.239	-3.566	-3.211
LNERT		-2.097 (0)	2016:Q1		-8.299*** (0)	2014:Q3		-4.239	-3.566	-3.211
LNGDP		-2.796 (5)	2002:Q2		-4.611*** (7)	2002:Q4		-4.239	-3.566	-3.211
LNOILP		-2.266 (5)	2009:Q2		-3.920** (4)	2015:Q3		-4.239	-3.566	-3.211
TOPN		-1.444 (5)	2003:Q2		-3.236* (7)	2002:Q2		-4.239	-3.566	-3.211
Model C: Shift in intercept and change in trend slope										
LNCPI		-4.295* (5)	2006:Q3					-5.050	-4.500	-4.180
LNERT		-2.894 (0)	2014:Q2					-5.070	-4.470	-4.200
LNGDP		-4.282* (5)	2012:Q2					-5.150	-4.450	-4.180
LNOILP		-3.091 (5)	2011:Q1					-5.050	-4.500	-4.180
TOPN		-3.584 (5)	2012:Q2					-5.150	-4.450	-4.180
LNCPI					-7.973*** (3)	2006:Q2		-5.050	-4.500	-4.180
LNERT					-9.014*** (0)	2015:Q4		-5.110	-4.500	-4.210
LNGDP					-5.613*** (7)	2014:Q3		-5.070	-4.470	-4.200
LNOILP					-5.871*** (7)	2013:Q3		-5.070	-4.470	-4.200
TOPN					-4.974** (7)	2009:Q4		-5.050	-4.500	-4.180

Notes: ***, **, and * denote significance at 1%, 5% and 10% level, respectively. The lag order is presented in the bracket ( ).

Furthermore, the results of Model C (*i.e.* the model with shift in intercept and change in trend slope) reveal that at level, there is no strong evidence of unit root in the presence of a structural break for LNCPI and LNGDP, while unit root is strong for oil price, exchange rate and trade openness. However, taking their first differences, we find that all the variables are stationary. The breakpoint estimates identified by the model include 2006:Q3 for CPI,

2014:Q2 for exchange rate, 2012:Q2 for GDP, 2011:Q1 for oil price and 2012:Q2 for trade openness. The 2006:Q3 identified for CPI could be explained by a gigantic policy change in the foreign exchange market by introducing the Dutch Auction System (RDAS) in 2006 to boost the supply of foreign exchange through pricing shares in Nigeria. The 2014:Q2 breakpoint identified for exchange rate and 2012:Q2 for GDP could be explained by the recent upheavals and oscillations in the crude oil prices, which result into trade deficit and shrinking of revenues for Nigeria. In addition, the breakpoint in 2011:Q1 is traceable to the political turmoil in Bahrain, Egypt, Libya, Yemen and other Middle East and North African oil producing countries. This significantly impacted the price of crude oil in the international crude oil market. Overall, based on the model with change in intercept and the model with shift in intercept and change in trend slope, the results of the unit root tests show that all the variables are stationary in their first differences.

Table 3 reports the results of the Bayer-Hanck combined cointegration test. In performing this test, we treat each of the variables as the dependent variable; hence, we have five equations. The results of this test show that the Fisher statistics for the two tests exceed the critical value at 5% level of significance. This means that the null hypothesis of no cointegration is certainly rejected, indicating the presence of a valid long-run relationship among CPI, exchange rate, GDP, oil price, and trade openness for the period 2000Q1 to 2017Q4 in Nigeria. We test for the robustness of the Bayer-Hanck combined cointegration test by using the ARDL bounds testing approach proposed by Pesaran *et al*. (2001), which accommodates structural breaks. The results of the ARDL bounds testing approach as revealed in Table 4, aptly find cointegrating for all four cases, which validates the results of the Bayer-Hanck combined cointegration test that a valid long-run relationship is found among the variables.

Table 3

Results of Bayer-Hanck Cointegration Test

Model Estimated	EG–JOH	EG–JOH–BO–BDM	Cointegration
$LNCPI = f(LNERT, LNGDP, LNOILP, TOPN)$	56.207[**]	65.532[**]	Yes
$LNERT = f(LNCPI, LNGDP, LNOILP, TOPN)$	55.658[**]	57.375[**]	Yes
$LNGDP = f(LNERT, LNCPI, LNOILP, TOPN)$	56.672[**]	112.69[**]	Yes
$LNOILP = f(LNGDP, LNERT, LNCPI, TOPN)$	55.772[**]	112.29[**]	Yes
$TOPN = f(LNOILP, LNGDP, LNERT, LNCPI)$	55.371[**]	78.144[**]	Yes

*Notes: ** denotes that the null hypothesis of no cointegration is rejected 5%. Critical value at 5% level is 10.576 for EG–JOH and 20.143 for EG–JOH–BO–BDM. Maximum lag order is 4.*

Table 4

Results of Cointegration using ARDL Bounds Testing Approach

Variable	LNCPI	LNERT	LNGDP	LNOILP	TOPN
F-Statistics	7.983[**]	2.811	6.717[**]	10.895[**]	4.013[**]
Break Points	2006:Q3	2014:Q2	2012:Q2	2011:Q1	2012:Q2
Critical value	5% Level				
Lower bounds	2.86				
Upper bounds	4.01				

*Notes: ** denotes that the null hypothesis of no cointegration is rejected at 5% level of significance. The critical value is determined where $k = 4$ independent variables with unrestricted intercept and*

no trend. The maximum lag order is 7 and the optimal lag order is selected by the Akaike Information Criterion (AIC), which is 6.

Table 5 reveals the long-run estimates based on the ARDL approach. Following the measurement of the exchange rate, which corresponds to the nominal effective exchange rate, an increase in the exchange rate signals appreciation of domestic currency with respect to a weighted basket of currencies of trading partners. This literarily means that exports become more expensive and imports become cheaper leading to loss of international competitiveness. The empirical results presented in Table 5 indicate that in the long run the exchange rate has a negative, inelastic and statistical relationship with inflation at 5% significance level ($\alpha_1 = -0.364, p < 0.10$). This result implies that keeping other things constant, a 1% increase in the exchange rate (*i.e.* an appreciation of domestic currency) would decrease inflation by roughly 0.364%. Moreover, to examine whether the pass-through is complete or incomplete in the long run, we employ a Wald test (coefficient restrictions). In this test, the null hypothesis is that $\alpha_1 = -1$, which means that the pass-through in Nigeria is complete during the period of economic reforms and the alternative is given as $\alpha_1 > -1$, which means that the pass-through is incomplete during the period of economic reforms. The value for F-statistic is as high as 15.662 and that of t-statistic is -3.958. These results definitely reject the null hypothesis that the pass-through is complete in the long run at a 1% level. The rejection of two test statistics strongly indicates that the pass-through of the exchange rate, in the long run, is incomplete in Nigeria during the period of economic reforms. This confirms that the prices are sticky during the period when inflation is relatively low. It also validates the effectiveness of monetary policy and inflation targeting policy of the Central Bank of Nigeria. Therefore, our finding agrees with Oyinlola and Babatunde (2009); Oyinlola and Egwaikhide (2011); Bada *et al.* (2016) who report incomplete pass-through of the exchange rate to inflation in Nigeria in the long run. However, this result is not congenial with the recent finding by Balcilar *et al.* (2019), who report that the pass-through in Nigeria is complete in the long run. The difference between our result and that of Balcilar *et al.* (2019) confirms the role of low and stable inflation on the ERPT.

Table 5

Long-Run Coefficients

Dependent Variable = ΔLNCPI			
Variable	Coefficient	*t*-Statistic	*p*-value
LNERT	-0.364[*]	-1.800	0.086
LNGDP	1.439[***]	11.45	0.000
LNOILP	-0.149[***]	-3.864	0.001
TOPN	0.007[***]	3.882	0.001
Residual Diagnostic Tests	Statistic	*p*-value	
$\chi^2 - ARCH$	0.336: [1]	0.564	
$\chi^2 - SERIAL$	4.418: [1]	0.056	
$\chi^2 - RESET$	1.468: [1]	0.157	
$\chi^2 - NORMAL$	32.380	0.000	

*Notes: *** and ** denote significance at 1% and 5% level of significance respectively. $\chi^2 - ARCH$, $\chi^2 - SERIAL$, $\chi^2 - RESET$ and $\chi^2 - NORMAL$ are tests for serial-correlation, heteroscedasticity, functional as well as normality test. [ ] represents the optimal lag selection for diagnostic tests with an unrestricted constant and no trend.*

Furthermore, we find that the long-run coefficient of output is positive, elastic, and easily passing a statistically significant test at 1% level ($\alpha_2 = 1.439, p < 0.01$). This result divulges that a 1% increase in output would lead to approximately 1.439% increase in inflation, pointing to the fact that output stimulates demand, which consequently increases prices in the long run. The results also suggest a negative and positive relationship between oil price and the degree of trade openness with inflation. This relationship is inelastic and statistically significant at 1% level ($\alpha_3 = -0.149, p < 0.01$) and ($\alpha_4 = 0.007, p < 0.01$). By implication, a 1% increase in oil price reduces inflation by 0.149%, while a 1% increase in the degree of trade openness makes inflation to rise by 0.007%. Therefore, our finding on the positive effect of output on inflation is consistent with Baharumshah *et al.* (2017), who find that output is inflationary in the Sudanese economy, but delink the anti-inflationary role of oil price. Also, the findings corroborate Balcilar *et al.* (2019) that the effect of output growth in the Nigerian inflationary process is positive and elastic, while oil price noticeably reduces inflation in Nigeria. Similarly, the result of the negative effect of oil price agrees with Aliyu *et al.* (2009) and Bada *et al.* (2016), who confirm that a rise in oil price is anti-inflationary in Nigeria. Furthermore, the positive effect of the degree of trade openness on inflation concurs with the conventional wisdom that the pass-through increases with the degree of global integration of a country. To this extent, our finding is consistent with Aliyu *et al.* (2009) and Usman and Musa (2018), who find that the pass-through of exchange rate increases with the increase in openness of the economy towards the global economic integration. On the contrary, our result disagrees with Lane and Milesi-Ferretti, (2007); Wynne and Kersting (2007); and Mukhtar (2010) who validate the inverse relationship between openness and inflation as hypothesized by Romer (1993).

Table 5 also provides the residual-based diagnostic tests. Based on the results of the Breusch-Godfrey Lagrange multiplier test for serial correlation at lag 1 and the ARCH test for conditional heteroscedasticity at lag 1, the null hypothesis cannot be rejected at 5% significant level. Therefore, we strongly conclude that there is no evidence of serial correlation and conditional heteroscedasticity. The result based on the Ramsey's regression specification error test (RESET) test at lag 1 indicates that the functional form of the model is correct and well-specified, while the account of the model's error term shows no evidence of normal distribution.[11] In addition, the plots of cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) in Figure 2 provide that the model parameters are stable as the CUSUM and CUSUMSQ lie within the 5% critical bounds.

---

[11] *Following the central limit theorem, the issue of normal distribution of model's error terms is a weak assumption especially in time series. This is because as a sample size increases, the series tends toward normal distribution*

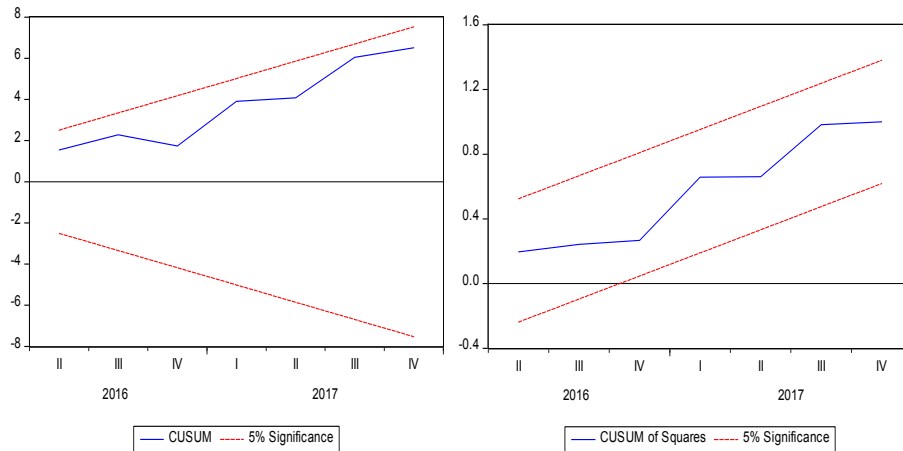## The CUSUM and CUSUM of Squares Tests

Table 6 reports the error correlation model and the short-run coefficients. The results of the breakpoints included in the estimation of the long run reveal that only the breakpoint for 2011:Q1 is statistically significant, easily passing a statistical test at 10% level. This finding, therefore, implies that the crude oil price shocks arising from the political turmoil in Bahrain, Egypt, Libya, Yemen and other Middle East and North African oil producing countries have positive and statistically significant effects on the inflationary pressure in Nigeria. Moreover, the coefficient of the error correction term is -0.476, which is negative and statistically significant at 1% level. This suggests that inflation apparently converges to the equilibrium long-run level by roughly 47.6% adjustment speed on a quarterly basis via the changes in the exchange rate, GDP, oil price and degree of openness of the economy to trade.

Table 6

Error Correction Regression Coefficients

Dependent Variable = ΔLNCPI			
Variable	Coefficient	t-Statistic	*p*-value
Constant	-14.876***	-6.853	0.000
ΔLNCPI$_{t-1}$	-0.078	-0.763	0.454
ΔLNCPI$_{t-2}$	-0.331**	-3.038	0.006
ΔLNCP$_{t-3}$	-0.470***	-4.619	0.000
ΔLNCPI$_{t-4}$	-0.082	-0.660	0.516
ΔLNCPI$_{t-5}$	-0.176*	-1.767	0.091
ΔLNCP$_{t-6}$	-0.166	-1.429	0.167
ΔLNERT$_t$	-0.082**	-2.258	0.034
ΔLNERT$_{t-1}$	0.101**	2.664	0.014
ΔLNERT$_{t-2}$	0.010	0.256	0.801
ΔLNERT$_{t-3}$	-0.050	-1.243	0.227
ΔLNERT$_{t-4}$	-0.0654*	-1.856	0.077
ΔLNERT$_{t-5}$	-0.061	-1.635	0.116

Dependent Variable = ΔLNCPI			
Variable	Coefficient	t-Statistic	*p*-value
ΔLNERT$_{t-6}$	-0.276***	-3.888	0.001
ΔLNGDP$_t$	0.591	1.557	0.134
ΔLNGDP$_{t-1}$	-1.165**	-2.380	0.026
ΔLNGDP$_{t-2}$	-0.469	-1.105	0.281
ΔLNGDP$_{t-3}$	-0.393	-1.074	0.294
ΔLNGDP$_{t-4}$	-2.016**	-2.854	0.009
ΔLNGDP$_{t-5}$	-2.552***	-3.409	0.003
ΔLNOILP$_t$	-0.193***	-3.328	0.003
ΔLNOILP$_{t-1}$	-0.079	-1.399	0.176
ΔLNOILP$_{t-2}$	0.129**	3.014	0.006
ΔLNOILP$_{t-3}$	0.076**	2.263	0.034
ΔLNOILP$_{t-4}$	-0.090*	-1.955	0.063
ΔLNOILP$_{t-5}$	-0.077	-1.630	0.117
ΔLNOILP$_{t-6}$	0.205***	4.454	0.000
ΔTOPN$_t$	0.004**	2.256	0.034
ΔTOPN$_{t-1}$	0.002	1.080	0.292
ΔTOPN$_{t-2}$	-0.006***	-4.145	0.000
ΔTOPN$_{t-3}$	-0.001	-0.775	0.446
ΔTOPN$_{t-4}$	-0.002*	-1.889	0.072
ΔTOPN$_{t-5}$	-0.001	-0.992	0.332
ΔTOPN$_{t-6}$	-0.007***	-6.205	0.000
D2006Q3	0.025	1.246	0.226
D2014Q2	0.007	0.723	0.477
D2012Q2	-0.015	-1.263	0.220
D2011Q1	0.033*	1.973	0.061
ect$_{t-1}$	-0.476***	-6.868	0.000

Note: *** and ** denote significance at 1% and 5% level of significance, respectively.

Furthermore, the short-run empirical result for exchange rate is negative, inelastic and statistically significant at 5% level ($\psi_{1,0} = -0.082, p < 0.05$). This result indicates that a 10% increase in exchange rate (appreciation) causes inflation to decrease by 0.82% in the short run. Similar to the case of long run, we apply the Wald test (coefficient restrictions) based on the null hypothesis $\psi_{1,0} = -1$ against the alternative $\psi_{1,0} > -1$. The value of F-statistic is 50.959 and t-statistic is -7.139, which both reject the null hypothesis of complete pass-through at 1% significance level. This rejection strongly implies that the short-run pass-through during the period of economic reforms in Nigeria is incomplete. However, when comparing the magnitude of the pass-through we observe that the short-run pass-through is about five times lower than the long-run one. This finding concurs with the sticky-price monetary model, which argues that the existence of imperfect information and inertia in consumer habits are the major reasons why prices are sticky and do not change instantaneously, but adjust gradually over time. Therefore, this finding validates the earlier findings documented by Aliyu *et al.* (2009), Oyinlola and Babatunde (2009) Oyinlola and Egwaikhide (2011), Lariau, *et al.* (2016) and Balcilar *et al.* (2019) that the short-run ERPT is incomplete in Nigeria.

In addition, the short run coefficient of GDP suggests a positive, elastic and significant effect only in the first lag ($\psi_{2,0} = 1.165, p < 0.05$). This result points to the fact that a 1% increase in output would cause inflation to rise by 1.165%. Also, the results reveal that oil price and degree of openness to trade have negative and positive relationships with inflation in the short run. These relationships are inelastic and easily pass the significance test at 1% and 5% levels, $(\psi_{3,0} = -0.193, p < 0.01)$ and $(\psi_{4,0} = 0.004, p < 0.05)$, respectively. This means that a 0.193% decrease in inflation is linked to 1% rise in oil price, while 0.004% increase in inflation is traceable to 1% increase in trade openness leaving other things unchanged. Comparing these findings with the previous studies, we find that the inflationary effect of output growth is consistent with Balcilar *et al.* (2018), who find the same situation in South Africa between 1922M01 and 2013M07. However, the finding contradicts Balcilar *et al.* (2019), who discover that the short-run effect of output growth is disinflationary. Our finding on the short-term effect of oil price concurs with Balcilar *et al.* (2017), who reveal that oil price is related to real output growth in South Africa. Also, the positive relationship between openness and inflation cast a doubt on the validity of Romer's (1993) hypothesis. Therefore, the result is consistent with Samimi *et al.* (2012), who find an inverse relationship flowing between trade openness and inflation using panel data.

5. Summary and Conclusion

The main objective of this study is to measure the long-run and short-run exchange rate pass-through to inflation during the period of economic reforms, which coincides with the introduction of inflation targeting. This period is characterized by a relatively low and stable level of inflation. To achieve this objective, we consider the possible structural breaks by applying the recently developed minimum Lagrange multiplier (LM) unit root test with one break proposed by Lee and Strazicich (2013), Bayer and Hanck (2013) combined cointegration test, and the estimation approach based on the ARDL bounds test proposed by Pesaran *et al.* (2001). Quarterly time series data including CPI, nominal effective exchange rate, GDP, oil price, and trade openness are used in the study for the period 2000Q1 to 2017Q4. The results of the unit root test display that all the variables are integrated of the same order. The cointegration test results establish a valid cointegration among the variables. The robustness of the combined cointegration test is performed using the bounds testing approach, which accommodates structural breaks. The results also confirm that there is cointegration among the variables in the presence of structural breaks. In addition, the results based on the flexible unrestricted error correction model reveal that the long-run and short-run pass-through of the exchange rate to inflation is incomplete in Nigeria during the period of economic reforms. The results also reveal that the pass-through is stronger in the long run compared to the short run, thus confirming the monetary theory of short-term price stickiness. Moreover, the results show that output growth is positively impacting inflation, while the oil price is negatively impacting inflation. Finally, the effect of trade openness on inflation is positive, which casts doubt on Romer's hypothesis in Nigeria. Furthermore, the deviation from the long-run equilibrium is corrected by 47.6% speed of adjustment in every quarter. Therefore, a vital policy implication of our findings is that the degree of the pass-through of the exchange rate ought to be considered in deciding the choice of the monetary policy framework. In addition, efforts should be made to stimulate the supply-side of the economy in order to reduce demand-pull inflation and, consequently, promote competition and efficiency, upon which the gains from trade openness are based.

References

Aleem, A. and Lahiani, A., 2014. Monetary policy credibility and exchange rate pass-through: Some evidence from emerging countries. *Economic Modelling*, *43*, pp.21-29.

Aliyu, S. U. R., and Englama, A., 2009. Is Nigeria ready for inflation targeting? *MPRA* [online] Available at: <http://mpra.ub.uni-muenchen.de/14870/>.

Aliyu, S.U.R., Yakub, M.U., Sanni, G.K., and Duke O.O., 2009. Exchange Rate Pass-through in Nigeria: Evidence from a Vector Error Correction Model. *MPRA Paper* No. 2553. [online] Available at: <https://mpra.ub.uni-muenchen.de/25053/>.

Bada, A.S., Olufemi, A.I., Tata, I.A., Peters, I., Bawa, S., Onwubiko, A.J. and Onyowo, U.C., 2016. Exchange rate pass-through to inflation in Nigeria. *CBN Journal of Applied Statistics*, 7(1), pp.49-70.

Balcilar, M., Usman O., and Agbede, E.A., 2019. Revisiting the Exchange Rate Pass-Through to Inflation in Africa's two largest economies: Nigeria and South Africa. *African Development Review*, 31(2), pp.245-257.

Baharumshah, A.Z., Sirag, A. and Mohamed Nor, N., 2017. Asymmetric Exchange Rate Pass-through in Sudan: Does Inflation React Differently during Periods of Currency Depreciation?. *African Development Review*, *29*(3), pp.446-457.

Bailliu, J. and Bouakez, H., 2004. Exchange rate pass-through in industrialized countries. *Bank of Canada Review*, *2004*(Spring), pp.19-28.

Bamidele, A., 2007. Pre-requisite for Inflation Targeting Country Experiences and Lessons for Nigeria, being a paper delivered at the *Inflation Targeting Workshop* organized by the Central Bank of Nigeria Learning Centre Lagos, July.

Banerjee, A., Dolado, J. and Mestre, R., 1998. Error-correction mechanism tests for cointegration in a single-equation framework. *Journal of time series analysis*, *19*(3), pp.267-283.

Bayer, C. and Hanck, C., 2013. Combining non-cointegration tests. *Journal of Time Series Analysis*, *34*(1), pp.83-95.

Boswijk, H.P., 1994. Testing for an unstable root in conditional and structural error correction models. *Journal of econometrics*, *63*(1), pp.37-60.

Cheikh, N.B. and Louhichi, W., 2016. Revisiting the role of inflation environment in exchange rate pass-through: A panel threshold approach. *Economic Modelling*, *52*, pp.233-238.

Campa, J.M. and Goldberg, L.S., 2005. Exchange rate pass-through into import prices. *Review of Economics and Statistics*, *87*(4), pp.679-690.

Campa, J.M. and Goldberg, L.S., 2002. *Exchange rate pass-through into import prices: A macro or micro phenomenon?* (No. w8934). National Bureau of Economic Research.

Choudhri, E.U. and Hakura, D.S., 2006. Exchange rate pass-through to domestic prices: Does the inflationary environment matter?. *Journal of International Money and Finance*, *25*(4), pp.614-639.

Corsetti, G. and Pesenti, P., 2005. International dimensions of optimal monetary policy. *Journal of Monetary economics*, *52*(2), pp.281-305.

Cunningham, R.M., Friedrich, C., Hess, K. and Kim, M.J., 2017. *Understanding the time variation in exchange rate pass-through to import prices* (No. 2017-12). Bank of Canada Staff Discussion Paper.

Dickey, D.A. and Fuller, W.A., 1979. Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American statistical association*, *74*(366a), pp.427-431.

Dube, S., 2016. Exchange Rate Pass-Through (ERPT) and Inflation-Targeting (IT): Evidence from South Africa-Exchange rate pass-through and inflation targeting: evidenze dal Sud Africa. *Economia Internazionale/International Economics*, *69*(2), pp.121-150.

Elliot, G., Rothenberg, T. J. and Stock, J. H., 1996. Efficient tests for an autoregressive unit root, *Econometrica*, 64, pp.813-836.

Engle, R.F. and Granger, C.W., 1987. Co-integration and error correction: representation, estimation, and testing. *Econometrica: journal of the Econometric Society*, pp.251-276.

Essien, E.A., 2005. Exchange rate pass-through to inflation in Nigeria. *West African Journal of Monetary and Economic Integration*, *5*(1), pp.89-103.

Fisher, R.A., 1992. Statistical methods for research workers. In *Breakthroughs in statistics* (pp. 66-70). Springer, New York, NY.

Gagnon, J.E. and Ihrig, J., 2004. Monetary policy and exchange rate pass-through. *International Journal of Finance & Economics*, *9*(4), pp.315-338.

Granato, J., Lo, M. and Wong, M.S., 2007. A note on Romer's openness-inflation relation: the responsiveness of AS and AD to economic openness and monetary policy. *Applied Economics*, *39*(2), pp.191-197.

Hajnal, M., Molnár, G. and Várhegyi, J., 2015. *Exchange rate pass-through after the crisis: the Hungarian experience* (No. 121). MNB Occasional Papers.

Hanif, M.N. and Batool, I., 2006. Openness and inflation: A case study of Pakistan. *MPRA Paper* No. 10214. [online] Available at: <http://mpra.ub.uni-muenchen.de/10214/>.

Johansen, S., 1995. A stastistical analysis of cointegration for I (2) variables. *Econometric Theory*, *11*(1), pp.25-59.

Jin, J.C., 2000. Openness and growth: an interpretation of empirical evidence from East Asian countries. *Journal of International Trade & Economic Development*, *9*(1), pp.5-17.

Junttila, J. and Korhonen, M., 2012. The role of inflation regime in the exchange rate pass-through to import prices. *International Review of Economics & Finance*, *24*, pp.88-96.

Kabundi, A.N. and Mlachila, M.M., 2018. *Monetary policy credibility and exchange rate pass-through in South Africa*. International Monetary Fund, WP/18/173.

Kenny, G. and McGettigan, D., 1998. Exchange rates and import prices for a small open economy: the case of Ireland. *Applied Economics*, *30*(9), pp.1147-1155.

Kurtović, S., Siljković, B., Denić, N., Petković, D., Mladenović, S.S., Mladenović, I. and Milovancevic, M., 2018. Exchange rate pass-through and Southeast European economies. *Physica A: Statistical Mechanics and its Applications*, *503*, pp.400-409.

Kwiatkowski, D., Phillips, P.C., Schmidt, P. and Shin, Y., 1992. Testing the null hypothesis of stationarity against the alternative of a unit root: How sure are we that economic time series have a unit root?. *Journal of econometrics*, *54*(1-3), pp.159-178.

Lane, P.R. and Milesi-Ferretti, G.M., 2007. The external wealth of nations mark II: Revised and extended estimates of foreign assets and liabilities, 1970–2004. *Journal of international Economics*, *73*(2), pp.223-250.

Lariau, A., El-Said, M. and Takebe, M.M., 2016. *An Assessment of the Exchange Rate Pass-Through in Angola and Nigeria*. International Monetary Fund.

Lee, J. and Strazicich, M.C., 2013. Minimum LM unit root test with one structural break. *Economics Bulletin, 33(4),* pp. 2483-2492.

Mann, C.L., 1986. Prices, profit margins, and exchange rates. *Federal Reserve Bulletin*, *72*, pp.366.

Marazzi, M., Sheets, N., Vigfusson, R., Faust, J., Gagnon, J., Marquez, J., Martin, R., Reeve, T. and Rogers, J., 2005. Exchange rate pass-through to US import prices: some new evidence. *International Finance Discussion Papers*, *833*.

Menon, J., 1993. Exchange rate pass-through: Australian imports of motor vehicles. *International Economic Journal*, *7*(3), pp.93-109.

Menon, J., 1995. Exchange rate pass-through. *Journal of Economic Surveys*, *9*(2), pp.197-231.

McCarthy, J., 2007. Pass-through of exchange rates and import prices to domestic inflation in some industrialized economies. *Eastern Economic Journal*, *33*(4), pp.511-537.

Mukhtar, T., 2012. Does trade openness reduce inflation? Empirical evidence from Pakistan. *Journal of Economic Cooperation and Development*, *33*(2), pp.33-52.

Ng, S. and Perron, P., 2001. Lag length selection and the construction of unit root tests with good size and power. *Econometrica*, *69*(6), pp.1519-1554.

Ogundipe, A.A. and Samuel, E., 2013. Exchange rate pass-through to consumer prices in Nigeria. *Journal of Business Management and Applied Economics* 3(4), pp.1-10.

Oladipo, O., 2007. Exchange rate pass-through: a case study of a small open economy. *Global Economy Journal*, *7*(3), p.1850112.

Omisakin, O., 2009. Exchange rate pass through to domestic price and output in Nigeria. *International Business Management*, *3*(3), pp.38-42.

Otani, A., Shiratsuka, S. and Shirota, T., 2003. *The decline in the exchange rate pass-through: evidence from Japanese import prices*. Institute for Monetary and Economic Studies, Bank of Japan.

Oyinlola, M.A. and Babatunde, M.A., 2009. A bound testing analysis of exchange rate pass-through to aggregate import prices in Nigeria: 1980-2006. *Journal of Economic Development*, *34*(2), p.97.

Oyinlola, M.A., and Egwaikhide, F.O., 2011. Exchange rate pass-through to domestic prices in Nigeria: A dynamic investigation", *Journal of Social Sciences*, 6, p.87-95.

Perron, P., (1989). The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis, *Econometrica* 57, pp. 1361-1401.

Pesaran, M.H., Shin, Y. and Smith, R.J., 2001. Bounds testing approaches to the analysis of level relationships. *Journal of applied econometrics*, *16*(3), pp.289-326.

Phillips, P.C. and Perron, P., 1988. Testing for a unit root in time series regression. *Biometrika*, *75*(2), pp.335-346.

Pollard, P.S. and Coughlin, C.C., 2004. Size matters: asymmetric exchange rate pass-through at the industry level. *University of Nottingham Research Paper*, (2004/13).

Poloamina, I.D., Babatunde, M., and Oyinlola, M.A., 2009. Is exchange rate pass-through to import prices in Nigeria a short-run phenomenon? *Ibadan Journal of the Social Sciences,* 7(1), pp. 36-48.

Rogoff, K., (1996). The Purchasing Power Parity Puzzle, *Journal of Economic Literature*, XXXIV, 647-668.

Romer, D., 1993. Openness and Inflation: Theory and Evidence, *Quarterly Journal of Economics,* 4, pp. 869-903.

Samimi, A. J., Ghaderi, S., Hosseinzadeh, R., & Nademi, Y., 2012. Openness and inflation: New empirical panel data evidence. *Economics Letters*, *117*(3), pp.573-577.

Savensson, L. E. O., 1999. Price Level Targeting vs. Inflation Targeting: A Free Lunch? *Journal of Money, Credit, and Banking,* 31 (August), pp.277-295.

Taylor, J., 2000. Low Inflation, Pass-Through and the Pricing Power of Firms. *European Economic Review,* 44, pp.1389-1408.

Uchendu, O. A., 2000. Inflation Targeting: A Monetary Policy Management Framework for the Attainment of Price Stability in Nigeria, *CBN Economic and Financial Review (EFR),* 38(2), June, pp.1-16.

Usman, O. and Musa, S.M., 2018. Revisiting Exchange Rate Pass-Through to Consumer Price Inflation in Nigeria: A Cointegrated Vector Autoregressive Approach. *Academic Journal of Economic Studies*, *4*(1), pp.60-67.

Wynne, M.A., Kersting, E.K., 2007. Openness and inflation. Staff Papers. Federal Reserve Bank of Dallas. No. 2.

Zivot, E. and Andrews, D.W.K., 1992. Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis. *Journal of business & economic statistics*, 10: pp.251-270.

Zubair, A., George, O. and Sanusi, A.R., 2013. Exchange Rate Pass-Through to Domestic Prices in Nigeria: An Empirical Investigation. *Central Bank of Nigeria Economic and Financial Review*, *51*(1).