INFLATION BEFORE AND AFTER CENTRAL BANK INDEPENDENCE: THE CASE OF COLOMBIA ♦

Jesús Otero ♣
Facultad de Economía
Universidad del Rosario, Colombia
✉ jotero@claustro.urosario.edu.co

Manuel Ramírez
Facultad de Economía
Universidad del Rosario, Colombia
✉ maramire@claustro.urosario.edu.co

ABSTRACT

In this paper we model the Colombian inflation rate in terms of excess demand effects from asset, goods and factor markets. In contrast to previous results for a group of industrial economies, we find that domestic factors are a far more powerful influence on inflation than are external factors. The paper pays particular attention to the potential effects of the Constitutional Reform of 1991, which created a Central Bank independent from other parts of government. We find that the creation of an independent Central Bank did change some of the parameters of the model, as the disequilibria in goods and monetary markets were found to have a smaller effect on inflation after Central Bank independence was granted.

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♣ Corresponding author
Jesús Otero
Facultad de Economía
Universidad del Rosario
Calle 14 # 4-69
Bogotá, Colombia
☏ (+571) 297 02 00 Ext. 661
🖷 (+571) 344 57 63

1. Introduction

One of the main subjects of concern for policymakers and economists alike is the behaviour of inflation. Empirical evidence for a large number of countries reveals that high and variable rates of inflation are not consistent with sustained economic growth, because they shorten the planning horizon of investors and reduce the rate of productivity growth in the economy (see e.g. Fisher, 1993; Barro, 1995). Understanding the main factors that affect the dynamics of inflation is thus crucial to help policymakers design measures to achieve a stable macroeconomic environment, and gain insight about the effects of their policies.

Economic theory suggests alternative views to explain the sources of inflation. A first view is associated with the monetarist school, according to which the main cause of inflation lies in expansions of the money supply in excess of real productivity growth. A second view focuses on the external factors that affect the domestic price level in an open economy, either through the transmission of import prices inflation in foreign currency terms into domestic inflation, or through the influence of the exchange rate on prices (via prices of imported intermediate and final goods). A third view lays emphasis on internal theories, which may be further subdivided into labour market theories and excess demand theories. The former highlights the role of the wage, being the result of labour demand and supply interactions, as a component of producers costs, while the latter refers to excess demand pressure effects. There has recently been an increasing interest in studying the effect of institutional factors, like the impact on inflation rate resulting from the degree of Central Bank independence from other branches of government (see e.g. Alesina and Summers, 1993; Cukierman, 1992).

This paper investigates the determination of the rate of inflation in Colombia in terms of the explanations mentioned above. The basic idea of the paper is that inflation can be associated with excess money supply, demand pressure effects, imported inflation, and wage inflation. The analysis of the main determinants of inflation in Colombia has certainly been a topic of dynamic research throughout the years; see e.g. Misas et. al. (1999) and the references therein. To our knowledge, however, existing literature has not investigated inflation dynamics accounting for all explanations, as it has mainly focussed on only one of the possible origins of inflation. In addition, research on Central Bank independence has typically been addressed within the context of cross section or panel data, but not within a pure time series model as the one we present in this paper.[1]

Our empirical modelling exercise is implemented in two steps. First, we use multivariate cointegration techniques to test for the existence of long-run equilibrium relationships in three different systems of equations, describing the monetary sector, the foreign sector, and the labour sector. Second, the deviations from the estimated cointegrating relationships are included as determinants in an inflation model. In other words, we expect the inflation rate to adjust to deviations from the long-run cointegrating relationships derived in the first step of the analysis. Modelling inflation in terms of disequilibria from several sectors is important in numerous studies of other economies; see e.g. Surrey (1989) for the US and the UK, Juselius (1992) for Denmark, and Hendry (2000, 2001) for the UK, among others. Surrey (1989) and Juselius (1992) find that, within the industrial economies considered, the external influence on domestic prices is a far more powerful influence compared to the domestic influence. Hendry (2000, 2001) finds that most theories of inflation help explain UK inflation over the last century and a quarter, providing no support for any "single-cause" explanation (in particular, excess money is not found to play a key role). The study of the Colombian case is therefore interesting, because it enables us to determine the importance of internal, external and institutional influences on domestic prices within the context of a developing economy.

Our paper differs in one important aspect from these earlier works. The analysis of the

[1] See e.g. Berger et. al. (2001) for a review of recent research on central bank independence.

Colombian experience allows us to assess the effects of the Constitutional reform of 1991, which radically modified the structure and functions of the Central Bank with the aim of creating an institution independent from the government's executive branch. The Constitution of 1991 created the Board of Directors of the Central Bank, which consists of seven members: the Minister of Finance who presides the Board without veto power; five Co-Directors who are appointed by the President and serve for a minimum period of four years and a maximum of twelve; and the Governor of the Bank who is elected by the Co-Directors for a minimum period of four years and a maximum stay of twelve. The terms of the Co-Directors are staggered so that no President can appoint the entire Board at any time. The Constitution stated that the main objective of the Bank was to control inflation, and that it had to coordinate its policies with government macroeconomic policies . The Constitution also forbids the Bank from lending to non-financial private agents - Loans to the government are only permitted after an unanimous vote by the Board, something which has not happened up to now. The institutional changes introduced by the Constitution of 1991 have now been in place long enough to allow for some hope of identifying the effects they may have had on inflation.

Colombian inflation was on the order of 25% per year in the 1970s and 1980s. This, according to Dornbusch and Fischer (1993), constitutes a classical example of a moderate but persistent inflation process. Since the 1991 reform the inflation rate has been exhibiting a sustained declining path, and is now in the one-digit inflation ground (about 7.5%). Gómez et. al. (2002) and Restrepo (2000) associate this marked change of the inflation rate with the high level of Central Bank independence. Our modelling exercise indicates that the independence of the Central Bank did not affect the autonomous level of inflation, but it changed the response of inflation to disequilibria in the goods and money markets. In particular, the disequilibria in the goods and money markets are found to have a larger effect on inflation before Central Bank independence, suggesting that monetary policy simply accommodated these disequilibria.

The paper is organised as follows. Section 2 presents the cointegration properties of three different systems of equations, describing first the monetary sector, then the foreign sector, and finally the labour sector. Section 3 estimates a model for the determination of the inflation rate in Colombia using results derived from the cointegration analysis. The model allows for the potential impact on inflation of the creation of an independent Central Bank. Section 4 concludes.

2. The analysis of long-run equilibrium relationships

Our empirical analysis of the determinants of the inflation rate in Colombia begins with the investigation of the long-run data structure. In particular, we focus on the cointegration properties of three systems of equations, describing the monetary sector, the foreign sector, and the labour sector. For each of these sectors, we apply the Johansen (1988, 1995) full information maximum likelihood procedure, as it allows for the estimation of multiple cointegrating vectors that can be interpreted as equilibrium relationships among the variables.

The Johansen procedure is based on a Vector Error Correction (VEC) model:

$$\Delta Y_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i} + \Pi Y_{t-1} + \Psi D_t + \varepsilon_t, \qquad t = 1,...,T, \tag{1}$$

where Y_t is a set of j endogenous variables, μ is a drift parameter, D_t is a set of centred seasonal dummy variables, and $\varepsilon_t \sim iid(0, \Sigma)$. Here Π is a $(j \times j)$ coefficient matrix, which in the presence of cointegration can be factorised according to the number r of linearly independent cointegrating vectors:

$$\Pi = \alpha\beta',$$

where α and β are both $(j \times r)$ matrices of full rank, with β containing the r cointegrating vectors and α carrying the corresponding loadings in each of the r vectors. In the Johansen procedure a test for the null hypothesis of r cointegrating relations is equivalent to a test of the hypothesis that Π has less than full rank. Next, we present the cointegration analysis for the three sectors mentioned above.

2.1. The monetary sector

In this section we postulate a small monetary system of the Colombian economy with the purpose of obtaining a measure of excess money, which will then be used to analyse the effect of an inflationary monetary policy on prices. In particular, we estimate such measure of money market disequilibrium as the residuals of a long-run money demand equilibrium relationship, assuming that the money supply is given.

Our starting point is thus a standard log-linear specification of the demand for money:

$$m^d = a_1 y + a_2 p - a_3 R,$$

where m^d is money demand in nominal terms, y is a measure of the volume of real transactions, p is an appropriate price level, and R is an interest rate on the alternatives of money. This specification is a summary way of representing the primary motives for holding money: transactions, precautionary and speculative; see e.g. Goldfeld and Sichel (1990) for a review of economic theories of money demand.

In the empirical analysis below, m^d, y, and p, are nominal $M1$, Gross Domestic Product (in prices of 1994), and the consumer price index, respectively. R is the yield of 90-day CDT certificates offered by banks and financial corporations representing opportunity costs for holding money.[2] The money, output and price series are considered in logarithms and denoted

[2] We also tried to estimate a model using the $M3$ definition of money, with the opportunity cost measured as the spread between the own rate on money and the money alternative. The own rate is calculated as a weighted average of the interest rates on the individual components of $M3$. As for the money alternative, the foreign interest rate is incorporated in the model to represent the opportunity cost of holding domestic currency with respect to foreign currency. In this case, however, we did not find evidence of cointegration.

$m1_t$, y_t, and p_t; the interest rate series (R_t) is not considered in logarithms in order to allow the interest rate elasticity to vary with the level of the interest rate. Data are seasonally unadjusted quarterly observations from 1980:1 to 2000:3 (see the appendix for definitions and sources of the data used in the paper).

Figure 1 plots the series in levels and in first differences. Visually all series appear $I(1)$ at least, and preliminary analysis of the data using the augmented Dickey-Fuller (1979, 1981) [ADF] tests supports this.[3]

The empirical model for the monetary sector is the VEC model (1) for the set of endogenous variables $Y_t = [m1_t, y_t, p_t, R_t]'$. The specification of the deterministic component of the model includes an unrestricted constant, centred seasonal dummy variables, and a linear trend that is restricted to lie in the cointegrating space as a proxy for financial innovations.[4] The model is estimated over the whole sample period 1980:1 to 2000:3, and the lag length is selected by starting with $k = 5$ lags, and sequentially testing from the highest order using Likelihood Ratio tests statistics. The final model had a lag order of four (the Akaike information criterion gave the same number of lags), and this specification was then subjected to diagnostic checking: serial correlation, ARCH effects and normality (see Table 1a).[5] These test statistics show no sign of misspecification, except for a minor problem with normality in the price equation which is failed at a five per cent significance level.

Johansen (1988, 1995) cointegration tests are shown in Table 2a, which reports the λ_i eigenvalues, the λ-*max* and the *trace* statistics. Both statistics, adjusted for degrees of freedom to take into account the small sample bias and lag structure (see Reimers, 1992), show evidence of two cointegrating vectors. Using the long-run structural modelling techniques advanced in Pesaran and Shin (2002), the following vectors were identified:

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[3] The ADF test results are not reported here to save space, but are available from the authors upon request.
[4] See e.g. Arrau et. al. (1995) on the use of a deterministic trend to capture the role of financial innovation.
[5] Estimations are done in PcGive and PcFiml 9.0 (see Hendry and Doornik, 1997).

$$m1_t = y_t + p_t \underset{(0.146)}{-0.715R_t} \underset{(0.0004)}{-0.005Trend}$$

and

$$y_t = \underset{(0.037)}{0.598m1_t} - \underset{(0.041)}{0.482p_t},$$

where standard errors are given in parentheses.[6] The first cointegrating vector (denoted Mm_t) is interpreted as a long-run money demand equation which satisfies income and price homogeneity, so that the residuals from this equation can be thought of as a measure of money market disequilibrium, assuming that the money supply is given. The second cointegrating vector (denoted Gm_t) can be interpreted as a proxy for aggregate demand with positive real money effects.[7]

Figures 2a and 2b plot the two identified restricted cointegrating vectors partialling out (i.e. correcting for) the full-sample short-run dynamics (see Johansen and Juselius, 1992). The two cointegrating vectors appear stationary from visual inspection. Further inspection of the graphs reveals that the magnitude of the disequilibrium in the money market has fluctuated within a range of approximately ± 10 per cent, and much smaller disequilibria are observed for the excess demand vector (approximately ± 5 per cent).

2.2. The foreign sector

In this section we turn our attention to the international aspects of inflation. There are two basic channels by which external factors directly affect the domestic price level. The first

[6] For exact identification we imposed a unit coefficient on $m1$ and long-run price homogeneity on the first vector, and a unit coefficient on y and long-run exclusion of the trend term on the second vector. Having imposed these exactly identifying restrictions, we then tested the validity of a proportional effect between y and $m1$ on the first vector, and long-run exclusion of R on the second vector, producing a $\chi^2(2) = 1.005$ (*p-value* = 0.605).

[7] Notice that in the second vector the estimated coefficients of $m1$ and p are almost of the same magnitude and with opposite sign. However, testing this hypothesis along with income homogeneity (on the first vector) and long-run interest rate exclusion (on the second vector) yields a significant value of the Likelihood Ratio statistic (*p-value* = 0.025). Therefore, we decided not to impose this additional restriction.

channel relates to the transmission of import prices inflation in foreign currency terms into domestic inflation, and the second one relates to the influence of the exchange rate on prices (through prices of imported intermediate and final goods). We are therefore interested in establishing whether there is a long-run tendency for the Colombian prices to follow foreign prices measured in a common currency, and this involves testing the hypothesis of Purchasing Power Parity (PPP).[8]

PPP states in its absolute version that exchange rates should equalise relative price levels in different countries. Relative, as opposed to absolute PPP requires that percentage changes in exchange rates and relative prices be equal over time. PPP gives an indication of the evolution of domestic versus foreign prices. Thus, if over a period of time domestic prices have been rising more than foreign prices, an exchange rate adjustment is called for in order to restore a country's international competitiveness. However, even if commodity arbitrage opportunities will tend to move the exchange towards PPP, such adjustment may not be instantaneous because of the presence of transportation costs and barriers to trade. This view leads to the suggestion that PPP may not hold in the short run but in the long run. Essentially, this is a question of whether deviations from PPP, i.e. the real exchange rate (defined as the nominal exchange rate multiplied by the ratio of the foreign price level to the domestic price level), may be characterised as a stationary process.

Our empirical analysis of long-run PPP starts off with the j-dimensional Vector Error Correction (VEC) model (1) with exogenous $I(1)$ variables as considered by Pesaran et. al. (2000). This generalisation is particularly useful when applied to a small open economy like Colombia, as it seems reasonable to assume that some of the $I(1)$ variables in the cointegrating VAR model, in our case foreign prices, are long-run forcing variables in the sense that in the long run they are not caused by the other variables in the model.

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[8] PPP is often attributed to Cassel (1916). See e.g. Froot and Rogoff (1995) for a literature review.

We use seasonally unadjusted quarterly observations on the consumer price index (p_t), the peso-US dollar nominal exchange rate (er_t), and the US consumer price index (pus_t). The data set covers the period 1980:1 to 2000:3, and the variables are considered in logarithms. Another measure commonly used for the price level when testing PPP is the producer price index; in any case, our findings are similar regardless of whether we use consumer or producer price indices.

Preliminary analysis of the data using the ADF tests suggested that all series are $I(1)$ with a drift when considered in levels. We then proceed to carry out a cointegration analysis on the three $I(1)$ variables er_t, p_t, and pus_t, treating the last variable as exogenous.[9] Table 1b reports the diagnostic tests for the residuals of the unrestricted VEC model (1) with an exogenous $I(1)$ variable, using a lag length of $k = 2$, allowing the intercept term to enter the model unrestrictedly, and including centred seasonal dummy variables to account for seasonal effects. The diagnostic test statistics for the VEC model reveal no misspecification, except for a minor problem with normality in the exchange rate equation, and rejection of the hypothesis of no ARCH effects (of up to fourth order) with respect to the same equation at a one per cent significance level. The ARCH effects reported in our model, however, are not very serious for the cointegration analysis (see the Monte Carlo evidence in Lee and Tse, 1996).

Cointegration results are shown in Table 2b.[10] The λ-*max* and the *trace* statistics are compared with the critical values tabulated by Pesaran et. al. (2000) allowing for the presence of exogenous $I(1)$ variables; these critical values are larger than the corresponding critical values when all the $I(1)$ variables in the model are treated as endogenous. Both the λ-*max* and the *trace*

[9] In their analyses of PPP, Juselius (1992), Johansen and Juselius (1992) and Pesaran et. al. (2000) also include domestic and foreign interest rates in their models, so that it is possible to test whether the interest rate differential is stationary. For our purposes, however, it turned out that the three-month treasure bill rate in the U.S. secondary market is stationary (at least during our sample period) so that this variable cannot possibly be cointegrated with the domestic interest rate.

[10] All the results reported in this section are carried out using Microfit 4.0. See Pesaran and Pesaran (1997).

statistics support $r = 1$ cointegrating vector. We exactly identify the cointegrating vector by normalising with respect to er_t. Having imposed this exactly identifying restriction, we then test the validity of the hypothesis of relative PPP, which implies a tendency for the nominal exchange rate (er_t) and the relative price ratio ($usp_t - p_t$) to be tied together in the long run. In other words, if the real exchange rate (defined as $er_t + usp_t - p_t$) is stationary the evidence is supportive of the view that relative PPP holds in the long run. The Likelihood Ratio test statistic for testing this over-identifying restriction is $\chi^2(2) = 15.750$, which is well above the 0.01 critical value of the chi-square distribution with two degrees of freedom (i.e. 9.210).

We now examine the validity of a less stringent form of the PPP hypothesis that was suggested by Taylor (1988), who considered whether the linear combination given by $er_t + a \cdot usp_t - a \cdot p_t$ is a cointegrating vector. This linear combination implies a long-run equilibrium relationship between the exchange rate and the relative price ratio, but it should also be noticed that PPP is weakened to symmetry rather than proportionality. Some motivation for allowing $a \neq 1$ is provided by Taylor (1988) by considering models of measurement error and/or transportation costs. The Likelihood Ratio test statistic for testing this over-identifying restriction is $\chi^2(1) = 0.905$, which is insignificant (*p-value* = 0.341). Hence, we are unable to reject this modified version of the PPP hypothesis. Imposing the restriction discussed above yields the cointegrating vector:

$$\begin{aligned} er_t = \; & \underset{(0.078)}{1.283} p_t - \underset{(0.078)}{1.283} usp_t, \end{aligned}$$

which is denoted PPP_t (coefficient standard errors in parentheses); see Figure 2c.

2.3. The labour sector

In this section we estimate a model of wage determination for the Colombian economy adapted from Layard et. al. (1991). The variables used in the model are average wages in current prices

(w_t), prices as measured by the consumer price index (p_t), labour productivity ($prod_t$) calculated as the ratio of total constant price GDP to employment, and the rate of unemployment (U_t).[11] This set of variables allows us to investigate the extent to which real wages are determined by "inside" factors (i.e. labour productivity) relative to "outside" factors (i.e. the unemployment rate). That real wages should depend upon labour productivity is a condition that can be derived from the classic theory of the firm, which postulates a positive relationship between these two variables. As to the unemployment rate, it enters the model to capture the idea that real wages may be affected by the conditions prevailing in the labour market. In particular, real wages tend to increase in labour markets with low unemployment, because employers find it hard to attract new workers, and the bargaining power of unions and workers is strong. On the other hand, real wages tend to decrease in labour markets with high unemployment, because unions and workers find themselves in a weak position, and firms can easily attract new workers. The relationship between real wages and the rate of unemployment would then be expected to be negative.

The data are seasonally unadjusted observations covering the period 1980:1 - 2000:3, and are considered in logarithms, with the exception of the unemployment rate which is considered in percentage terms. ADF test statistics were calculated as an indicator of the order of integration of the variables under consideration, with the hypotheses that w_t, $prod_t$, U_t and p_t are each integrated of order one being accepted.

The next step in our analysis is an investigation of the cointegration properties of the labour sector variables. Here we carry out a cointegration analysis on the set of variables $Y_t = \left[w_t, p_t, prod_t, U_t\right]^{/}$, all of which are treated as endogenous $I(1)$ variables. The

[11] The series of wages, employment and unemployment are for the four main metropolitan areas of the country, that is; Bogota, Cali, Medellin and Barranquilla. The choice of these series is dictated by the availability of data; more comprehensive series are only available for more recent periods of time.

specification of the model contains an unrestricted constant, and centred seasonal dummy variables to account for seasonal effects, and for estimation two lags have been chosen.

Table 1c shows misspecification tests for the estimated VEC(2) model. The diagnostic statistics reveal no misspecification except some ARCH effects in the unemployment equation (at a five per cent significance level), and rejection of normality with respect to the wage, inflation and productivity equations (also at a five per cent level). The normality failure, however, is not so serious for the cointegration tests reported below (see Cheung and Lai, 1993, who find that the Johansen tests perform reasonably well in the presence of excess kurtosis, and Johansen, 1995, p. 29, who points out that although the cointegration analysis is based on Gaussian likelihood, the asymptotic properties only depend on the assumption that the errors are *i.i.d.*).

The λ-*max* and the *trace* test statistics reported in Table 2c indicate that a cointegrating rank $r = 0$ is rejected, whereas $r = 1$ cannot be rejected (both statistics with degrees of freedom adjustment). Hence we proceed under the assumption that there is one cointegration relation, which may include essential parts of a wage formation equation. The restricted cointegrating relation is estimated as: [12]

$$w_t = p_t - \underset{(0.990)}{2.999} U_t,$$

where the standard error is given in parenthesis.

Therefore, our findings imply that in the long run real wages are negatively related to the rate of unemployment, and are not determined by labour productivity. This last result does not give support to the new classical argument that postulates a positive relationship between wages and labour productivity. At first sight this finding might seem peculiar, but it might be explained

[12] For exact identification we normalised with respect to *w*. Having imposed this exactly identifying restriction, we then tested the validity of a unit price elasticity and that labour productivity does not enter the cointegrating relation , producing a $\chi^2(2) = 3.564$ (*p-value* = 0.168).

by the fact that in Colombia wage increases (especially those of public servants) have responded more to institutional factors than to changes in productivity. Historically, wage growth has been strongly influenced by a government fixed minimum wage which usually responded to past inflation. However, during the nineties government policy attempted to link minimum wage increases to expected inflation plus labour productivity, but a Constitutional Court ruling declared that past inflation also had to be taken into account.[13]

In Figure 2d the restricted cointegration vector (denoted Lbm_t) is plotted, after partialling out the full-sample short-run dynamics. From the figure it can be seen that the deviations of nominal wages from the implied long-run relationship have fluctuated within a range of approximately -30 to 40 per cent. It is interesting to notice that since the mid 1990's nominal wages have been consistently above the steady-state level, and this result might help explain the rapid increase in unemployment during the last years (see Figure 1).

3. Modelling the rate of inflation in Colombia

In this section we model inflation as responding to disequilibria in the money market, the goods market, the labour market, and the external sector, as defined by the cointegrating vectors estimated in the previous section. Our empirical modelling exercise attempts to measure the potential impact on inflation of the creation of a Central Bank which is independent from other parts of government.

Following a "general to specific" approach to model building, we formulate a dynamic inflation model in which the inflation rate (Δp_t) is initially regressed on a constant, and one and two lagged values of excess money (Mm_t), excess demand (Gm_t), deviations from PPP (PPP_t), wage deviations from the implied long-run relation (Lbm_t), money growth ($\Delta m1_t$),

[13] Notice also that *prod* is an imprecise measure of productivity, as it is calculated as the ratio of total GDP to employment in the main four metropolitan areas.

variations in the nominal exchange rate (Δer_t), and wage inflation (Δw_t) (market disequilibria are mean-adjusted). The model also includes four lags of the inflation rate, centred seasonal dummies, and an impulse dummy variable (*d*862) that takes the value of one in the second quarter of 1986 (and zero elsewhere) to account for an atypically low inflation rate in that date (mainly the result of low agricultural prices).

As expected, the resulting unrestricted model is initially overparameterised, so a more parsimonious representation could be obtained by excluding regressors with small *t*-ratios $\left(|t| < 1.2\right)$. The finally selected inflation model is reported in Table 4*a* (the overall *F*-statistic for testing that reduction is equal to 0.615; p-value = 0.841). In terms of diagnostic test statistics the estimated model performs satisfactorily (all diagnostic tests are insignificant), and the coefficients are estimated with the theoretically correct sign. In particular, excess money, excess demand and deviations from PPP have an effect on inflation of approximately 14%, 11%, and 2%, respectively. On the other hand, the estimated coefficient on the wage disequilibrium variable is positive (as expected) but not significant. This suggests that inflation does not respond to nominal wage deviations from the steady-state wage relation, at least during the sample period under consideration.[14] The estimate of the autonomous inflationary component is approximately 4.9% per quarter, after all other effects have been accounted for. In contrast to the results in Surrey (1989), Juselius (1992) and Hendry (2000, 2001), our findings indicate that in the case of Colombia internal influences (i.e. Mm_{t-1} and Gm_{t-1}) are more important in explaining inflation than external influences (i.e. PPP_{t-1}). Another feature of the model is the magnitude of the autonomous inflationary component in the Colombian economy, which is more than four times the estimate obtained by Juselius (1992) for Denmark; Hendry (2000, 2001) does not find evidence of autonomous inflation

[14] Considerable care should be exercised when this finding is used for policy analysis. Although wages above the steady-state level do not appear to be inflationary, they might lead to higher unemployment.

when modelling UK inflation.

Model *b* in Table 4 attempts to measure the impact on inflation of the creation, through the 1991 Constitution, of an independent Central Bank. In order to do this, we create a step dummy variable (i.e. dBR_t) which takes the value of one from 1992:1 to 2000:3 and zero otherwise.[15] This dummy variable was then included as an additional regressor in the inflation equation, along with its interaction with our estimated measures of market disequilibria, that is $Mm_t \times dBR_t$, $Gm_t \times dBR_t$, and $PPP_t \times dBR_t$. The estimated coefficients on $PPP_t \times dBR_t$ and dBR_t were insignificant, implying that Central Bank independence did not affect the response of prices to deviations from PPP nor the autonomous level of inflation. The estimated coefficients on $Mm_t \times dBR_t$ and $Gm_t \times dBR_t$ turn out to be negative and statistically different from zero, which indicates that the disequilibria in the goods and monetary markets had a larger effect on inflation before Central Bank independence (that is, 0.416 *vs*. 0.081 for the goods market, and 0.295 *vs*. 0.124 for the monetary market). These larger effects prior to Central Bank independence suggest that the monetary policy simply accommodated the disequilibria in the goods and in the money markets; such accommodation is now effectively forbidden by the Constitution, as referred to in the Introduction.

4. Concluding Remarks

This paper has estimated an inflation model for Colombia in terms of disequilibria in the monetary sector, the foreign sector, and the labour sector. The model also allows us to study the potential effect on inflation of the creation of an independent Central Bank, granted through the constitutional mandate of 1991.

Within the context of vector autoregressive models for these three sectors, evidence

[15] Notice that *dBR* does not start in 1991 in order to allow for the lag between the proclamation of the Constitution, and the moment in which the Board of Directors of the Central Bank started to operate.

has been found for the monetary sector of the existence of two cointegration vectors, which can be interpreted as measures of excess money and of excess demand. The cointegration analysis of the foreign sector provided support for a weak version of PPP between Colombia and the US. And from the analysis of the labour sector we determined a long-run equilibrium relationship among wages, prices and unemployment. Relating the estimated measures of market disequilibria to the inflation rate, coefficients are estimated with the theoretically correct sign: excess money, excess demand and deviations from PPP have a positive effect on inflation, while wages above the steady-state do not appear to have an effect. Our findings indicate that disequilibria in the money and goods markets are a far more powerful influence on inflation than are external factors. Another important result was the rather high autonomous part of the Colombian rate of inflation. The study of the Colombian case thus offers an interesting contrast with previous results obtained for the US, the UK, and Denmark, where external factors were found to be the main driving force behind inflation, and the autonomous inflation component was much smaller (or even insignificant).

The Constitutional Reform of 1991, which gave greater political independence to the Central Bank, did change some of the parameters of the model, as the disequilibria in goods and monetary markets were found to have a larger effect on inflation before Central Bank independence was granted. This larger effects suggest that monetary policy accommodated the disequilibria in the goods and in the money markets; such accommodation is now effectively forbidden by the Constitution.

Table 1. Diagnostic statistics

(a). Monetary sector model

Statistic	$m1$	y	p	R
F *ar*	1.497 [0.216]	1.630 [0.180]	0.781 [0.543]	0.362 [0.835]
F *arch*	0.781 [0.543]	0.164 [0.956]	1.340 [0.268]	0.109 [0.979]
χ^2 *nd*	1.255 [0.534]	0.172 [0.918]	6.689 [0.035]	4.527 [0.104]

(b). Foreign sector model

Statistic	p	er
F *ar*	0.333 [0.855]	0.479 [0.751]
F *arch*	1.804 [0.139]	4.785 [0.002]
χ^2 *nd*	4.225 [0.121]	8.313 [0.016]

(c). Labour sector model

Statistic	w	p	$prod$	U
F *ar*	0.958 [0.437]	0.273 [0.894]	1.064 [0.381]	0.315 [0.867]
F *arch*	0.883 [0.480]	1.964 [0.111]	1.634 [0.177]	3.612 [0.011]
χ^2 *nd*	7.940 [0.019]	7.389 [0.025]	8.969 [0.011]	0.289 [0.866]

Notes:
F *ar* is the Lagrange Multiplier F-test for residual serial correlation of up to fourth order.
F *arch* is the fourth order Autoregressive Conditional Heteroscedasticity F-test.
χ^2 *nd* is a Chi-square test for normality.
Numbers in square brackets are the probability values of the test statistics.

Table 2. Eigenvalues, test statistics, and critical values

(a). Monetary sector model

λ_i	H_0	H_1	λ-*max* statistic	H_0	H_1	λ-*trace* statistic
0.448	$r = 0$	$r = 1$	37.38^{***}	$r = 0$	$r \geq 1$	79.65^{***}
0.337	$r \leq 1$	$r = 2$	25.87^{**}	$r \leq 1$	$r \geq 2$	42.27^{*}
0.189	$r \leq 2$	$r = 3$	13.21	$r \leq 2$	$r \geq 3$	16.39
0.049	$r \leq 3$	$r = 4$	3.19	$r \leq 3$	$r \geq 4$	3.19

(b). Foreign sector model

λ_i	H_0	H_1	λ-*max* statistic	H_0	H_1	λ-*trace* statistic
0.229	$r = 0$	$r = 1$	21.07^{**}	$r = 0$	$r \geq 1$	24.39^{**}
0.040	$r \leq 1$	$r = 2$	3.32	$r \leq 1$	$r \geq 2$	3.32

(c). Labour sector model

λ_i	H_0	H_1	λ-*max* statistic	H_0	H_1	λ-*trace* statistic
0.308	$r = 0$	$r = 1$	26.89^{*}	$r = 0$	$r \geq 1$	50.78^{**}
0.188	$r \leq 1$	$r = 2$	15.17	$r \leq 1$	$r \geq 2$	23.89
0.077	$r \leq 2$	$r = 3$	5.82	$r \leq 2$	$r \geq 3$	8.72
0.039	$r \leq 3$	$r = 4$	2.90	$r \leq 3$	$r \geq 4$	2.90

Notes:
r denotes the number of cointegration vectors.
*, ** and *** denote statistical significance at the 10, 5 and 1 per cent levels, respectively.

Table 3. Inflation models (OLS estimates)

a. Linear model			*b*. Linear model with Central Bank independence		
Variables	Coeff.	(s.e.)	Variables	Coeff.	(s.e.)
Constant	0.035	(0.005)	*Constant*	0.037	(0.007)
Mm_{t-1}	0.135	(0.037)	dBR_t	0.002	(0.005)
Gm_{t-1}	0.110	(0.059)	Mm_{t-1}	0.295	(0.098)
PPP_{t-1}	0.021	(0.005)	$Mm \times dBR_{t-1}$	-0.171	(0.097)
Δp_{t-4}	0.283	(0.094)	Gm_{t-1}	0.416	(0.158)
			$Gm \times dBR_{t-1}$	-0.335	(0.166)
			PPP_{t-1}	0.019	(0.017)
			$PPP \times dBR_{t-1}$	0.005	(0.020)
			Δp_{t-4}	0.215	(0.104)
R^2	0.804			0.818	
F ar	1.344	[0.263]		1.248	[0.300]
F arch	2.231	[0.076]		0.736	[0.572]
χ^2 *nd*	0.652	[0.722]		1.594	[0.451]
F het	1.312	[0.238]		0.805	[0.690]
F Reset	1.748	[0.191]		1.261	[0.266]

Notes:
F *ar* is the Lagrange Multiplier F-test for residual serial correlation of up to fourth order. F *arch* is the fourth order Autoregressive Conditional Heteroscedasticity F-test. χ^2 *nd* is a Chi-square test for normality. *F het* is an *F* test for heteroscedasticity, and F *Reset* is Ramsey's RESET test statistic.
The models include three centred seasonal dummy variables, and an impulse dummy *d*862 as additional regressors.

Figure 1. Graphs of the series

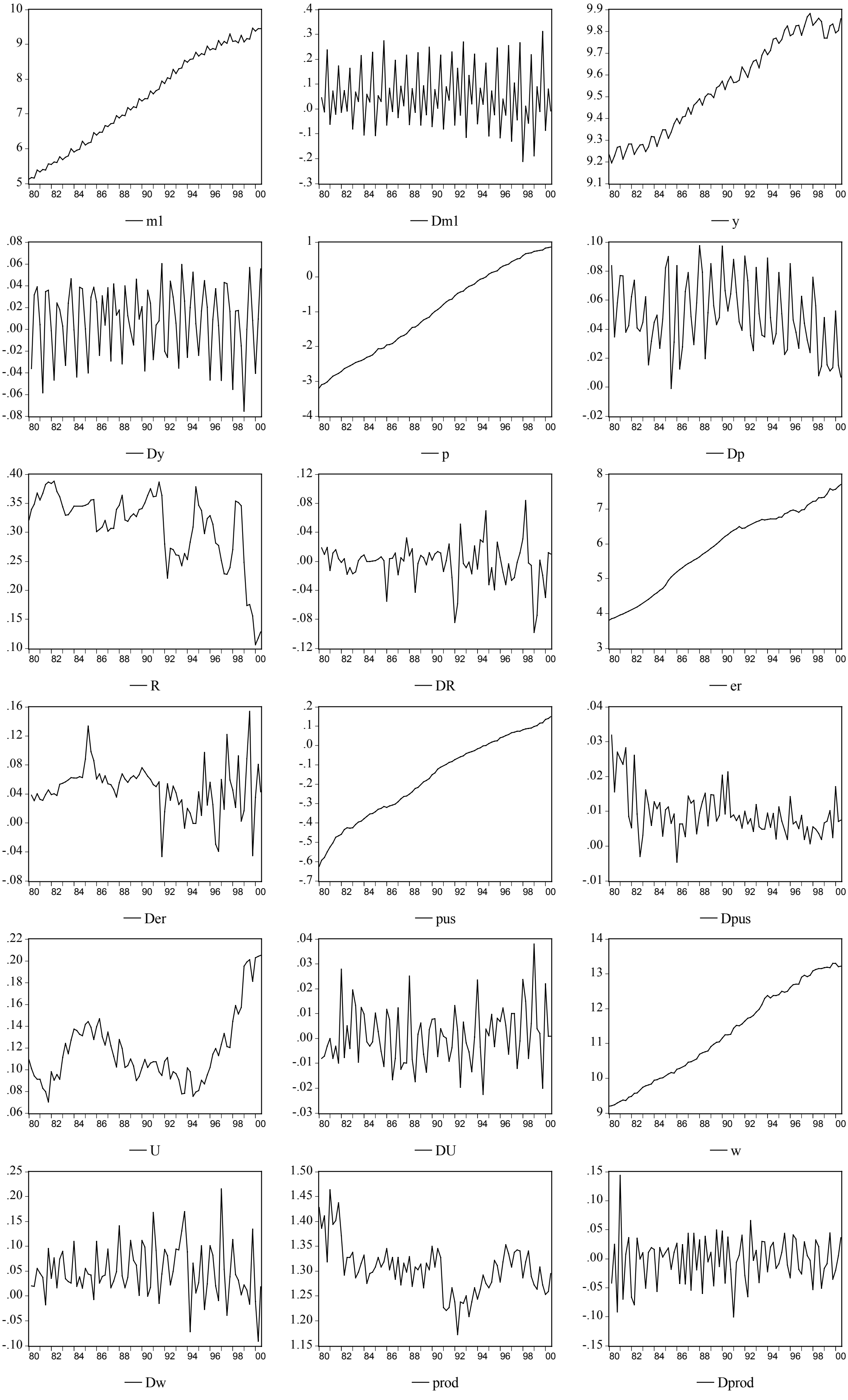
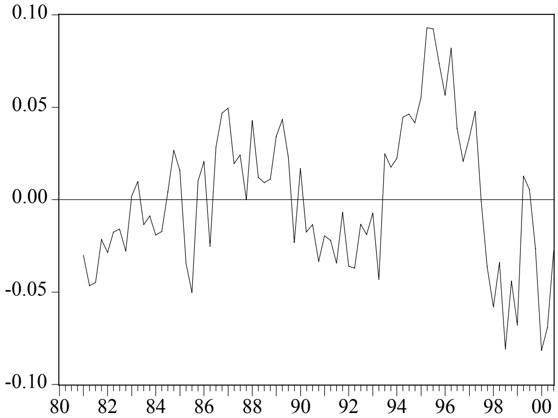
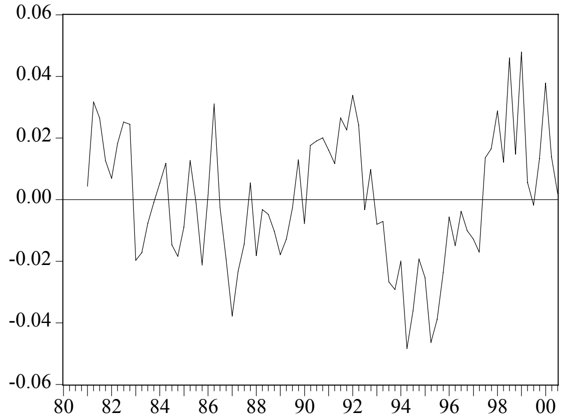


Figure 2. Restricted cointegrating vectors

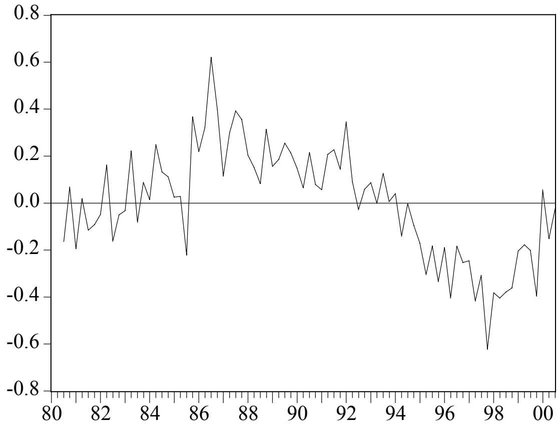
(a). Excess money (*Mm*)



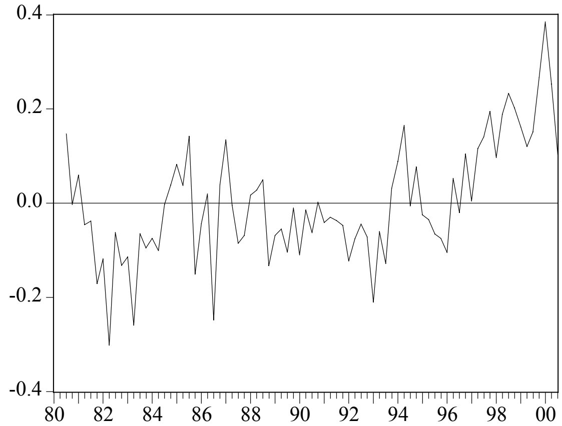
(b). Excess demand (*Gm*)



(c). Deviations from purchasing power parity (*PPP*)



(d). Deviations of nominal wages from the long-run relationship (*Lbm*)



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Data Appendix

*M*1 is the M1 definition of money, that is currency plus demand deposits. The source is Banco de la República.

Y is the series of Gross Domestic Product (in prices of 1994). From 1980 to 1993 the source is Departamento Nacional de Planeación (DNP), and from 1994 onwards the source is Departamento Administrativo Nacional de Estadística (DANE).

R is the yield of 90-day certificates offered by banks and financial corporations. The source is Banco de la República.

P is the consumer price index. The source is Banco de la República.

ER is the peso-dollar nominal exchange rate (in pesos per dollar). The source is Banco de la República.

PUS is the consumer price index in the United States as taken from the Federal Reserve Bank of St. Louis web site at *www.stls.frb.org*.

U is the rate of unemployment in the main four metropolitan areas of the country. The source is Departamento Administrativo Nacional de Estadística (DANE).

W is the series of average nominal wage in the main four metropolitan areas of the country. The source is Departamento Administrativo Nacional de Estadística (DANE)

PROD is labour average productivity, calculated as the GDP series divided by total employment in the main four metropolitan areas of the country. The source of the total employment series is Departamento Administrativo Nacional de Estadística (DANE).

All variables are seasonally unadjusted and are available from the authors upon request.