

Price-Wage Spiral in Bangladesh: Evidence from ARDL Bound Testing Approach

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Abstract: This paper explores the interrelationship between price and nominal wage during the period of Fiscal Year 1975-76 to Fiscal Year 2009-10. An ARDL bound testing approach suggested by Pesaran et al. (2001) is employed to investigate the short-run dynamics and long-run relationship between consumer price index (CPI) and nominal wage rate, while accounting for the underlying macro determinants of price inflation, namely, the bilateral (TK/US dollar) nominal exchange rate, domestic credit and real gross domestic product (GDP). Our study shows that nominal wage rate and domestic credit exhibit a positive relation with the price level both in the short-run and in the long run. Moreover our result is robust as FMOLS, DOLS and Johansen-Juselius procedure ensuring that there is a long-run relationship between price and other macroeconomic variables under study. Again, the augmented Granger causality test ensures that wage and domestic credit have the impact on the inflationary process in both short-run and long-run. These results endorse the fiscal view of inflation which is more important for the stability of price level in Bangladesh.

Keywords: wage, price, bound testing, cointegration

JEL Classification: C 32, E31, E370, E5, E64

1. Introduction

Price dynamics and its movement is one of the most discussant issues in macroeconomics. The interdependencies between price and wage and their affect on aggregate demand are often analyzed by economists and policymakers. Continuous rise in price level or inflationary process and its determinants have long been debated in literatures. Research on recent inflationary process in developing countries often leads the economists to distinguish the views into the 'fiscal view' and the 'balance of payment view'. Advocates of fiscal view claim that inflation and excessive monetary growth is closely related while they opined that the leading reason for this excessive money growth is the financing of budget deficit while the developing countries very often experienced with the deficit in their budget as in most cases government expenditure always surpasses the total revenues.¹ This reason accelerates not only the inflationary process, but also the adverse effect of this inflation often tend to increase the relationship between money growth, budget deficit and inflation. On the other hand, the proponents of balance of payments view emphasize on the induced impact of the exchange rate movements on the inflationary process. This type of movements like depreciation of exchange rate are often become the result of development of balance of payment problem which in turn give excessive government

expenditure or sharp deterioration of terms of trade.² The final result rests into prolonged inflation.

For gauging inflationary pressure, policymakers and financial analysts pay close attention to the behavior of nominal and real wages. It is widely believed that if wage rate rises faster than that of productivity, then, it is the wage that creates an upward pressure on prices as higher wage costs pass on the prices. With uncertainty about the unemployment rate's reliability as an early-warning device for rising inflation, recent attention has turned to wage and compensation growth for a labor market indicator of inflation. Surprisingly, the recent shift in attention to higher wages as the cause of higher prices leaves unexplained the problem of how wages get high in the first place. In the real world, the price and wage rigidities, which is important for the evolution of macroeconomics for the responses to shocks, is expected to translate into persistent responses of real wages, as well as of price and wage inflation to shocks hitting the economy. Evaluating the degree of persistence of such responses is thus an issue of paramount importance.

The reason for studying the relationship between wage and price level lies into the policy formulation of productivity enhancement, inflation control, consumption stimulation and for formulation of monetary and fiscal policy. Bangladesh experiences an increasing trend both in CPI for industrial labor and nominal wage rate index.³ During the FY 2004-05, the average inflation rate is 6.48% where in FY 2010-11 it is 8.80%. In FY 2008-09, it is reduced to 6.66% but later it increases. During this stage, we see an upward trend in nominal wage rate index as it rose from 3293 to 5782, an average increase of 74%. This gives us a snapshot about the wage and price level relationship. As we discussed earlier the causes of rise in price level and it will be discussed later in next section, it urges us to identify the relationship between the price level and wages.

With this line, this paper tries to examine the relationship between domestic price level and nominal wage rate index in view to identify the price dynamics to search for short-run and long-run relationship between price and wage. To do this, we employed ARDL bound testing approach suggested by Pesaran et al. (2001). The following sections describe theoretical and empirical findings, Inflation and the rationale for the choice of variables, Data description and Methodology, Empirical Analysis and Concluding Remarks.

2. Theoretical and Empirical Findings on the Link between Wages and Prices

A positive aggregate demand shock (such as an increase in government spending, or, the devaluation of the exchange rate) results an upward pressure on the demand for domestic goods and thus increases the price level. The higher the price level, the higher would be the demand for labor which is associated with higher aggregate demand. This higher demand for labor would create an upward pressure on the equilibrium nominal wage and thus increase the nominal wage (Blanchard, 1986). On the other hand, if the firms try to increase mark-up profits (i.e., mark-up prices above wages) or, if workers try to increase their real wages, or if both firms and workers try to maintain the same price and wage in the face of a negative supply shock, this may give rise to 'cost-push' inflation (Blanchard, 1986). Moreover, according to the 'new-structuralist' tradition, an increase in the real wage demanded by workers may trigger an unstable spiral between nominal wages and prices (Agénor and Hoffmaister, 1997). Montiel (1997) explored the

dynamic relationship among prices, base money, nominal exchange rate, nominal wages and real output for Argentina, Brazil and Israel. In his study, he found by employing vector autoregression for Israel, the main driver of inflation is wage when it was motivated by the changes in base money and nominal exchange rate. Again for Argentina and Brazil, he found that base money and exchange rate was the main driver of inflation. Again, Agénor and Hoffmaister (1997) examined the short-run relationship between money growth, exchange rate depreciation, nominal wage growth, output gap and inflation in Chile, Korea, Mexico and Turkey. They found that the affect is persistent in Korea and Mexico. Different results are found from various studies estimating the pass-through of exchange rate movements for developing countries. Loungani and Swagel (2001) found positive relationship between the exchange rate depreciation and the inflation rate in such kind of countries. The authors estimated this effect using a panel of 53 developing countries: African countries – 16, Asian – 11, South American – 19, and Mediterranean – 7. The case of the developing counties with the floating exchange rate suggests that the impact of exchange rate depreciation on the price changes is positive and statistically significant. The same results are obtained in the studies estimating the relationship between the exchange rate and inflation separately for individual developing countries. In his article, Mwase (2006) showed negative exchange rate effect on inflation, which becomes stronger in the long-run compared to the short-run. Considerable attention was paid to the exchange rate impacts, as one of the major drivers, on inflation in the papers by Kuijs (1998); Callen and Chang (1999); Ubide (1997); Cerisola and Gelos(2009); Leigh and Rossi (2002); Hossain (2002); Sacerdoti and Xiao (2001); Khan and Schimmelpfennig (2006) analyzing inflation dynamics respectively for Nigeria, India, Mozambique, Brazil, Turkey, Bangladesh, Madagascar and Pakistan. These papers suggested a negative relationship between the exchange rate and inflation in developing economies. Grigorian et al. (2004) identified interrelations of consumer price, exchange rate and nominal wages. By studying three inter-related markets (foreign exchange, money and labor), the authors analyzed the dynamic effects of the exchange rate on prices in Armenia. Their estimation showed higher responsiveness of inflation to the exchange rate rather than to the other determinants (money supply and nominal wages). The paper suggested a negative correlation, both in the short- and long-run.

Although the role of wages in the inflationary process is well recognized, few empirical studies have been conducted in the developing countries context which tried to capture the effects. Montiel (1989) has found positive relationship though other studies (Dornbusch and Wolf, 1990; De Haan and Zelhorst, 1990; Moser, 1995; Loungani and Swagel, 1995) did not find any significant relationship as the latter studies suggested the wage-price spiral relationship was existed during the investigation. Again, some studies like Hoffmaister and Roldos (1997) investigated the indirect effect of nominal wage on inflation by employing the wage shocks on prices through aggregate nominal shocks, but they did not find any specific contribution of wages. So, based on the studies, we would try to investigate the relationship among the nominal wages and inflation.

As for empirical studies in Bangladesh, a number of studies have been conducted on the determinants of inflation and the relationship between inflation and other macroeconomic variables. Hossain (1987) examined the relationship between fiscal deficit and inflation from 1973 to 1983 using quarterly data. He found that government expenditures adjusted themselves to inflation more rapidly to government revenues and increase the size of fiscal deficit during

inflation. He stressed that inflation-induced fiscal deficits (if financed by money creation) may generate a self sustaining process. In his studies, he showed that the rate of growth of exchange rate, or, the depreciation of exchange rate had the positive impact on inflation. He also checked the causality between depreciation and inflation and found that devaluation did cause inflation, though the conclusion was subjected to the choice of lag length (Hossain, 2000). Hoque (1990) empirically examined the expectation augmented version of Phillips curve with specified attention paid toward wage rate changes and overseas price. He showed that foreign price impulse, domestic excess demand, wages and expectation of inflation are the ultimate determinants of the development of the inflation. His findings suggested that overseas prices affect the inflation rate significantly. Taslim (1982) also examined the inflationary process in Bangladesh and found that money supply and the devaluation was the root cause of inflation. Other studies like Ahsan (1974); Bose (1973); Rahim (1973); Siddique (1975) have also emphasized the monetary roots of inflation in a war-ravaged economy. Bose (1973) exerted that an increase in money supply might be taken as an approximation to the increase in money expenditure, the expansion in money supply surely exerted a very powerful inflationary pressure.

This paper would try to investigate the short-run and long-run relationship between prices, nominal wages, real income, exchange rate and domestic credit. To do this, we employ the recent technique in identifying the relationship by ARDL bound testing approach by Pesaran et al. (2001). One of the advantages of this technique is that it does not consider the order of integration of the variables. The other method such as VAR requires, the variables should be order of integration of 1. Our empirical study is based on the data from FY 1975-76 to FY 2009-2010. All of the data are collected for various issues of Economic Trends published by Bangladesh Bank, various Economic Reviews published by Ministry of Finance, Government of Bangladesh (GoB).

3. Inflation, Nominal Wage and other Macro Determinants in Bangladesh

Though we are investigating the relationship between inflation and nominal wages, sometimes nominal wages thought of as inflation allowances as wage increases only when inflation occurs (Fares and Ibrahim, 2009; Metwally and Al-Sowaidi, 2004). It means that nominal wage growth comes after the price inflation. But, it is not implausible to think that wage rate also creates an upward pressure on price level. So, it needs an empirical investigation to search for the factors that affect nominal wages. In other words, what are the factors that drive the wage and price inflation? To illustrate this question we need to look upon the trend of nominal wage rate and price index. Figure 1 shows the trend of the growth rate of CPI and Nominal wage rate from the Fiscal Year 1976 to Fiscal Year 2009. The figure clearly expresses two distinct features: as wage rate increases, then price also increases, and, we can see that there is a visual cointegration between these two variables.

We see a downward trend of both between FY 1987 to FY1993 but after FY2000, both of them are increasing. From FY1987 to FY1993, Bangladesh has stepped herself into the trade liberalization era. As Alvarez and Braun (2006) pointed out that trade liberalization are positively related to decline in aggregate price distortion. Again from FY 2000 to FY 2000 we see an upward trend in both the wage rate and CPI. Again, the growth rate of nominal wage is

higher than that of CPI. This obviates that there is a relationship between these two trends but the direction of causality is not readily apparent.

As studies on price dynamics are relatively rare, so we would look upon the studies on the determinants of inflation. Inflation in Bangladesh has varied sharply since the early 1980s. Hossain (2002) showed that during 1972-1975 the inflation rate was more than 40% per annum. This period was characterized by fixed exchange rate where Bangladesh imported inflation from anchor country. After this period, during 1990s inflation remained at about 10% per annum. Inflation decreased significantly after 1991 and remained at less than 5% per annum until 1996. This period was characterized by trade liberalization and the exchange rate of taka was adjusted frequently.

The two proponents of inflationary process-fiscal view and balance of payments view recommend choosing a number of indicators that causes inflation to any empirical analysis. Proponents of fiscal view argued that fiscal deficit is the exogenous sources of inflation as fiscal deficit occurs only when expenditure surpassed the revenues. Like other developing countries, Bangladesh has to rely on domestic credit which is inflationary and its recent economic history shows this evidently. During the FY 2002-2003 to FY 2010-11, total domestic debt to government as a percentage of total government expenditure is nearly 60-70%⁴. So, it can apparently be said that government debt plays a dominant role in creating inflationary process through budget deficit. This makes a rationale for introducing domestic credit as a determinant in our empirical analysis. On the other hand, balance-of-payments view considers inflation to be mainly driven by exchange rate movements in the presence of accommodative monetary policy. That is to say that nominal exchange rate shocks which arise from adverse external conditions cause inflation, when monetary policy is passive. Hossain (2002) empirically examines the exchange rate responses to inflation. In his study, he finds that effect of devaluation on inflation is not significant. As Bangladesh was under the IMF-World Bank-supported structural adjustment programs throughout the 1990s and 2000s, in those periods devaluation was considered as a key component of macroeconomic policy package. In those periods inflation rate per annum dropped from 11% to 5% where periodic exchange rate management was taken to adjust balance of payment problems. Again foreign capital inflows to Bangladesh played an important role for exchange rate management. Although Bangladesh turns itself into managed-floating regime in 2003, nevertheless, critics hold that devaluation/depreciation is inflationary. Considering these facts, we would consider exchange rate as a determinant for the rise in price level. Along with these variables, we would consider real GDP as a determinant to capture the positive shock from aggregate demand.

4. Data Description and Methodology

In order to find out the empirical relationship between wage and price, we consider real GDP, nominal exchange rate and domestic credit by examining the available literatures besides wage and price. We have taken the data from the Fiscal Year 1975-76 to Fiscal Year 2009-10. The variable wage is defined as the nominal wage rate index calculated by Bangladesh Bureau of Statistics (BBS) using base year 1969-70. Price is defined as the Consumer Price Index (CPI) using the base year 1995-95.⁵ Domestic Credit is defined as the total domestic credit through the banking system, Real GDP is calculated using the same base year and nominal exchange rate is

defined as the amount of taka is obtained per dollar. All of the variables are in natural logarithmic form. Data are obtained from various issues of Monthly Economics Trends, Bangladesh Bank, BBS and various issues of Economic Reviews of Ministry of Finance, Government of Bangladesh (GoB). So, there are mainly five variables: nominal wage (w), price (p), real GDP ($rgdp$), domestic credit (dom) and exchange rate (exr).

In this paper test of cointegration are carried out using the autoregressive distributive lag model (ARDL) approach due to Pesaran et al. (2001) where we estimate five unrestricted error-correction considering each variable as a dependent variable. It follows the ARDL bounds testing approach to cointegration developed by Pesaran and Pesaran (1997); Pesaran et al. (2000), and latter on by Pesaran et al. (2001). Haug (2002) has argued that ARDL approach to cointegration provides better results for small sample data set such as in our case as compared to traditional approaches to cointegration i.e. Engle and Granger (1987); Johansen and Juselius (1990) and Phillips and Hansen (1990). Another advantage of ARDL bounds testing is that unrestricted model of ECM seems to take satisfactory lags that captures the data generating process in a general-to-specific framework of specification (Laureson and Chai, 2003). This method avoids the classification of variables as I (1) and I (0) by developing bands of critical values which identifies the variables as being stationary or non-stationary processes. Unlike other cointegration techniques (e.g., Johansen's procedure) which require certain pre-testing for unit roots and that the underlying variables to be integrated are the same order, the ARDL model provides an alternative test for examining a long-run relationship regardless of whether the underlying variables are purely I (0) or I (1), even fractionally integrated. Therefore, the previous unit root testing of the variables is unnecessary. Moreover, traditional cointegration method may also suffer from the problems of endogeneity while the ARDL method can distinguish dependent and explanatory variables. Thus, estimates obtained from the ARDL method of cointegration analysis are unbiased and efficient, since they avoid the problems that may arise in the presence serial correlation and endogeneity. Note also that the ARDL procedure allows for uneven lag orders, while the Johansen's VECM does not. However, Pesaran and Shin (1999) contented that, "appropriate modification of the orders of ARDL model is sufficient to simultaneously correct for residual serial correlation and problem of endogenous variables".

A two-step procedure is used in estimating the long-run relationship. In the first step, we investigate the existence of a long-run relationship predicted by theory among the variables in question. The short and long-run parameters are estimated in the second stage, when if the long-run relationship is established in the first step.

Suppose that at the first stage, theory predicts that there is a long-run relationship among p , w , dom , $rgdp$ and exr . Without having any prior information about the direction of the long-run relationship among the variables, the following unrestricted error correction (UEC) regressions are estimated:

$$\Delta p_t = \alpha_0 + \sum_{i=1}^{\rho} \beta_i \Delta p_{t-i} + \sum_{i=0}^{\rho} \delta_i \Delta w_{t-i} + \sum_{i=0}^{\rho} \sigma_i \Delta dom_{t-i} + \sum_{i=0}^{\rho} \gamma_i \Delta rgdp_{t-i} + \sum_{i=0}^{\rho} \varphi_i \Delta exr_{t-i} + \eta_1 p_{t-1} + \eta_2 w_{t-1} + \eta_3 dom_{t-1} + \eta_4 rgdp_{t-1} + \eta_5 exr_{t-1} + e_t \quad (1)$$

The above equation shows the unrestricted version of ARDL specification.⁶ F statistic is used for testing the existence of long-run relationships. The null hypothesis for testing the nonexistence of the first long-run relationship (i.e. $H_0 : \eta_1 = \eta_2 = \eta_3 = \eta_4 = \eta_5 = 0$) is denoted by $Fp(p/w, dom, rgdp, exr)$. The F test has a nonstandard distribution which depends upon: (i) whether variables included in the ARDL model are to be I (0) or I (1), (ii) the number of regressors, and (iii) whether the ARDL model contains an intercept and/or a trend. Two sets of critical values are reported in Pesaran et al. (2001): one set is calculated assuming that all variables included in the ARDL model are I (1) and the other is estimated considering the variables are I (0). If the computed F values fall outside the inclusive band, a conclusive decision could be drawn without knowing the order of integration of the variables.

If a stable long-run relationship is supported by the first step, then in the second stage, the augmented ARDL (m, n, o, p, q) model is estimated using the following:

$$p_t = \theta_0 + \sum_{i=1}^m \xi_i p_{t-i} + \sum_{i=1}^n \psi_i w_{t-i} + \sum_{i=1}^o \zeta_i dom_{t-i} + \sum_{i=1}^p \varsigma_i rgdp + \sum_{i=1}^q \tau_i exr_{t-i} + u_t \quad (2)$$

Again the maximum of lags (ρ) in Eq. 1 must retain to determine the numbers of lags (m, n, o, p, q) in Eq. 2 selected by the Akaike Information Criterion (AIC) or Schwartz Bayesian Information Criterion (SBIC) to determine the optimal structure for the ARDL specification.

After the estimation of the ARDL (m, n, o, p, q) specification and the calculation of the associated long-run multipliers, the final step is the estimation of the short-run dynamic coefficients via the following error correction model:

$$\Delta p_t = \alpha_1 + \sum_{i=1}^m \chi_i \Delta p_{t-i} + \sum_{i=0}^n \kappa_i \Delta w_{t-i} + \sum_{i=0}^o \nu_i \Delta dom_{t-i} + \sum_{i=0}^p \pi_i \Delta rgdp_{t-i} + \sum_{i=0}^q \varpi_i \Delta exr_{t-i} + \lambda ECM_{t-1} + \vartheta_t \quad (3)$$

where ECM_{t-1} is the error correction term resulting from the verified long-run equilibrium relationship and λ is a parameter indicating the speed of adjustment to the equilibrium level after a shock. The sign of the ECM_{t-1} must be negative and significant to ensure convergence of the dynamics to the long-run equilibrium. The value of the coefficient, λ , which signifies the speed of convergence to the equilibrium process, usually ranges from -1 and 0. -1 signifies perfect and instantaneous convergence while 0 means no convergence after a shock in process.

Moreover, Pesaran and Pesaran (1997) argued that it is extremely important to ascertain the constancy of the long-run multipliers by testing the above error-correction model for the stability of its parameters. The commonly used tests for this purpose are the cumulative sum (CUSUM) and the cumulative sum of squares (CUSUMQ), both of which have been introduced by Brown et al. (1975).

The presence of a single long-run equilibrium relation entails the existence of causality in at least one direction, and for this purpose, we apply the Granger causality test (Granger 1969) augmented by the error correction term, whose inclusion is absolutely essential; otherwise, the estimates from the VAR system in first differences would not be unambiguous (Engle and

Granger 1987). The advantage of using an error correction specification to test for causality is that, on the one hand, it allows testing for short-run causality through the lagged differenced explanatory variables and, on the other hand, for long-run causality through the lagged ECM_{t-1} term. A statistically significant ECM_{t-1} term implies long-run causality running from all the explanatory variables towards the dependent variable. In our case, the multivariate p^{th} -order vector error correction model is given by the following equation:

$$\Delta x_t = \varphi + \sum_{i=1}^n \Phi_i x_{t-i} + \lambda ECM_{t-1} + e_t \quad (4)$$

where x_t is the 5×1 vector of the variables included in the model (p , w , dom , $rgdp$, exr), φ is the 5×1 vector of constant terms, Φ_i is the 5×5 matrices which include the interaction coefficients of the variables involved in (4), λ is the 5×1 vector of coefficients for each of the error correction terms and e_t is the vector of disturbance terms.

5. Empirical Analysis and Discussion of Results

5.1 Stationarity Test

Although the ARDL cointegration approach does not require unit root tests, nevertheless we need to conduct this test to ensure that none of the variables are the integrated of order 2, i.e., $I(2)$, because, in case of $I(2)$ variables, ARDL procedures makes no sense. If a variable is found to be $I(2)$, then the computed F-statistics, as produced by Pesaran et al. (2001) and Narayan (2005) can no longer be valid.

Figure 2 shows the trend of two important variables under study, wage and price, which indicates the series are upward trending. This ensures the need for unit root tests. Table 1 and Table 2 show the unit root tests of the variables. The first table shows the Augmented Dickey-Fuller tests and then second shows the Phillips-Perron test.⁷

Table 1 depicts the Dickey-fuller unit root test. The hypothesis is the series has a unit root with none and with trend. In both cases, we can clearly accept the null hypothesis that the variables have a unit root at their level. So, we need to test with their 1st differencing and this ensures that all of the variables are $I(1)$. Table 2 shows the unit root test proposed by Phillips and Perron (1990). The results resemble with the former test and with these tests we can proceed to the ARDL cointegration tests as all of the variable are $I(1)$.

5.2 Cointegration Test

The cointegration relationship between the variables p , w , dom , $rgdp$ and exr is examined using the newly developed ARDL bound testing procedure. Two steps are used in this procedure in a stepwise fashion. In the first step, the order of lags on the first differenced variables in Equations : 1–4 are obtained from the unrestricted models by using the Akaike Information Criterion (AIC) or the Schwartz-Bayesian Criterion (SBC). Besides this, researcher usually depends on literature and convention to determine the maximum lag length. As Pesaran et al. (2001) caution, there is a

delicate balance between choosing ρ in equation (1) sufficiently large to mitigate the residual serial correlation problem and, at the same time, sufficiently small so that conditional ECM is not unduly over-parameterized, particularly in view to the limited time series data which are available. So, before estimating the equation (1), we have to determine the lag-length of the 1st differenced of the variables. Table 3 shows the lag-length selection of the 1st differenced of the variables. It shows that according to AIC, the maximum lag-length for our estimation of equation is 1. After determining the lag-length, we proceed to the cointegration test. As our model is unrestricted, we omit the trend term in equations.

Table 4 represents the F statistics of estimation of equation (1) using AIC. As earlier stated that we would perform the test using each of the variables as dependent variables, so, all-in-one we would get 5 equations. We performed F test for each of the model and Table 4 shows those results. After deciding on lag-length, the issue on the selection of critical values (CVs) becomes imperative. The CVs of the F test depends on the sample sizes. Narayan (2005) argues that CVs of Pesaran et al (2001) that is generated for larger sample size should not be used for smaller sample size. Narayan (2005) presents CVs of the F test for smaller sample sizes with 30-80 observations. With 35 observations in our sample, we report both the 95% and 90% critical values from Narayan (2005) in Table 4. The long-run relationship is clearly identified for model F_p and F_{exr} . As we give importance on model F_p , it passes both at 5% and 10% level of significance as the value of F-statistics is higher than critical value. Model F_p and F_{exr} represent the long-run relationship but model F_{dom} represents a positive error correction term which gives nothing. Model F_w shows no cointegration and model F_{rgdp} is inconclusive as the value lies in the middle of lower and upper boundary.

Table 5 represents the same results when we are dealing with SBIC for cointegration test. Like Table 4, Table 5 also represents that model F_p and F_{exr} show the presence of cointegration, while model F_{dom} and F_{rgdp} give inconclusive result and model F_w gives no sign of cointegration. As our concern is on model F_p , we find that there is a long-run relationship between price and wage along with domestic credit; real GDP and nominal exchange rate.⁸ Cointegration among the variables are also checked by the test proposed by Johansen and Juselius (1990). The Phillips-Perron test indicates that all of the variables are I(1) at their levels but I(0) at their 1st differenced form, which is the pre-condition for Johansen cointegration test. This test would provide a sensitivity check on the ARDL results. Following table, Table 6 provides the likelihood ratio and Trace statistics based cointegration test. The test is based on the hypothesis that there is unrestricted intercept and no trend. This test suggests that there is at least 1 cointegration equation among the variables under study when we consider price as dependent variable. The result is significant against 5% critical value. This test confirms that there is a long-run relationship among the variables.

Based on the results in Table 4 and 5, we need to proceed only with the ARDL model having price on the left-hand-side variable (Al-Nashar, 2011). Now, we need to determine the optimal lags for the ARDL model before confronting to the estimation of long-run model to determine the coefficients of regressors and the coefficient of error correction term. Studies on ARDL cointegration often use both the AIC and SBIC criterion to select the models. To check the robustness, we represent the models based on both criteria. The long run coefficients of the selected ARDL models are presented in Table 6. Although AIC selects an ARDL (1, 0, 0, 1, 1)

model, SBIC selects ARDL (1, 0, 0, 0, 0). For the robustness of our findings, we also present the FMOLS and DOLS results using the same variables.⁹

From the above table, we can clearly see that wage, w and domestic credit, dom has significant relationship with the price, p in both criteria. These results are consistent with the earlier studies. Inspecting the results from FMOLS and DOLS, we clearly see that, in both case, wage, w plays the dominant role in determination of price dynamics, p . Though the other coefficients show the mixed results, but, all of the long-run models ensure that wage dominantly determine the price level.

Table 8 provides the error correction represent of the selected ARDL models. Panel A shows Error Correction Model (ECM) estimations having price, p as dependent variable. Panel B of the table shows the diagnostic tests of each of the model based on Table 7 and 8. Panel A in Table 8 shows both of the ARDL models. The first model, ARDL (1, 0, 0, 1, 1) which is selected on AIC, shows the short-run coefficients of each of the regressors. The variable of interest, w , clearly shows that nominal wage rate significantly affect the short-run dynamics of price, p . The model also shows that, in short-run domestic credit, dom , also affect the price level. These results are consistent with the earlier results such as Helmy (2009). To examine the fiscal view of inflation, we introduced the variable, dom . In many developing countries, like Bangladesh, mounting pressure to spend is a normal phenomenon because the tax policy and the required tax revenues to finance this spending are poor. Again inefficient financial market (by which government could finance its expenditure through issuing bond) also contributes the large budget deficits. This deficit very often financed through by money creation which could lead to persistent inflation.¹⁰ Inflation is more sensitive to government expenditure than government revenues during inflationary periods. It suggests that government's credit from domestic source for financing the budget deficit always creates pressure on inflationary trend (Hossain, 1987). The relationship between real GDP and price shows the expected sign though the result is insignificant. In the short-run, real GDP could merely affect the inflation. Ghura and Hadjimichael (1996) demonstrated that there is an inverse relationship between economic growth and inflation rate. Gylfason and Herbertsson (2001) concluded that inflation rate, above 10% to 20% per year, is generally believed to be detrimental to economic growth, where low and stable inflation has a positive role in the economic growth (Hossain et al. 1998). There is a causal relationship between the price level and real GDP. The economic intuition behind this negative relationship rests on the adverse effect on investment and productivity that inflation causes (Helmy, 2009)¹¹. Again, the short-run effect of nominal exchange rate on price is negative and insignificant. With this link, Hossain (2002) investigates the exchange rate responses to inflation in Bangladesh from the period of 1972-73 to 1999. Findings suggest that the effect of devaluation on inflation was not significant. In a deregulated system, the causal relationship between exchange rate and price is ambiguous and it could only possible if we investigate the complex structure of exchange rate responses to inflation by general equilibrium model which we are not conducting in this study. In another study, he finds that devaluation has a short-run impact on inflation but its long-run effect on inflation is not significant. This concludes that devaluation is not the cause rather effect of inflation in the long-run (Hossain, 2000). Kenen and Pack (1980) pointed out that if price levels and exchange rates are both in motion then it is hard to determine the direction of causation. It arises only when both the prices and exchange rates are endogenous and determined within the macroeconomic system. Under flexible exchange rate, a

persistent depreciation is inflationary but under pegged or adjustable peg system, a nominal devaluation represents an adjustment to the inflation. It means that the relationship between exchange rate and inflation can be established statistically but not in working hypothesis.

However, the most important term in Table 8 is the sign and coefficient of the *ECM* term. The negative sign of the *ECM* term confirms the expected convergence process in the long-run dynamics of price and other macroeconomic variables. About 40% of the last year's disequilibrium is corrected in the current year, suggesting a good speed of adjustment in the relationship process following a shock. Once a shock has occurred, it takes price level in the period from 2 to 3 years to adjust with nominal wage rate and other variables to restore its long-run relation. Banerjee et al. (1998) stated that a highly significant error-correction term is the further proof of the existence of a stable long-term relationship.

Our study takes out four diagnostic tests: (A) the Lagrange multiplier test of residual correlation (B) the heteroscedasticity test based on the regression of squared residuals on square fitted values, (C) the Ramsey Regression Equation Specification Error Test (RESET) test using the square of the fitted values. (D) the normality test based on a test of skewness and kurtosis of residuals.

The diagnostic tests suggest that the estimation of long-run coefficients and *ECM* are free from serial correlation, heteroscedasticity and non-normality at 5% level of significance regardless of whichever ARDL model we select. However, the ARDL (1, 0, 0, 1, 1) model as selected by AIC, and the ARDL (1, 0, 0, 0, 0) as selected by SBIC are not free from functional-form misspecification. This is not quite unusual. We saw some functional-form errors in Pesaran et al. (2001) as well. Pesaran et al. (2001) assert that this type of functional-form misspecification may exist due to the presence of some non-linear effects or asymmetries in the adjustment process. Işık (2010) also experienced the specification error dealing with ARDL model. The value of adjusted R^2 of the both model ranges from 70% to 74%, relatively a high adjusted R^2 , which exemplifies a good fit of the models.

The second issue addressed is the stability of the long-run coefficients that are used to form the error-correction term in conjunction with the short-run dynamics. Some of the problems of instability encountered from inadequate modeling of the short-run dynamics depart itself from long-run relationship (Bahmani-Oskooee, 2001). So, it is useful to incorporate the short-run dynamics in testing for constancy of long-run parameters. In view of this, we applied CUSUM and CUSUMQ test proposed by Brown et al. (1975) to the residuals of each model. The test CUSUM shows if the cumulative sum goes outside the two critical lines, then the model is not free from instability. Similarly, CUSUMQ, which is based on squared recursive residuals, goes outside the critical lines, and then the model is unstable. Figure 3 and 4 show the CUSUM and CUSUMQ of each model based on AIC and SBIC.

Inspection of the both figures clearly indicates both models pass the parameter stability. Pesaran and Pesaran (1997) used the same procedure to test the stability of the model. Our results clearly indicate that parameters are stable during the investigating period at 5% level of significance.

5.3 Granger Causality Test

The existence of a long-run relationship among the variables indicates that there must be at least one direction of causality to hold the existence of long-run equilibrium relationship (Engle and Granger, 1987). However, it should be noted that it does not indicate the direction of temporal causality between the variables. Thus, in order to determine the short- and long-run causality, we performed the augmented form of Granger causality test using error-correction-based Granger causality models which are weak (short-run) Granger causality and long-run Granger causality (see Tang, 2008; Acaravci & Ozturk, 2010; Ghosh, 2010). Such a test is useful for several reasons: (i) a measure of statistical significance can be found by the statistical significance of the lagged differences of variables, and (ii) the presence of long-run relationship can be found by the statistical significance of the error-correction term. It is important to remember that only equations where the null hypothesis of no cointegration is rejected (i.e., price equation, and exchange rate equation) will be estimated with an error-correction term (Narayan and Smyth, 2008; Odhiambo, 2009; Uddin et al. , 2012; Halicioglu , 2009). Another important aspect of such test is that this test differs significantly from the short-run error-correction model since it uses the same lag structure while the short-run vector error-correction model may use different lag structure (Pesaran et al. , 2001; Narayan and Smyth , 2006; Halicioglu , 2009).

Table 9 shows the short-run and long-run Granger causality test within the *ECM* framework.¹² Our first equation, the equation of our interest, price equation (F_p) shows that wage and real GDP are found to be statistically significant indicating that both wage and real GDP Granger cause price in the short-run. We found a bidirectional causal relationship between price and wage, exchange rate and wage, and exchange rate and domestic credit while we found unidirectional relationship between price and real GDP, and domestic credit and real GDP. We did not find any significant relationship between price and exchange rate in the short-run.

Turning to the long-run causality result, where we consider only price and exchange rate equation (F_p and F_{exr}), we see that the coefficient of lagged-error correction term of price equation, F_p , is negative and statistically significant exemplifies that there is a long-run relationship among the price and other variables which resembles with our bound-testing results. On the other hand, the exchange rate equation, F_{exr} , shows that the coefficient of lagged-error correction term appears insignificant which contradicts the bound test result. So, nominal wage rate along with other variables Granger cause price but there is no evidence of reverse-causality.

The long-run relationship between price and other macroeconomics variables suggest that wage and domestic credit affect the price level both in the short-run and long-run. As Blanchard (1986) said that price level dynamics are indeed the result of attempts by workers to maintain (or increase or decrease as the case may be) their real wage and by firms to maintain (or increase or decrease) their markups. Furthermore, there is a direct relation between the inflexibility of real wages and markups to shifts in demand and the degree of price level inertia. The smaller the effect of shifts in the demand for goods on the markup, and the smaller the effect of shifts in the demand for labor on the real wage, the more slowly will the nominal price level adjust to offset aggregate demand disturbances. He also added an example and explained that after the initial increase in nominal money, the monetary authority attempts to maintain the higher level of output by partially accommodating the increase in nominal prices. In this case, the higher the degree of accommodation, the stronger and the longer lasting the wage price spiral. The

monetary authority is indeed able to maintain output at a higher level for a longer period of time but at the cost of a higher initial increase in prices and a longer period of inflation. Though Gordon (1988) showed the evidence through Granger-causality test that changes in wages and prices are not related, but, other empirical works (Barth and Bennett, 1975; Mehra, 1977) found that wages and prices are related with Granger-causality running either in both directions or only prices to wages. Mehra (1988) empirically tested the relationship between the rate of changes in prices and the unit cost of labor and found that they were cointegrated which are in line with our results.

Some authors like Taslim (1982); Hossain (2000); Mujeri et al. (2009) etc also examined the inflationary pressure through the view of both structuralist and monetarist views. They all used the money supply as a determinant of increase in price level. But Metin (1998) assumed that to factor out the causes of inflation, it is more reasonable to use base money or domestic credit instead of money supply, as, in many cases money supply may render itself as endogenous. We find that domestic credit is also responsible for creating upward pressure on price both in short-run and long-run. The recent evidence also supports our hypothesis. This mainly stem from the budget-deficit effect on inflation (Metin, 1998; Miller, 1985; Wijnbergen et al., 1989). As Friedman (1974) said that inflation is an old diseases and the cost of reducing inflation is unemployment and the reduction in output. But he also asserted that the cost of the later is not so far than the cost of former.

6. Concluding Remarks

This study has intended to assess how wages and prices interact with /and affect each other. But price and wage dynamics are at the heart of macroeconomic developments induced by changes in aggregate demand and/or aggregate supply. And so, examining price and wage inflation cannot be separated from the developments that take place in the various macro indicators. Therefore, the study has tested empirically the relationship between prices and wages, while accounting for the most important macro variables that contribute to price inflation dynamics in Bangladesh. As inflation or rise in price levels exacerbate the real income, so, examining the factors of rise in price is necessary for not only policy formulation but also to restrain the people from embracing themselves to poverty.

To this line, we examined the price-wage relationship along with other macroeconomic determinants which also affect the aggregate price level. We investigated the short-run and long-run relationship between price and nominal wage along with domestic credit and real GDP with ARDL bound testing approach of cointegration. These findings are in line with the previous studies (Metin, 1988; Adu and Marbuah, 2011; Montiel, 1989). In the short-run we found that real GDP is positively related to price level but in the long-run it appears negative though it is insignificant. A negative coefficient of real GDP on price indicates the favorable effect of easing capacity constraints towards curbing upward pressures on the domestic price level.

Investigating the cointegration by assuming each variable as dependent, we test five models and found only two functions which show the sign of cointegration: our base line price equation, F_p and exchange rate equation, F_{exr} . As more than one equation shows the cointegration so, it definitely urges for causality test. We performed the augmented form of Granger causality test

using error-correction-based Granger causality models which are weak (short-run) Granger causality and long-run Granger causality. We find that nominal wage rate Granger cause the price where there is a bidirectional relationship between wages and price. This proves the traditional wage-price spiral relationship explained by Blanchard (1987). We also find that there is a long-run relationship between price and other variables while we find no significant long-run relationship between exchange rate and other variables, which is one of the striking results in this study. Moreover, the long run relationship is also established by FMOLS and DOLS method, which show that both the wage and domestic credit play important role in determining price level.

Although many other factors directly and indirectly affect the price level, so, it needs an empirical investigation to factor out the determinants. As Agénor et al. (1997) empirically established that wage is important for understanding inflationary process in developing countries. Again they asserted that wage is also important for short-run fluctuations in output as wage movements affect both the supply side and demand side of the economy. In such a situation, developing countries like Bangladesh should give importance on the wage-price relationship. It is to be noted that price stability is one of the main goals of monetary policy. One of the limitations of this study is to disregard the rate of unemployment where we have no time series data on this variable. Again, we disregard the money supply, an important variable for the determinant of rising price level. We left this as many authors considered as endogenous. As price is considered as signal, so the effect of wage on price should be explored by general equilibrium analysis as price is affected by a number of macroeconomic variables. Again, we did not incorporate any shock to this bound testing approach which is kept for further research.

Endnotes

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1. As Milton Friedman said, "I always speak of the change in money supply as a proximate cause, and say that deeper causes must be found in what are the explanations for the rise in money supply."

2. A sizable empirical literature has also emphasized the possibility (again in the presence of accommodating policy stance) of vicious circles between inflation and public sector deficits, on the one hand, and between the exchange rate and prices, on the other. An inflation-budget deficit spiral may emerge if inflation itself has an adverse effect on the fiscal deficit, by raising expenditure faster than the revenue--as emphasized by Olivera-Tanzi effect. Likewise, a devaluation-inflation spiral may occur, if, following an initial exchange rate adjustment and an increase in domestic prices, policymakers attempt to maintain the real exchange rate constant to prevent adverse affect of competitiveness and external accounts. See, for instance, the evidence for Turkey discussed by Onis and Ozmucur (1990).
3. see Bangladesh Economic Review (2012).
4. see Bangladesh Economic Review (2012).
5. BBS and Bangladesh Bank did not publish general consumer price index (national) from 1971-72 to 1985-86. So, for all those periods, we use the estimated CPI (national) from Rahman (1994). We then extended our data set to 2009-2010 by changing base year to 1995-96 using the data from BBS and Bangladesh Bank and calculated the CPI General for national.
6. for the sake of space, we did not show the other models, where each of the variables appeared to be as dependent variables. The result of such types of models are presented in later table.
7. See Dickey and Fuller (1979, 1981), Phillips and Perron (1988).
8. We will address the causality of model F_p and F_{exr} of both table, Table 4 and 5 later.
9. Fully Modified OLS (FMOLS) is based on the assumption of equal weights, truncation lag=1, Non-trended case and Dynamic OLS (DOLS) is based on 2 period lead and lag using Newey standard errors.
10. Bangladesh experienced several episodes of inflation during the entire history. War-torn economy faced an inflationary periods during the early stages of 1980s. Large trade and budget deficit also ignite the inflationary phases in recent years. In each fiscal year, the ADP remain unused in most cases and before the ending of fiscal year, excessive disbursement of ADP creates inflation. Again, supply shock very often is a major determinant of inflation in Bangladesh. Over all, budget deficit is inflationary (Choudhary and Parai, 1991; Neyapti, 2003; Narayan and Narayan, 2005).
11. Helmy (2009), Abou-Ali et al. (2008) etc found the relationship between inflation and real growth to be actually non-linear; and detected the presence of a threshold. Meaning that there exists a threshold level below which inflation is considered a desired phenomenon, but above that threshold level, inflation becomes detrimental to growth. Below the threshold, inflation is compatible with a growing economy and resilient demand. Above the threshold, higher

inflation increases inflationary expectations and erodes the real value of returns on investment.

12. We present the coefficient of 1st differenced variable only. All of the models are based on AIC. The lag-length is based on AIC. We also performed the test based on SBIC. For the sake of place, we did not show the results. The results will be shown upon request.

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Table1: Unit Root Test(Dickey Fuller Test)

| Variable | Level | | 1st Difference | | 1st Difference | | 1st Difference | |
|----------|----------|----------|----------------|----------|----------------|----------|----------------|----------|
| | None | Decision | With Trend | Decision | None | Decision | With Trend | Decision |
| p | -3.251** | I(0) | -1.626 | I(1) | -2.796* | I(0) | -3.436* | I(0) |
| w | -2.491 | I(1) | -1.899 | I(1) | -3.358** | I(0) | -3.902** | I(0) |
| dom | -1.906 | I(1) | -2.271 | I(1) | -4.028*** | I(0) | -4.329*** | I(0) |
| rgdp | -0.400 | I(1) | -2.143 | I(1) | -5.743*** | I(0) | -5.686*** | I(0) |
| exr | -1.410 | I(1) | -0.910 | I(1) | -3.509** | I(0) | -3.656*** | I(0) |

N.B: *, ** and *** denote 10%, 5% and 1% level of significance respectively. The values are given in the cells are calculated values of the regressions. The null hypothesis is the series has a unit root which is tested against Mackinnon Critical values.

Table2: Unit Root Test(Phillips-Perron Test)

| Variable | Level | | 1st Difference | | 1st Difference | | 1st Difference | |
|----------|--------|----------|----------------|----------|----------------|----------|----------------|----------|
| | None | Decision | With Trend | Decision | None | Decision | With Trend | Decision |
| p | -2.513 | I(0) | -1.602 | I(1) | -2.793* | I(0) | -3.422* | I(0) |
| w | -2.112 | I(1) | -1.876 | I(1) | -3.486** | I(0) | -3.993** | I(0) |
| dom | -1.704 | I(1) | -2.261 | I(1) | -4.045*** | I(0) | -4.334*** | I(0) |
| rgdp | -0.372 | I(1) | -2.195 | I(1) | -5.745*** | I(0) | -5.687*** | I(0) |
| exr | -1.286 | I(1) | -1.316 | I(1) | -3.397** | I(0) | -3.472* | I(0) |

N.B: *, ** and *** denote 10%, 5% and 1% level of significance respectively. The values are given in the cells are calculated values of the regressions. The null hypothesis is the series has a unit root which is tested against Mackinnon Critical values.

Table3: Lag-length selection of the 1st differenced of the variables

| Lag | LL | LR | df | p | FPE | AIC | HQIC | SBIC |
|-----|---------|--------|----|-------|----------|-----------|-----------|-----------|
| 0 | 214.057 | | | | 9.6e-13 | -13.4876 | -13.4122 | -13.2563* |
| 1 | 254.077 | 80.03 | 25 | 0.000 | 3.7e-13* | -14.4566* | -14.0042* | -13.0689 |
| 2 | 276.108 | 44.06 | 25 | 0.011 | 5.2e-13 | -14.265 | -13.4357 | -11.7209 |
| 3 | 290.704 | 29.19 | 25 | 0.256 | 1.5e-12 | -13.5938 | -12.3875 | -9.89319 |
| 4 | 326.193 | 70.09* | 25 | 0.000 | 1.9e-12 | -14.2705 | -12.6872 | -9.41346 |

N.B: * indicates the lowest value under each criteria.

Table 4: Cointegration Test [based on AIC]

| Model | LHS variable | Forcing variable | F Statistics | 95% Critical bounds (with no trends) | | 90% Critical bounds | | Cointegration |
|---------------|--------------|----------------------------|----------------|--------------------------------------|-------|---------------------|-------|--------------------------|
| | | | | I(0) | I(1) | I(0) | I(1) | |
| 1. F_p | <i>cpi</i> | <i>w, dom, exr, rgdp</i> | 34.0650 | 3.276 | 4.630 | 2.696 | 3.898 | Present |
| 2. F_w | <i>w</i> | <i>cpi, dom, exr, rgdp</i> | 0.47978 | 3.276 | 4.630 | 2.696 | 3.898 | No Cointegration |
| 3. F_{dom} | <i>dom</i> | <i>cpi, w, exr, rgp</i> | ----- | 3.276 | 4.630 | 2.696 | 3.898 | Not Present ^a |
| 4. F_{exr} | <i>exr</i> | <i>cpi, w, dom, rgdp</i> | 8.6681 | 3.276 | 4.630 | 2.696 | 3.898 | Present |
| 5. F_{rgdp} | <i>rgdp</i> | <i>cpi, w, dom, exr</i> | 2.8373 | 3.276 | 4.630 | 2.696 | 3.898 | Inconclusive |

N.B: F statistics are bold when they are significant at the 5% level. The null hypothesis is $\eta_1 = \eta_2 = \eta_3 = \eta_4 = \eta_5 = 0$. They represent the coefficients of the lagged levels as per equation (1). Critical bounds for the F statistics are applicable for 35 observations with unrestricted intercept and no trend. See case III in Narayan (2005), p. 1988 for details. Inconclusive means if the value of F-statistics lie between the upper and lower critical value.. All of the models are assuming there is no trend and intercept.

^a Model 3 represent positive coefficient of error correction term of coinegration test, so, it is referred to no cointegration.

Table 5: Cointegration Test [based on SBIC]

| Model | LHS variable | Forcing variable | F Statistics | 95% bounds (with trends) | Critical (with no trends) | 90% bounds | Critical | Cointegration |
|---------------|--------------|----------------------------|----------------|--------------------------|---------------------------|------------|----------|------------------|
| | | | | I(0) | I(1) | I(0) | I(1) | |
| 1. F_p | <i>cpi</i> | <i>w, dom, exr, rgdp</i> | 75.3343 | 3.276 | 4.630 | 2.696 | 3.898 | Present |
| 2. F_w | <i>w</i> | <i>cpi, dom, exr, rgdp</i> | 0.51514 | 3.276 | 4.630 | 2.696 | 3.898 | No Cointegration |
| 3. F_{dom} | <i>dom</i> | <i>cpi, w, exr, rgp</i> | 4.2293 | 3.276 | 4.630 | 2.696 | 3.898 | Inconclusive |
| 4. F_{exr} | <i>exr</i> | <i>cpi, w, dom, rgdp</i> | 8.6681 | 3.276 | 4.630 | 2.696 | 3.898 | Present |
| 5. F_{rgdp} | <i>rgdp</i> | <i>cpi, w, dom, exr</i> | 2.8373 | 3.276 | 4.630 | 2.696 | 3.898 | Inconclusive |

N.B: F statistics are bold when they are significant at the 5% level. The null hypothesis is $\eta_1 = \eta_2 = \eta_3 = 0$. They represent the coefficients of the lagged levels as per equation (1). Critical bounds for the F statistics are applicable for 35 observations with unrestricted intercept and no trend. See case III in Narayan (2005), p. 1988 for details. Inconclusive means if the value of F-statistics lie between the upper and lower critical value. All of the models are assuming intercept and there is no trend.

Table 6: Johansen Tests for Cointegration

| Maximum Rank | Parameters | LL | Eigenvalue | Trace Statistics | 5% Critical Value |
|--------------|------------|-----------|------------|------------------|-------------------|
| 0 | 30 | 265.20903 | | 71.7328 | 68.52 |
| 1 | 39 | 283.01717 | 0.66128 | 36.0074* | 47.21 |
| 2 | 46 | 291.41605 | 0.39693 | 19.3187 | 29.68 |
| 3 | 51 | 298.144 | 0.33489 | 5.8628 | 15.41 |
| 4 | 54 | 301.074 | 0.16270 | 0.0034 | 3.76 |
| 5 | 55 | 301.075 | 0.00007 | | |

N.B: * denotes 5% level of significance. The Table indicates that there is at least 1 cointegration equation based on Trace statistics.

Table 7: Estimated Long Run Coefficients using the ARDL Approach based on AIC and SBIC [Dependent variable, p]

| | Model: ARDL(1, 0, 0, 1, 1): AIC | Model: ARDL(1, 0, 0, 0, 0): SBIC | FMOLS | DOLS |
|----------|---------------------------------------|--|--------------------|--------------------|
| Variable | Coefficient (Prob) | Coefficient (Prob) | Coefficient (Prob) | Coefficient (Prob) |
| w | 0.51183(0.000)*** | 0.49647(0.000)*** | 0.6945(0.000)*** | 1.03(0.031)** |
| dom | 0.2129(0.000)*** | 0.20307(0.000)*** | -0.0592(0.216) | -0.2034(0.479) |
| rgdp | -0.0065(0.667) | -0.0045(0.759) | 0.0367(0.005)*** | 0.0675(0.09)* |
| exr | -0.1278(0.420) | -0.081270(0.585) | 0.3108(0.000)*** | 0.1285(0.469) |
| F-stat | 34.065 | 75.3343 | | |

N.B: *, ** and *** indicates 10%, 5% and 1% level of significance respectively.

Table 8: Error Correction Representation of the Selected Model[Δp_t is dependent variable]

| Panel A | | | | | | | | |
|---|----------------------------|----------------|-----------------------------|----------------|-----------------------------|-------------------------|-----------------------------|----------------|
| | Regressor | | | | | | | |
| ARDL (1, 0, 0, 1, 1) | Δw_t | Δdom_t | $\Delta rgdp_t$ | Δexr_t | ECM_{t-1} | Adjacent R ² | SBC | AIC |
| Coefficient | 0.206 | 0.0857 | 0.0136 | 0.0720 | -0.40264 | 0.738 | 77.87 | 83.21 |
| Standard error | 0.0249 | 0.0173 | 0.0112 | 0.0791 | 0.069341 | | | |
| p- value | 0.000 | 0.000 | 0.233 | 0.370 | 0.000 | | | |
| $ECM = p -0.51183*w -0.21294*dom + 0.0065095*rgdp + 0.12787*exr$ | | | | | | | | |
| ARDL (1, 0, 0, 0, 0) | | | | | | | | |
| Coefficient | 0.2023 | 0.083 | -0.0018 | -0.03312 | -0.40754 | 0.694 | 78.74 | 82.55 |
| Standard error | 0.0258 | 0.0177 | 0.0061 | 0.056 | 0.07007 | | | |
| p- value | 0.000 | 0.000 | 0.763 | 0.565 | 0.000 | | | |
| $ECM = p -0.49647*w -0.20370*dom + 0.0045945*rgdp + 0.081270*exr$ | | | | | | | | |
| Panel B | | | | | | | | |
| | ARDL (1, 0, 0, 1, 1) AIC | | | | ARDL (1, 0, 0, 0, 0) SBIC | | | |
| | <i>F version</i> | <i>p-value</i> | <i>LM Version</i> | <i>p-value</i> | <i>F version</i> | <i>p-value</i> | <i>LM Version</i> | <i>p-value</i> |
| (A) Serial Correlation | F(1, 26) = 0.02 | 0.815 | $\chi^2(1) = 0.07$ | 0.787 | F(1,28) = 1.04 | 0.315 | $\chi^2(1) = 1.22$ | 0.269 |
| (B) Heteroscedasticity | F(1, 32) = 0.01 | 0.900 | $\chi^2(1) = 0.02$ | 0.896 | F(1,32) = 0.41 | 0.523 | $\chi^2(1) = 0.43$ | 0.508 |
| (C) Functional Form | F(1,26) = 4.59 | 0.042 | $\chi^2(1) = \mathbf{5.10}$ | 0.024 | F(1,28) = 7.62 | 0.10 | $\chi^2(1) = \mathbf{7.27}$ | 0.007 |
| (D) Normality | NA | | $\chi^2(2) = 1.99$ | 0.369 | NA | | $\chi^2(2) = 1.23$ | 0.540 |

N.B: Δ denotes the first order difference operator. Coefficients are bold when they are significant at the 5% or 1% level. AIC Akaike Information Criterion, SBIC Schwarz Information Bayesian Criterion . Here higher values of the criteria are preferred to lower ones by design. The nulls are (A) No serial correlation, (B) No Heteroscedasticity (C) No functional-form misspecification and (D) No non-normal errors for the respective tests. For details of diagnostics tests see Pesaran and Pesaran (1997), Ch. 18.

NA = Not applicable

Table 9:Granger Causality Test Results:

| Dependent variable | Short-run (or weak) Granger Causality Test | | | | | | Long-Run Granger Causality Test |
|--------------------|--|----------------|----------------|-----------------|-----------------|----------------|---------------------------------|
| | Δp | Δw | Δdom | $\Delta rgdp$ | Δexr | ECM_{t-1} | |
| Δp | - | 0.169(0.021)** | 0.066(0.310) | 0.023(0.006)* | -0.009(0.935) | -0.253(0.073)* | |
| Δw | 0.923(0.071)* | - | -0.054(0.760) | -0.026(0.217) | -0.438(0.098)* | - | |
| Δdom | 0.508(0.277) | -0.077(0.742) | - | 0.081(0.000)*** | -0.640(0.036)** | - | |
| $\Delta rgdp$ | 5.506(0.194) | -1.117(0.308) | 2.443(0.166) | - | 1.995(0.252) | - | |
| Δexr | -.032(0.935) | -0.272(0.095)* | -0.281(0.064)* | 0.029(0.101) | - | -0.338(0.214) | |

N.B: Figures in parenthesis indicate p-values. *, ** and *** represent 10%, 5% and 1% level of significance respectively.

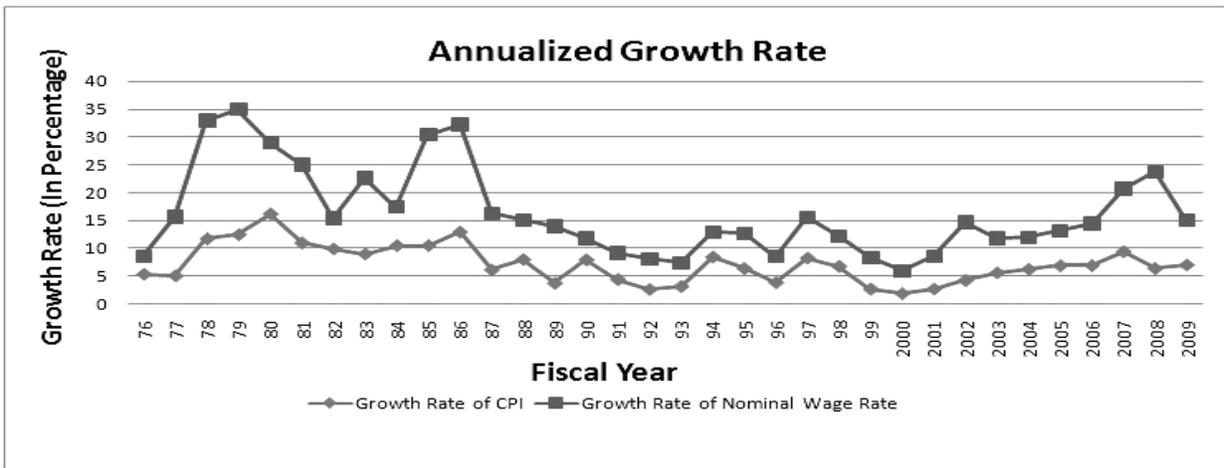


Figure 1: Annualized Growth rate of CPI and Nominal Wage Rate(Author's Calculation)

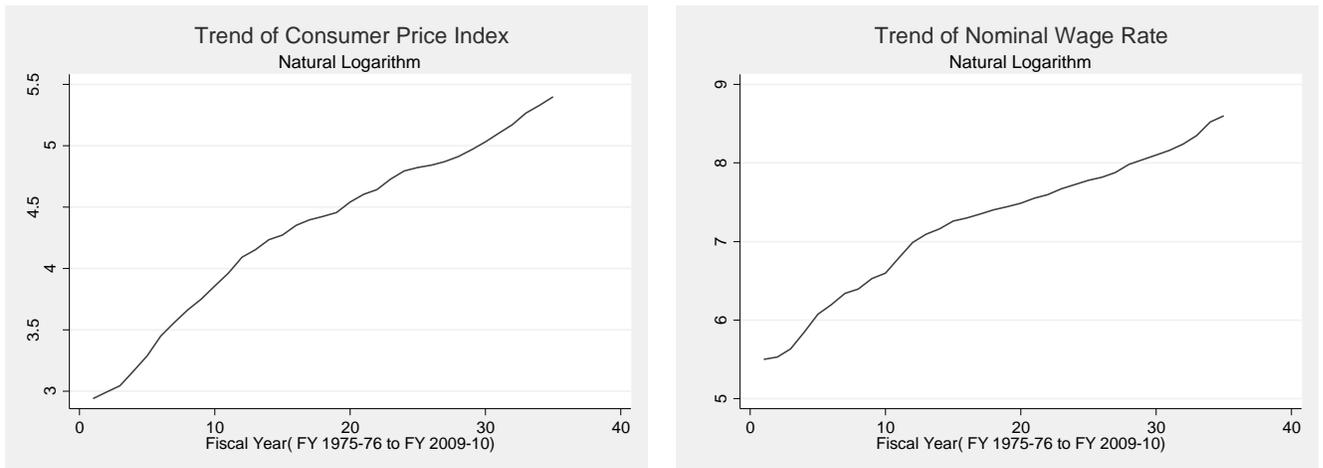


Figure 2: Trends of Nominal Wage Rate Index and Consumer Price Index

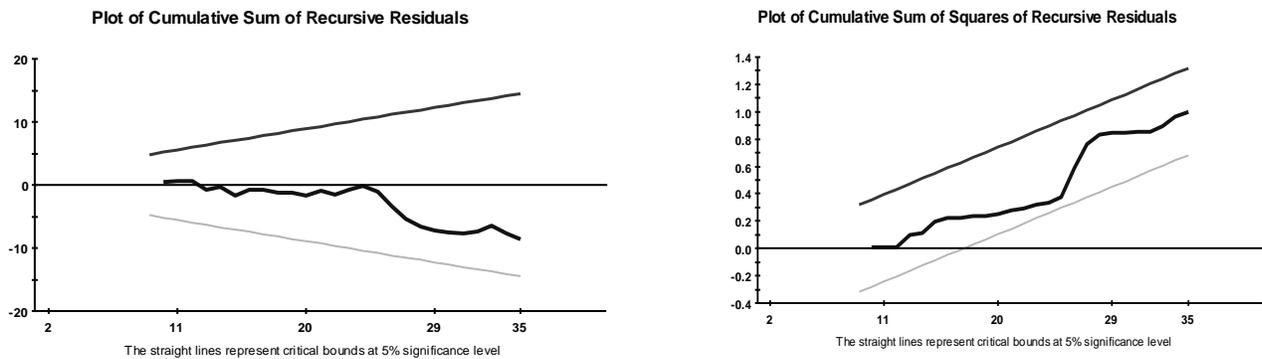


Figure 3: (a) CUSUM (b) CUSUMQ based on AIC[ARDL(1, 0, 0,1, 1)]

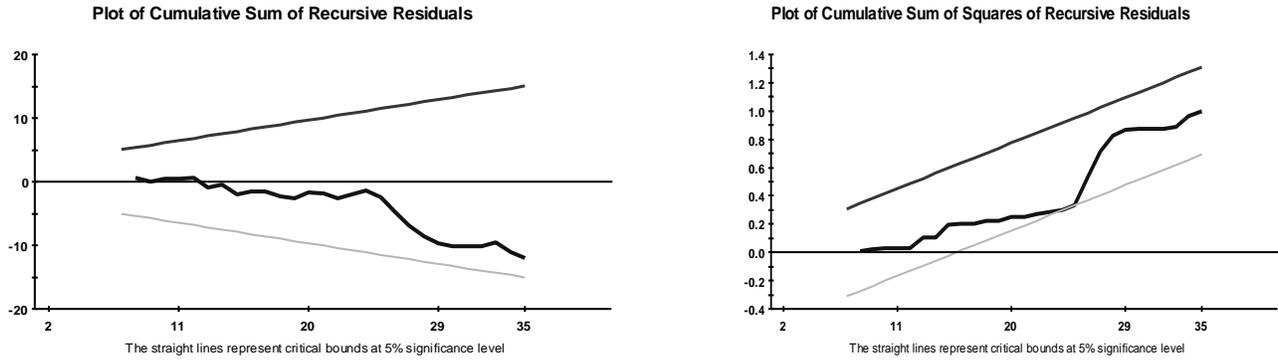


Figure 4: (a) CUSUM and (b) CUSUMQ based on SBIC [ARDL(1, 0, 0, 0, 0)]