ECONOMÍA

BORRADORES
DE
INVESTIGACIÓN

No. 19 Enero 2002

ISSN 0124-4396

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ON THE DETERMINANTS OF THE INFLATION RATE IN COLOMBIA: A DISEQUILIBRIUM MARKET APPROACH ♦

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ABSTRACT

We investigate the inflation rate in Colombia in terms of excess money, excess demand, deviations from PPP, and wage inflation. 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. We also find evidence of non-linear price behaviour in response to excess demand and deviations from PPP; prices appear to respond symmetrically to excess money. The constitutional reform of 1991, which gave greater political independence to the Central Bank, also did influence inflation, both changing some parameters of the model and the adjustment patterns.

RESUMEN

Se investiga la tasa de inflación en términos del exceso de dinero, el exceso de demanda, de las desviaciones con respecto a la PPA y de la inflación salarial. A diferencia de los resultados encontrados para un grupo de países industrializados, encontramos que los factores domésticos tienen una influencia mucho mayor sobre la inflación que los factores externos. También encontramos evidencia de comportamiento no lineal de los precios en respuesta a los excesos de demanda y a las desviaciones con respecto a la PPA; los precios responden simétricamente al exceso de dinero. La reforma constitucional de 1991 incrementó la independencia del Banco Central, también influenció la inflación, pero por medio de cambios en los en algunos parámetros y en los patrones de ajuste.

JEL classifications: C32, C51, E31, O54

Keywords: Inflation, cointegration, non-linearities, excess money, excess demand, imported inflation, wage inflation, Colombia

♦ We would like to thank Ana María Iregui, Marta Misas and Lina Sarmiento for providing necessary data. We are also indebted to Luis Eduardo Fajardo and Juan Carlos Guataquí for useful discussions on several of the ideas developed in the paper. Financial support from the Banco de la República (Central Bank of Colombia) is gratefully acknowledge. The views expressed in the paper are the responsibility of the authors, and should not be interpreted as reflecting those of the Board of Directors of the Banco de la República, or other members of its staff. Any remaining errors are ours.

1. Introduction

One of the main subjects of concern of 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 to 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 supply of money 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 inflation of import prices 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). Lastly, a third view lays emphasis on internal theories, which may be further subdivided into labour market theories and excess demand theories. The first of these highlights the role of the wage, being the result of labour demand and supply interactions, as a component of producers costs, while the second refers to excess demand pressure effects.

The purpose of this paper is to investigate the determination of the rate of inflation in Colombia in terms of the macroeconomic 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 for example Herrera (1985), Julio and Cobo (2000), López and Misas (1998), Misas and Oliveros (1994), and Misas et. al. (1999). To our knowledge, however, existing literature has not investigated inflation dynamics accounting for all three explanations, as it has mainly focussed on only one of the possible origins of inflation.[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 a model for the Colombian rate of inflation. 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.

The modelling approach adopted in this paper follows the analyses of inflation in the US and in the UK by Surrey (1989), and in Denmark by Juselius (1992). Both Surrey and Juselius investigate the origins of inflation in these countries by distinguishing between the same three kinds of explanation: pure monetarist theories, external theories, and internal

[1] Misas and Oliveros (1994) and Julio and Cobo (2000) study the dynamics of wages and prices; Herrera (1985) examines the effect of the devaluation of the nominal exchange rate on prices; López and Misas (1998) study inflation dynamics within the context of the Phillips curve; and Misas et. al. (1999) focus on monetary considerations. Gómez and Julio (2001) present an analysis of transmission mechanisms of monetary policy, although their empirical model is not based on a cointegrated VAR model.

theories. However, an important difference in their modelling approaches is that Surrey considers the variables in first differences, possibly to avoid the complications that arise from the nonstationarity of the series, thereby losing information on the long-run dynamics of the data. Juselius, on the other hand, uses multivariate cointegration techniques that allow not only the modelling of the short- and long-run dynamics of the variables simultaneously, but also their interrelations.

The findings of Surrey (1989) and Juselius (1992) indicate that within the industrial economies considered the external influence on domestic prices is a far more powerful influence compared to the domestic influence. The study of the Colombian case using an approach similar to Surrey (1989) and Juselius (1992) is therefore interesting, because it enables us to determine the importance of internal and external influences on domestic prices within the context of a developing economy.

Our paper differs from the earlier works of Surrey (1989) and Juselius (1992) in at least two respects. First, we examine the possibility of non-linear adjustment in the dynamics of the inflation rate. In particular, we discuss two types of non-linear behaviour. The first one is concerned with asymmetric effects of positive and negative deviations from equilibrium, and the second one distinguishes between small and large (both positive and negative) deviations from equilibrium. A number of reasons could potentially explain non-linear behaviour in the dynamics of the inflation rate, including government intervention in some markets, e.g. in the form of price controls or other regulations; in addition, the speed at which economic forces act to restore short-run discrepancies in the equilibrium relationships might differ depending upon the market under consideration. Second, the analysis of the Colombian experience allows us to assess the effects of the Constitutional reform of 1991, which provided for a greater independence of the Central Bank from the government's executive branch. The institutional changes introduced by the new Constitution have now been in place long enough to allow some hope of identifying the effects they may have had.

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 inflation model allows for linear and non-linear adjustment to the disequilibrium errors. Section 4 offers some concluding remarks.

2. The analysis of long-run equilibrium relationships

Our empirical analysis of the determinants of the rate of inflation in Colombia starts off with the investigation of the long-run structure of the data. 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 one of these sectors, we apply the Johansen (1988, 1995) full information maximum likelihood procedure, as it allows 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.

We now present the results of the cointegration analysis for the three economic 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 be then 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 *M*1, 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 the opportunity costs for holding money.[2] The money, output and price series are considered in logarithms and denoted $m1_t$, 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. The data are seasonally

[2] It was also tried to estimate a model using the *M*3 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 *M*3. As to 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.

unadjusted quarterly observations from 1980:1 to 2000:3 (see the data appendix for definitions and sources of the variables 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 (see Table 1a).

An additional feature of the money, output and price series is that they display a seasonal component, so that it seems sensible to study whether the seasonal patterns are regular (i.e. deterministic) or changing (i.e. stochastic). The existence of unit roots at seasonal frequencies can be interpreted as indication of a changing seasonal pattern. Hylleberg, Engle, Granger and Yoo (Hylleberg et. al., 1990) [HEGY] propose a test to test for the existence of both a unit root as well as seasonal roots. The results of applying the HEGY test are given in Table 2a. From the results in this table it is clear that the money and price series do not have unit roots at frequencies other than the zero (or long-run) frequency, and that the output series presents a unit root at the zero frequency as well as at the annual frequency. Hence, for the purposes of our cointegrating analysis it seems appropriate to model the unadjusted data using seasonal dummies.

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.[3] 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 3a).[4] 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 4a, 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. The unrestricted cointegrating vectors are reported in the left hand side panel of Table 5a. The first vector seems to be a money demand equation (possibly with a unit coefficient on prices) although notice that the interest rate semi elasticity does not have the expected negative sign. At first sight the second vector does not have a clear economic interpretation but later on, after a number of restrictions motivated by economic theory are tested on the cointegrating vectors, it is shown that it can be interpreted as an aggregate demand equation.[5]

Next, we identify the two cointegrating vectors. Exact identification of β (in $\Pi = \alpha\beta'$)

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[3] See e.g. Arrau et. al. (1995) on the use of a deterministic trend to capture the role of financial innovation.

[4] Estimations are done in PcGive and PcFiml 9.0 (see Hendry and Doornik, 1997).

[5]To check whether the long-run relationships found by the Johansen procedure are stable we calculated recursive eigenvalues of the VEC model. Figure A1 in Appendix A plots the time paths of the eigenvalues with their asymptotic 95% error bounds for each sub-sample. The plots of the recursive eigenvalues are reasonably stable over time, and also show that there are two cointegrating vectors.

requires at least r restrictions (including the normalising restrictions) on each of the r cointegrating relationships (see Pesaran and Shin, 1999). Here, we distinguish between *non-testable* and *testable* restrictions. *Non-testable* restrictions are the exactly identifying restrictions; these are needed to achieve exact identification of the cointegration space. On the other hand, *testable* restrictions refer to additional over-identifying restrictions the validity of which can be tested using standard Likelihood Ratio tests. Under the assumption of $r = 2$ cointegrating vectors, we need to impose two (*non-testable*) restrictions on each of the two vectors to exactly identify them. To do so, we denote the two cointegrating relationships associated with $Y_t = [m1_t, y_t, p_t, R_t, trend]'$, by:

$\beta_1 = [\beta_{11}, \beta_{12}, \beta_{13}, \beta_{14}, \beta_{15}]'$,

and

$\beta_2 = [\beta_{21}, \beta_{22}, \beta_{23}, \beta_{24}, \beta_{25}]'$,

respectively. Notice that there are five elements in each of the two vectors. These are the coefficients of the four endogenous variables, $m1_t$, y_t, p_t, R_t, as well as the coefficient associated to the trend term. To identify the two cointegrating vectors, we begin by imposing long-run price homogeneity on the first vector and long-run exclusion of the trend term on the second one. Formally, we thus have the following exactly identifying restrictions:

$\beta_{11} = 1$, $\beta_{13} = -1$ (on the first vector),

and

$\beta_{22} = 1$, $\beta_{25} = 0$, (on the second vector).

Having imposed exactly identifying restrictions on the two vectors, we then test the validity of further over-identifying restrictions:

$\beta_{12} = -1$,

and

$\beta_{24} = 0$.

The two *testable* over-identifying restrictions refer to a proportional effect between y and $m1$ on the first vector (i.e. $\beta_{12} = -1$), and long-run exclusion of R_t on the second vector (i.e. $\beta_{24} = 0$). The reason for testing these restrictions is that, after imposing the exactly identifying restrictions, the estimated coefficient associated to y_t in the first vector was close to one, whereas the exclusion of R_t on the second vector allows us to interpret the cointegrating relationship as an aggregate demand equation.[6] The Likelihood Ratio test statistic for testing these two over-identifying restrictions is distributed as a $\chi^2(2)$ under the null hypothesis, giving a value of 1.005 which is insignificant (*p-value* = 0.605). Imposing the restrictions discussed above, yields the restricted cointegrating vectors reported in the right hand side panel of Table 5a.

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]

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[6] For a similar identification approach see e.g. Johansen and Juselius (1994).

[7] Notice that the estimated coefficients on $m1$ 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

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).

Ordinary Least Squares (OLS) estimates of the error correction models are reported in Table 6a. We report only the estimated coefficients associated with the error correction terms, as we are mainly interested in the characteristics of the short-run adjustment (the other regressors appear in the notes of the table). The results show significant feedback from the disequilibrium error in the money market in the $\Delta m1_t$, Δp_t, and ΔR_t equations, and from the disequilibrium error in the goods market in all four equations. Of primary interest to us is the finding that both the disequilibria in the money and in the goods markets turn out to be statistically significant for the inflation rate (here measured by the changes in the consumer price index). Later on we will investigate whether these variables are still significant when a more extensive information set is used.

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 channel relates to the transmission of inflation of import prices 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 adjustment of the exchange rate 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

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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.

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

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.

We use seasonally unadjusted quarterly observations on the consumer price index (p_t), the nominal exchange rate (er_t), and the U.S. consumer price index (pus_t). The data set covers the period 1980:1 to 2000:3, and the variables are considered in logarithms (see Figure 1). 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 Augmented Dickey-Fuller (ADF) tests suggested that the nominal exchange rate and the price indices are $I(1)$ with a drift when considered in levels (see Table 1b).[9]

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.[10] Table 3b 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 4b, which reports the λ_i eigenvalues, the λ-*max* and the *trace* statistics.[11] 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* statistics support $r = 1$ cointegrating vector.[12]

We now examine the validity of the PPP hypothesis using the long-run structural modelling techniques proposed by Pesaran and Shin (1999). To do so, let us denote Z_t the vector of endogenous and exogenous $I(1)$ variables $Z_t = [er_t, p_t, usp_t]'$, and $\beta_1 = [\beta_{11}, \beta_{12}, \beta_{13}]'$ the corresponding cointegrating relationship associated with it. In line with

[9] When we applied the HEGY to *pus* the results also supported the presence of a unit root at the zero (or long-run) frequency but not at seasonal frequencies (see Table 2b).

[10] 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.

[11] The results of the cointegration analysis reported in this section are carried out using Microfit 4.0 (see Pesaran and Pesaran, 1997).

[12] Figure A2 in the Appendix plots the time paths of the eigenvalues calculated from a recursive estimation of the model. The graph of the first recursive eigenvalue show that there is one cointegrating vector. The second eigenvalue is nearly zero.

our discussion in the previous section, we exactly identify the cointegrating vector by imposing $\beta_{11} = 1$ which refers to normalisation with respect to the nominal exchange rate. The resulting cointegrating vector is reported in the left hand side panel of Table 5b. Notice that the coefficient on the domestic price level is negative (i.e. -2.064), while the coefficient on the foreign price level is positive (i.e. 5.748). This finding already supports a weaker form of PPP with $\beta_{12} < 0$ and $\beta_{13} > 0$, but not necessarily unity in magnitude (see e.g. MacDonald 1993, and Chen 1995).

Other variations of the PPP hypothesis can be tested using the above exactly identified model. Consider first the hypothesis of relative PPP which implies a tendency for the nominal exchange rate (er_t) and the relative price ratio ($p_t - usp_t$) to be tied together in the long run. In other words, if the real exchange rate (defined as $er_t - p_t + usp_t$) is stationary the evidence is supportive of the view that relative PPP holds in the long run. The PPP hypothesis is thus whether the vector $\beta_1 = [1, -1, 1]^{/}$ is a cointegrating vector. The Likelihood Ratio test statistic for testing this over-identifying restriction is distributed as a $\chi^2(2)$ under the null hypothesis, giving a value of 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 $\beta_1 = [1, -a, a]^{/}$ 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 distributed as a $\chi^2(1)$ under the null hypothesis, giving a value of 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 restricted cointegrating vector reported in the right hand side panel of Table 5b, which is denoted PPP_t (see Figure 2c).

OLS estimates of the error correction model are reported in Table 6b along with a number of diagnostic test statistics. From the table it can be seen that the Δp_t equation performs best, explaining about two-thirds of the domestic price variation over the sample period. The error correction term associated with this equation allows us to investigate the external transmission effects as a result of deviations from PPP. The estimate of this turns out to have a significant but small positive impact on domestic price changes, suggesting an equilibrating but slow adjustment for Colombian prices in response to changes in the nominal exchange rate and in foreign prices. The error correction term in the Δer_t equation is positive but statistically insignificant.

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) (see also Marcellino and Mizon 2000; and Nymoen 1989). 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).[13] 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); see Layard et. al. (1991). 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, in labour markets with low unemployment real wages tend to increase, because employers find it hard to attract new workers, and the bargaining power of unions and workers is strong. In labour markets with high unemployment, on the other hand, real wages tend to decrease 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 (see Figure 1). As an indicator of the order of integration of the variables under consideration Augmented Dickey-Fuller test statistics were calculated, with the hypotheses that w_t, $prod_t$, U_t and p_t are each integrated of order one being accepted (see Table 1c). The HEGY tests in Table 2c support this finding, and also reject the presence of unit roots at seasonal frequencies.

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 $Y_t = [w_t, p_t, prod_t, U_t]'$, all of which are treated as endogenous $I(1)$ variables. The 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 3c 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 4c, 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.[14] The unrestricted standardised cointegration vector given in the first row of the left hand side panel of Table 5c

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[13] 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.

[14] All eigenvalues seem reasonably stable when recursively estimated over time; see Figure A3 in the Appendix.

suggests that there may be a unit price elasticity. The Likelihood Ratio test statistic for testing this over-identifying restriction is distributed as a $\chi^2(1)$ under the null hypothesis, giving a test value of 3.468 which is not significant (*p-value* = 0.063). Hence the restriction is accepted by the data.

Next, we consider the hypothesis that labour productivity does not enter the cointegrating relation; the reason for testing this additional restriction is that, after imposing a long-run unit price elasticity, the estimated coefficient associated to $prod_t$ was close to zero (i.e. 0.219). The Likelihood Ratio test statistic for testing these two over-identifying restrictions is distributed as a $\chi^2(2)$ under the null hypothesis, giving a test value of 3.564 which is insignificant (*p-value* = 0.168). Imposing both a long-run unit price elasticity and long-run exclusion of the productivity variable, yields the restricted cointegration vector reported in the right hand side panel of Table 5c.

Our findings thus 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 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 had also to be taken into account.[15]

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).

OLS estimates of the error correction models together with a number of diagnostic tests are reported in Table 6c. As we are mainly interested in the characteristics of the short-run adjustment, we report only the coefficients associated with the deviations from the wage relation (other regressors are omitted to save space). The results show that the error correction term has a significant negative impact on current wage changes, suggesting an equilibrating adjustment process for wages in response to changes in domestic prices and unemployment. There is a small negative relationship between inflation and the disequilibrium error indicating that wages above the steady-state level are deflationary rather than inflationary. This result is surprising considering the role of wages as part of production costs, although in the next section we shall investigate the hypothesis of wage inflation on a more extensive information set. Unemployment rises with increases in wages above the steady-state level, and productivity growth does not respond to deviations from the long-run wage relation.

3. Modelling the rate of inflation

In this section we estimate a model of inflation in terms of the deviations from the

[15] 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.

steady-state relations derived in our cointegration analysis of the monetary sector, the foreign sector and the labour sector. The model accounts for the possible role played by inflationary inertia and by seasonal factors (both deterministic and stochastic), and at a later stage of the analysis we also investigate whether the steps taken by the Constitutional Reform of 1991 in the sense of granting greater political independence to the Central Bank, had an effect on the estimated parameters of the model.

Following a "general to specific" approach to model building, the inflation rate (Δp_t) was initially regressed on a constant, excess money (Mm_{t-1}), excess demand (Gm_{t-1}), deviations from PPP (PPP_{t-1}), and wage deviations from the implied long-run relation (Lbm_{t-1}) (these variables are measured as deviations around their mean values to facilitate the analysis that follows). The model also includes four lags of the inflation rate, centred seasonal dummies, and an impulse dummy variable ($d862$) 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 regression is initially overparameterised, so a more parsimonious representation could be obtained by excluding some of the regressors based on Wald tests for zero restrictions.

Model 1 in Table 7 is the resulting linear inflation model after eliminating insignificant variables. In terms of diagnostic test statistics the estimated model performs satisfactorily, and the coefficients are estimated with the theoretically correct sign. In particular, excess money, excess demand and deviations from PPP have a positive effect on the rate of inflation. 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.[16] Lastly, the estimate of the constant term indicates that the autonomous (or inertial) component of the rate of inflation is approximately 3.5% per quarter, after all other effects have been accounted for. In contrast to the earlier results reported by Surrey (1989) and Juselius (1992) for the US, the UK and Denmark, 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}). Additionally, our estimate of the autonomous inflationary component for the Colombian economy is more than three times Juselius (1992) estimate for Denmark.

In recent years various authors have examined non-linearities in the behaviour of error correction models (see Granger and Lee, 1989; Granger and Teräsvirta, 1993; Escribano and Granger, 1998; Escribano and Pfann, 1998; and Escribano and Aparicio, 1999, among others). For instance, Granger and Lee (1989) partition the error correction term into its positive and negative components, and feed them back into the short-run dynamic equations, whereas Escribano and Granger (1998) and Escribano and Aparicio (1999) use a cubic error correction term.

Models 2 and 3 in Table 7 are estimated based on different types of non-linear adjustment. First, following Granger and Lee (1989), in model 2 we take the deviations of Mm_t, Gm_t, and PPP_t around their mean values, and partition them into their positive and negative components (denoted by x^+ and x^-, where $x = Mm_t$, Gm_t, and PPP_t, respectively).

[16] 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 do result in higher unemployment (see Table 6c).

In order to visualise the adjustment mechanisms observed for the rate of inflation, Figure 3 plots the estimated asymmetric price response against the relevant market disequilibrium (lagged once). It appears that prices increase rapidly when there is a positive excess demand for goods, and decrease slowly when the excess demand is negative. There does not seem to be evidence of differential price adjustment depending upon the type of disequilibrium in the monetary market (notice that the cross-plot in this case is a straight line). Lastly, one can also notice that the evidence of asymmetric price adjustment in response to deviations from PPP is very weak.

However, imposing a unique equilibrium around zero may be too restrictive. In order to relax this assumption, in model 3 we follow Escribano and Granger (1998) and Escribano and Aparicio (1999) who allow for x^2 and x^3, where $x = Mm_t$, Gm_t, and PPP_t, respectively, to enter the inflation equation. This type of non-linear adjustment is more flexible than the Granger and Lee (1989) type of asymmetric adjustment as it allows for the possibility of more than one equilibrium points. Figure 4 shows the estimated non-linear adjustments in prices against Mm_{t-1}, Gm_{t-1}, and PPP_{t-1}. The non-linear adjustment results support the findings of asymmetric adjustment around a unique (at the zero point) equilibrium. Evidence of non-linear price behaviour is observed following disequilibria in the goods market, and also in response to large deviations (both positive and negative) from PPP. Prices appear to respond symmetrically to positive and negative (both small and large) disequilibria in the money market.

The empirical analysis so far has not addressed an interesting policy question that arises in the Colombian context: The fact that in 1991 a major Constitutional reform took place in the country, 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.[17] Thus, it is of particular interest to test whether the Constitutional reforms regarding Central Banking had an effect on the estimated parameters of our inflation model. 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.[18] 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 results presented in model 4 of Table 7 indicate that the disequilibria in the goods and in the 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. The estimated coefficient on $PPP_t \times dBR_t$ was not found to be significant implying that central bank independence did not affect the response of prices to deviations from PPP.

[17] 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 also stated that the main objective of the Bank was to control inflation, and that it had to coordinate its policies with the macroeconomic policies of the government.

[18] 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.

Having found a differentiated effect on inflation of the disequilibria in the money and in the goods markets before and after central bank independence, it is worth considering the possibility of non-linear adjustments in each sub-period. To investigate the possibility of asymmetric adjustment, we partition x into x^{+} and x^{-}, where $x = Mm_t$, $Mm_t \times dBR_t$, Gm_t, and $Gm_t \times dBR_t$. To examine the possibility of non-linear adjustment, we create the squares and cubes of x, where $x = Mm_t$, $Mm_t \times dBR_t$, Gm_t, and $Gm_t \times dBR_t$.

Figures 5a,b and 6a,b plot the estimated asymmetric and non-linear cubic polynomial adjustments, respectively, against Gm_{t-1}. The plots offer strong evidence in favour of non-linear price adjustments, and also suggest that the adjustment patterns changed with the introduction of an independent Central Bank. Figures 5c,d plot the estimated asymmetric adjustments against Mm_{t-1} before and after Central Bank independence, respectively. The figures show that the evidence of asymmetric adjustment is very weak in both sub-periods. Figures 6c,d report the estimated non-linear cubic polynomial adjustments against Mm_{t-1}, from which one can notice evidence of non-linear adjustment before Central Bank independence, at least for large positive values of Mm_{t-1}.

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. Within the context of linear vector autoregressive models for these three sectors, evidence 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 derived with the help of the cointegration analysis to the inflation rate, the 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 also indicate that the disequilibria in the money and in the goods markets are a far more powerful influence on inflation than are external factors. 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.

The linear inflation model was subsequently extended to allow for the possibility of asymmetric and non-linear price adjustment in response to market disequilibria. The non-linear extension is particularly useful as it allows us to examine the asymmetric effects of positive and negative deviations from equilibrium, as well as the differential effects of small and large discrepancies. We find evidence of non-linear price behaviour in the case of the disequilibrium in the goods market (with prices increasing rapidly following a positive excess demand, and decreasing slowly when the excess demand is negative), and also in response to large deviations (both positive and negative) from PPP. We did not find evidence of non-linear adjustment in the case of the disequilibrium in the money market. Lastly, the Constitutional Reform of 1991, which gave greater political independence to the Central Bank, did change the parameters of the model, as the disequilibria in the goods and in the monetary markets were found to have a smaller effect on inflation after Central Bank independence. The adjustment patterns also appear to have changed with the introduction of an independent Central Bank.

Despite the interesting properties of the non-linear models considered in the paper, an important question that needs to be answered is how successful they are (compared to the linear version) for forecasting inflation. It is our intention to perform an evaluation of the forecast performance of the models estimated in this paper in future research.

Table 1.Dickey and Fuller unit root τ-tests

(a). Monetary sector variables

Variable	Model	Lags	τ	F *ar*
*m*1	C, T	4	-1.619	[0.464]
y	C, T	5	-2.505	[0.147]
p	C, T	4	-0.325	[0.374]
R	C	1	-1.955	[0.482]

(b). Foreign sector variables

Variable	Model	Lags	τ	F *ar*
p	C, T	4	-0.325	[0.374]
pus	C, T	3	-1.925	[0.684]
er	C, T	1	-0.918	[0.294]

(c). Labour sector variables

Variable	Model	Lags	τ	F *ar*
p	C, T	4	-0.325	[0.374]
U	C	4	-0.617	[0.392]
prod	C, T	4	-2.554	[0.096]
w	C, T	4	-1.224	[0.959]

Notes:
C, T indicates that the Dickey-Fuller regression contains a constant term and trend. The 5% critical value from MacKinnon (1991) is -3.465.
C indicates that the Dickey-Fuller regression contains a constant term. The 5% critical value from MacKinnon (1991) is -2.897.
The ADF regressions (except those for *R*, *er* and *w*) also include seasonal dummy variables.
F *ar* is the Lagrange multiplier F-test for residual serial correlation of up to fourth order. Numbers in square brackets are the probability values of the test statistics.

Table 2. HEGY test for seasonal integration

(a). Monetary sector variables

Variable	Lags	$t(\pi_1)$	$t(\pi_2)$	$F(\pi_3,\pi_4)$	$F(\pi_2,\pi_3,\pi_4)$	F *ar*
$m1$	1	-1.875	-2.156	10.129 **	8.479 **	[0.393]
y	2	-2.705	-2.561 *	3.282	4.343	[0.247]
p	1	-1.168	-2.899 **	9.817 **	9.278 **	[0.060]

(b). Foreign sector variables

Variable	Lags	$t(\pi_1)$	$t(\pi_2)$	$F(\pi_3,\pi_4)$	$F(\pi_2,\pi_3,\pi_4)$	F *ar*
p	1	-1.168	-2.899 **	9.817 **	9.278 **	[0.060]
pus	1	-1.595	-3.039 **	11.909 **	10.709 **	[0.244]

(c). Labour sector variables

Variable	Lags	$t(\pi_1)$	$t(\pi_2)$	$F(\pi_3,\pi_4)$	$F(\pi_2,\pi_3,\pi_4)$	F *ar*
p	1	-1.168	-2.899 **	9.817 **	9.278 **	[0.060]
U	1	-1.000	-2.909 **	16.121 **	12.842 **	[0.410]
$prod$	1	-2.482	-4.606 **	19.261 **	19.164 **	[0.055]

Notes:
The tests (except those for *U*) are based on an auxiliary regression that includes constant, trend, and seasonal dummy variables. In the case of *U* the auxiliary regression includes constant and seasonal dummy variables
F *ar* is the Lagrange multiplier F-test for residual serial correlation of up to fourth order. Numbers in square brackets are the probability values of the test statistics.
The critical values for the $t(\pi_1)$, $t(\pi_2)$ and $F(\pi_3,\pi_4)$ tests are taken from Franses and Hobijn (1997). The critical values for the $F(\pi_2,\pi_3,\pi_4)$ test are taken from Ghysels et. al. (1994). * and ** denote statistical significance at the 10 and 5 per cent, respectively.

Table 3. 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 4. 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 5. Cointegration analysis: Standardized beta eigenvectors

(a). Monetary sector model

	Unrestricted					Restricted				
	*m*1	*y*	*p*	*R*	*Trend*	*m*1	*y*	*p*	*R*	*Trend*
1	1.000	-2.054	-0.744	-0.569	-0.001	1.000	-1.000	-1.000	0.715	0.005
2	0.921	1.000	-1.368	1.897	0.013	-0.598	1.000	0.482		
3	-4.803	31.089	1.000	-14.19	-0.081					
4	0.368	0.451	-1.647	1.000	0.055					

(b). Foreign sector model

	Unrestricted			Restricted		
	er	*p*	*usp*	*er*	*p*	*usp*
1	1.000	-2.064	5.748	1.000	-1.283	1.283
2	-13.644	1.000	61.104			

(c). Labour sector model

	Unrestricted				Restricted			
	w	*p*	*prod*	*U*	*w*	*p*	*prod*	*U*
1	1.000	-1.049	-0.775	1.665	1.000	-1.000		2.999
2	-0.858	1.000	5.084	2.722				
3	6.648	-7.344	1.000	-12.937				
4	-0.136	0.117	-0.118	1.000				

Table 6. Error correction model (OLS estimates)

(a). Monetary sector model

Variables	$\Delta m1_t$		Δy_t		Δp_t		ΔR_t	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
mm_{t-1}	-0.365	0.158	0.007	0.078	0.207	0.075	-0.295	0.133
gm_{t-1}	-0.929	0.290	-0.489	0.143	0.285	0.138	-0.688	0.244
R^2	0.956		0.843		0.752		0.405	
F ar	0.815	[0.521]	1.728	[0.156]	0.727	[0.577]	1.066	[0.382]
F arch	0.229	[0.921]	0.179	[0.949]	1.197	[0.323]	0.825	[0.515]
χ^2 *nd*	0.446	[0.800]	0.243	[0.885]	4.639	[0.098]	5.696	[0.058]
F het	0.509	[0.966]	0.486	[0.974]	0.717	[0.817]	0.871	[0.648]
F Reset	0.061	[0.806]	2.199	[0.143]	2.645	[0.109]	9.367	[0.003]

(b). Foreign sector model

Variables	Δer_t		Δp_t	
	Coeff.	S.E.	Coeff.	S.E.
ppp_{t-1}	0.025	0.015	0.029	0.007
R^2	0.189		0.687	
F ar	0.443	[0.777]	0.402	[0.806]
F arch	4.013	[0.006]	2.863	[0.030]
χ^2 *nd*	9.832	[0.007]	5.850	[0.054]
F het	2.816	[0.005]	3.335	[0.001]
F Reset	0.357	[0.552]	0.512	[0.477]

(c). Labour sector model

Variables	Δw_t		Δp_t		$\Delta prod_t$		ΔU_t	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
lbm_{t-1}	-0.116	0.040	-0.056	0.013	0.025	0.031	0.014	0.007
R^2	0.349		0.685		0.347		0.533	
F ar	0.789	[0.536]	0.256	[0.905]	2.139	[0.085]	0.567	[0.688]
F arch	1.073	[0.377]	1.922	[0.118]	0.371	[0.828]	3.535	[0.011]
χ^2 *nd*	9.550	[0.008]	7.664	[0.022]	10.417	[0.006]	0.374	[0.830]
F het	0.496	[0.918]	0.874	[0.584]	1.357	[0.208]	1.085	[0.390]
F Reset	0.874	[0.353]	0.070	[0.793]	0.140	[0.709]	0.271	[0.604]

Notes:
The ECM in part A also includes intercept, $\Delta m1_{t-j}$, Δy_{t-j}, Δp_{t-j}, ΔR_{t-j}, $j = 1,2,3$, and seasonal dummies.
The ECM in part B also includes intercept, Δer_{t-1}, Δp_{t-1}, Δusp_{t-1}, and seasonal dummies.
The ECM in part C also includes intercept, Δw_{t-1}, Δp_{t-1}, $\Delta prod_{t-1}$, ΔU_{t-1}, and seasonal dummies.
F het is an *F* test for heteroscedasticity, and F *Reset* is Ramsey's RESET test statistic. The other diagnostic tests are discussed in the notes of Table 3.

Table 7. Inflation models (OLS estimates)

Model 1 Variable	Coeff	Model 2 Variable	Coeff	Model 3 Variable	Coeff	Model 4 Variable	Coeff
Intercept	0.035	*Intercept*	0.033	*Intercept*	0.036	*Intercept*	0.037
	(0.005)		(0.006)		(0.006)		(0.007)
Mm_{t-1}	0.135	Mm^{+}_{t-1}	0.135	Mm_{t-1}	0.155	dBR_{t}	0.002
	(0.037)		(0.060)		(0.052)		(0.005)
Gm_{t-1}	0.110	Mm^{-}_{t-1}	0.136	Mm^{2}_{t-1}	0.027	Mm_{t-1}	0.295
	(0.059)		(0.052)		(0.177)		(0.098)
PPP_{t-1}	0.021	Gm^{+}_{t-1}	0.142	Mm^{3}_{t-1}	-0.195	$Mm{\times}dBR_{t-1}$	-0.171
	(0.005)		(0.090)		(0.899)		(0.097)
Δp_{t-4}	0.283	Gm^{-}_{t-1}	0.066	Gm_{t-1}	0.180	Gm_{t-1}	0.416
	(0.094)		(0.096)		(0.086)		(0.158)
		PPP^{+}_{t-1}	0.023	Gm^{2}_{t-1}	0.605	$Gm{\times}dBR_{t-1}$	-0.335
			(0.012)		(0.658)		(0.166)
		PPP^{-}_{t-1}	0.020	Gm^{3}_{t-1}	-8.038	PPP_{t-1}	0.019
			(0.011)		(5.376)		(0.017)
		Δp_{t-4}	0.287	PPP_{t-1}	0.029	$PPP{\times}dBR_{t-1}$	0.005
			(0.099)		(0.016)		(0.020)
				PPP^{2}_{t-1}	-0.001	Δp_{t-4}	0.215
					(0.029)		(0.104)
				PPP^{3}_{t-1}	-0.053		
					(0.117)		
				Δp_{t-4}	0.227		
					(0.106)		
R^2	0.804		0.806		0.814		0.818
F ar	1.344 [0.263]		1.554 [0.198]		1.716 [0.159]		1.248 [0.300]
F arch	2.231 [0.076]		1.924 [0.119]		2.250 [0.075]		0.736 [0.572]
χ^2 *nd*	0.652 [0.722]		0.663 [0.718]		0.960 [0.619]		1.594 [0.451]
F het	1.312 [0.238]		1.046 [0.432]		0.706 [0.803]		0.805 [0.690]
F Reset	1.748 [0.191]		1.942 [0.168]		1.613 [0.209]		1.261 [0.266]

Notes:
Model 1 is the linear model.
Model 2 is the linear model with asymmetric adjustment.
Model 3 is the linear model with non-linear (cubic) adjustment.
Model 4 is the linear adjustment model with Central Bank independence dummy (i.e. *dBR*) and interaction terms.
The models also include three centred seasonal dummy variables and the impulse dummy variable *d*862. Numbers in parentheses are standard errors and numbers in square brackets are the probability values of the test statistics. The diagnostic tests are discussed in the notes of Tables 3 and 6

Figure 1. Graphs of the series

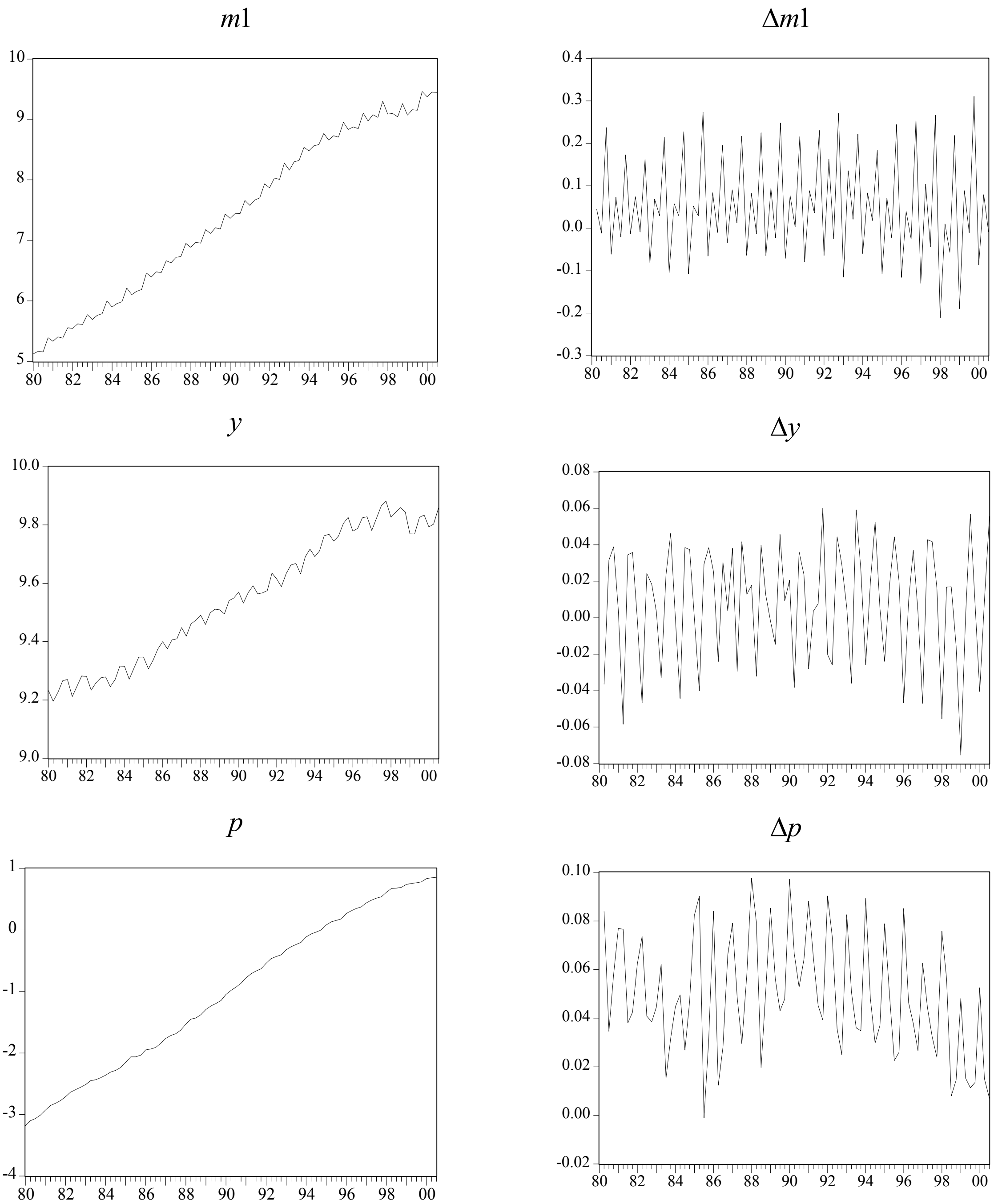


Figure 1 (Continued). Graphs of the series

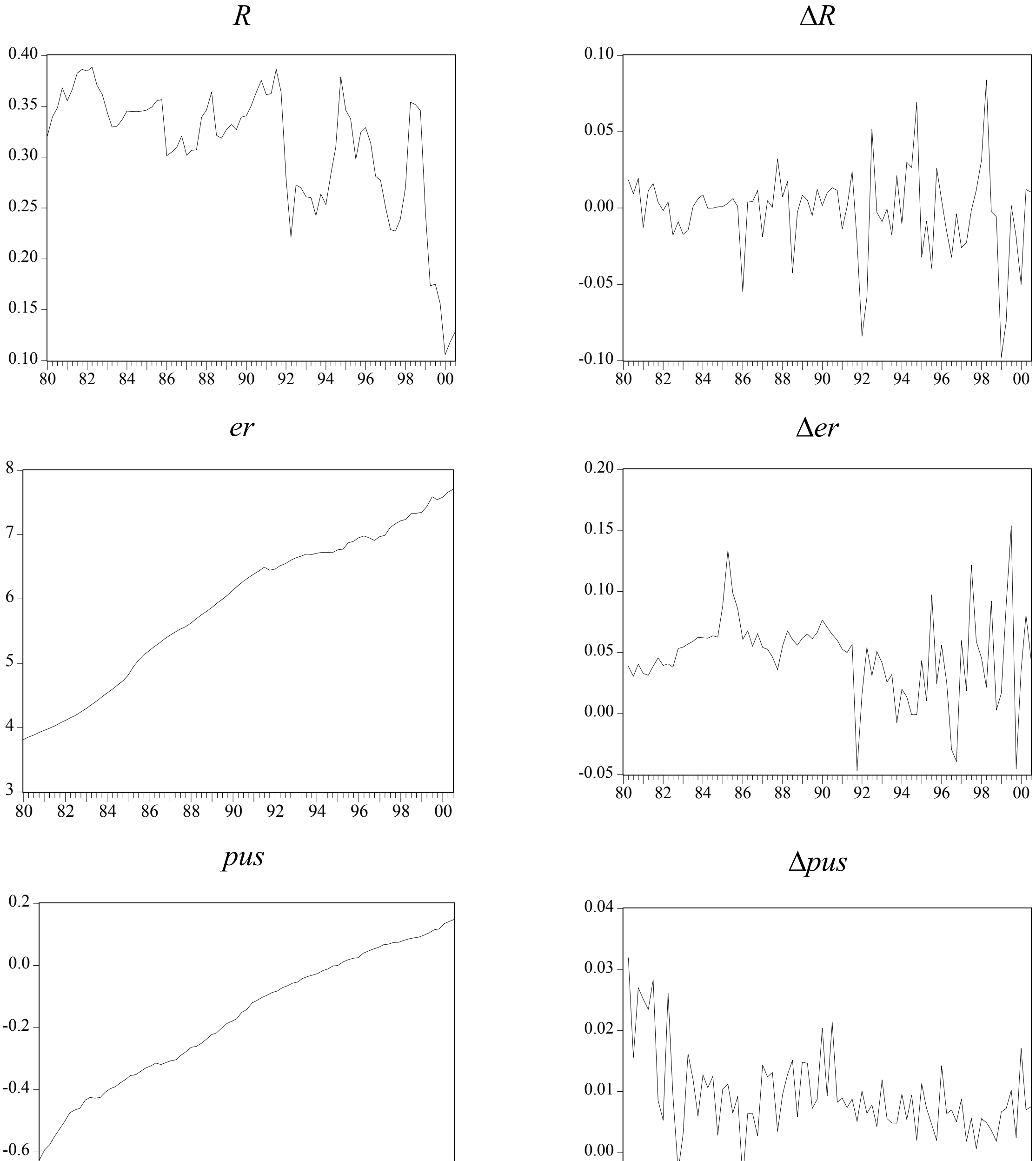


Figure 1 (Continued). Graphs of the series

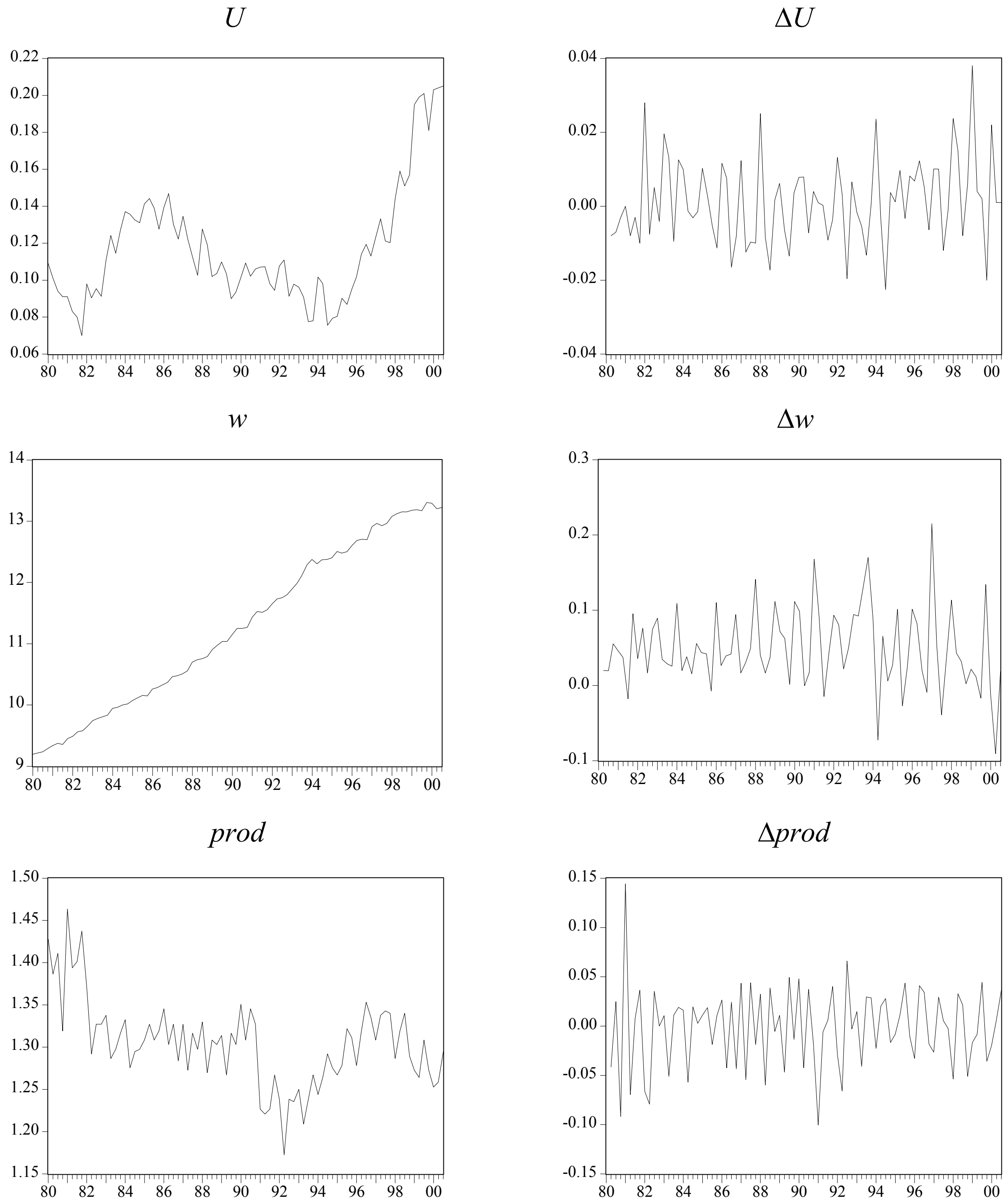
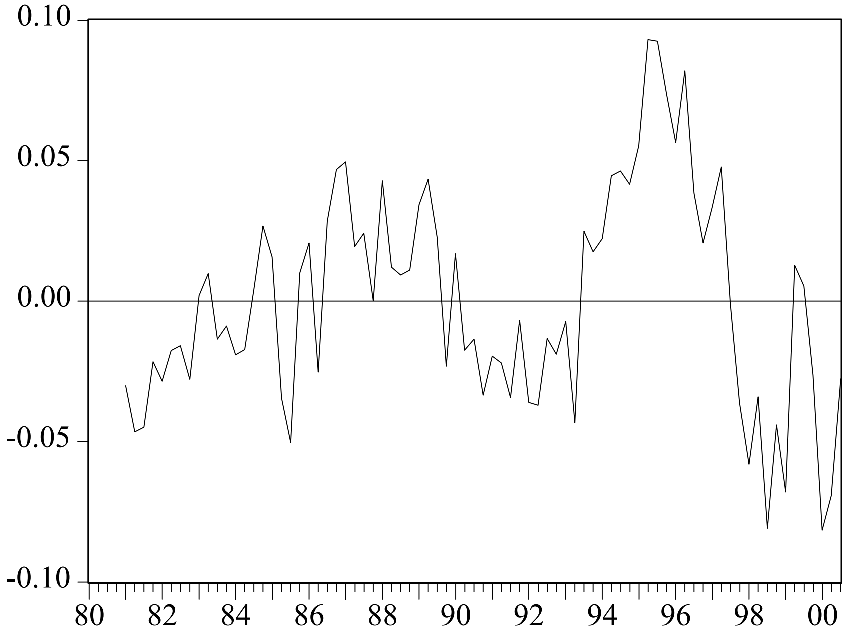


Figure 2. Restricted cointegrating vectors
(Partialling out the full-sample short-run dynamics)

(a). Excess money (*Mm*)



(b). Excess demand (*Gm*)

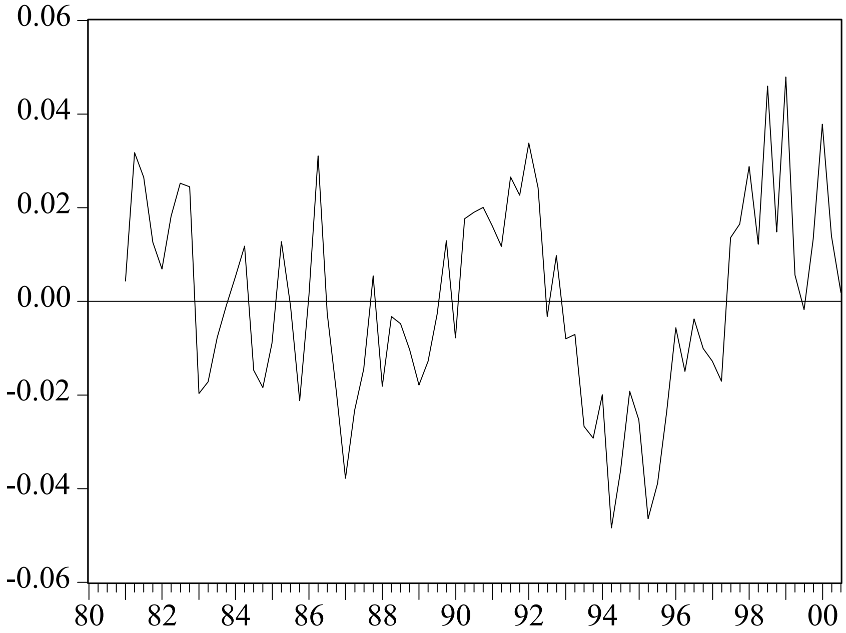
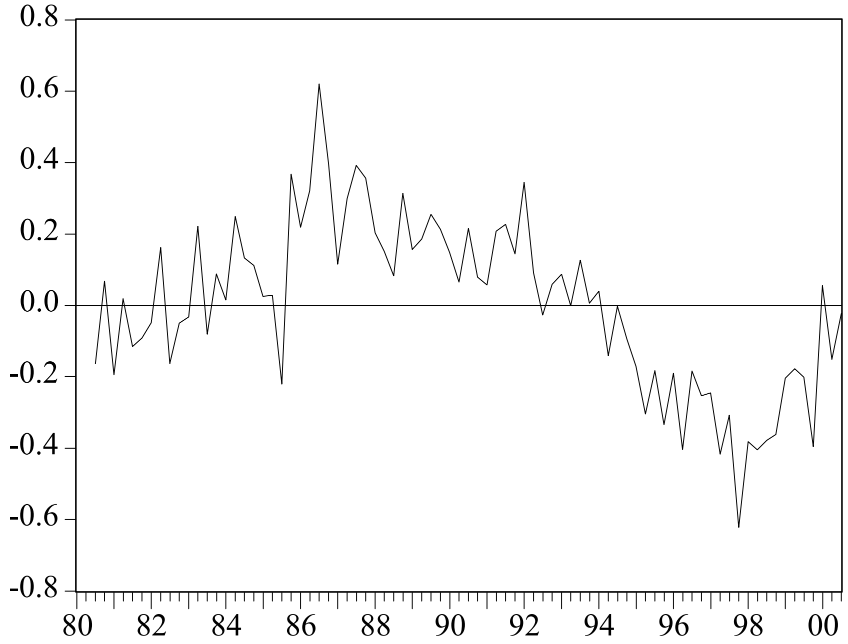


Figure 2 (Continued). Restricted cointegrating vectors
(Partialling out the short-run dynamics)

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



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

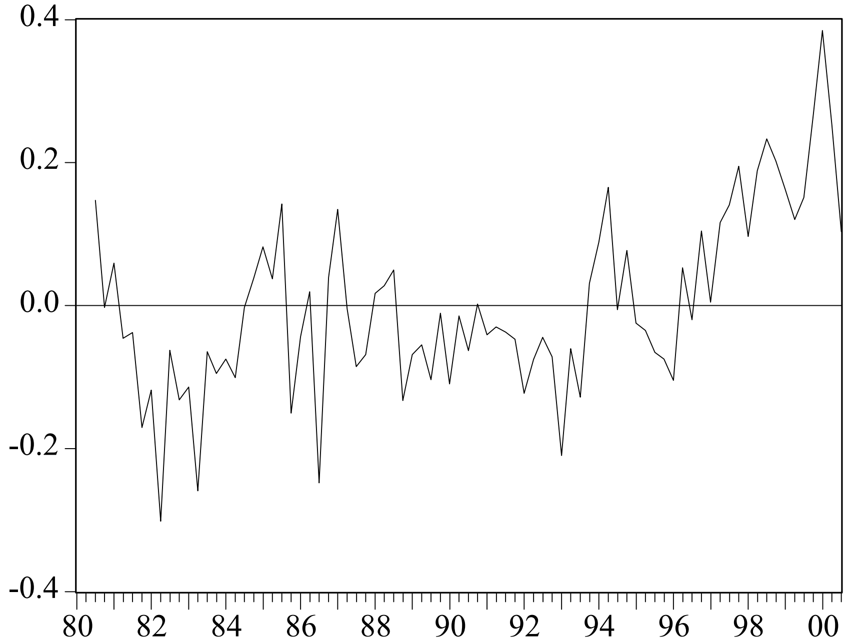
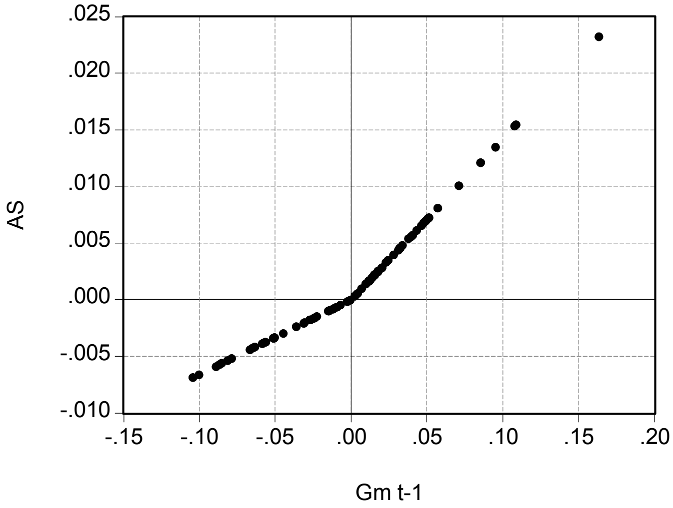
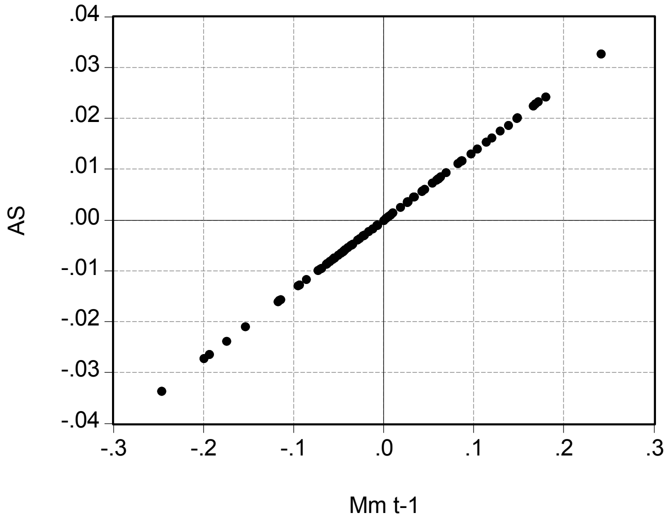


Figure 3. Asymmetric price adjustment in response to market disequilibria



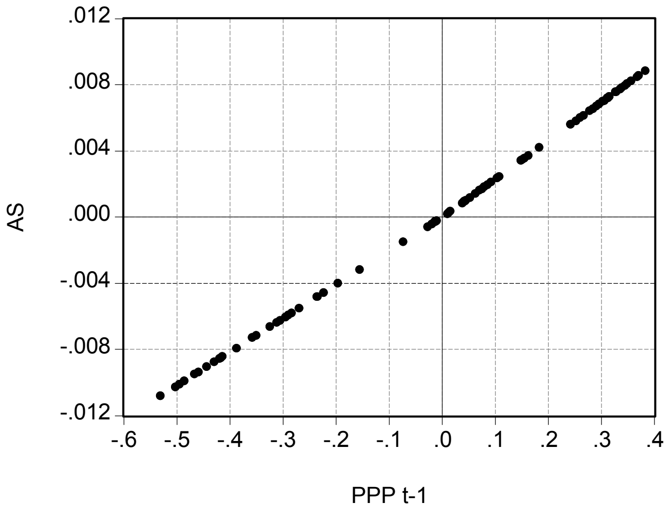
(a) The asymmetric adjustment in prices (AS) is calculated as:

$$0.142Gm^{+}_{t-1} + 0.066Gm^{-}_{t-1}.$$



(b) The asymmetric adjustment in prices (AS) is calculated as:

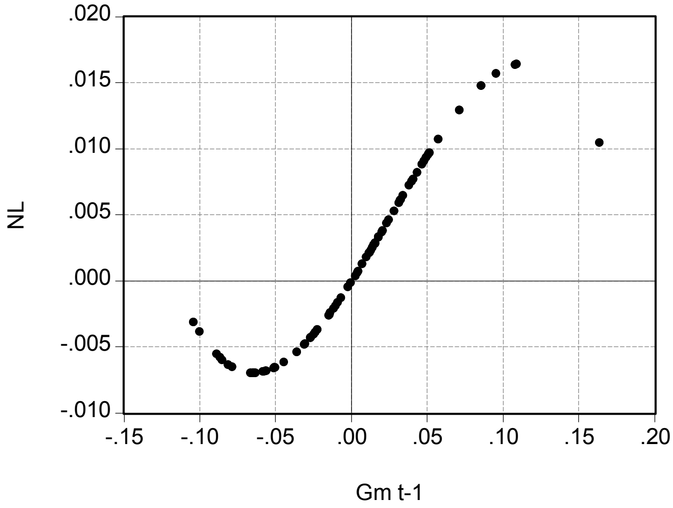
$$0.135Mm^{+}_{t-1} + 0.136Mm^{-}_{t-1}.$$



(c) The asymmetric adjustment in prices (AS) is calculated as:

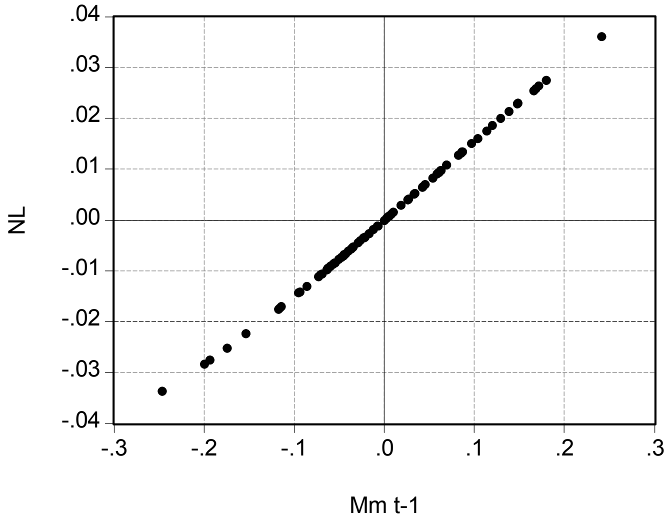
$$0.023PPP^{+}_{t-1} + 0.020PPP^{-}_{t-1}.$$

Figure 4. Non-linear price adjustment in response to market disequilibria



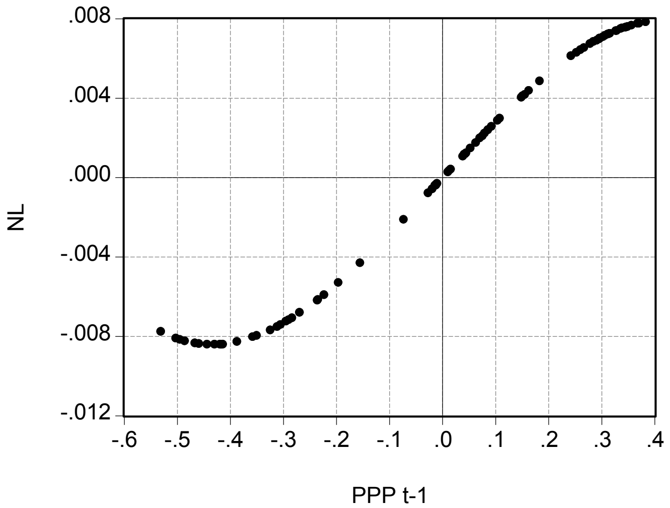
(a) The non-linear adjustment in prices (NL) is calculated as:

$$0.180Gm_{t-1} + 0.605Gm_{t-1}^{2} - 8.038Gm_{t-1}^{3}$$



(b) The non-linear adjustment in prices (NL) is calculated as:

$$0.155Mm_{t-1} + 0.027Mm_{t-1}^{2} - 0.195Mm_{t-1}^{3}$$



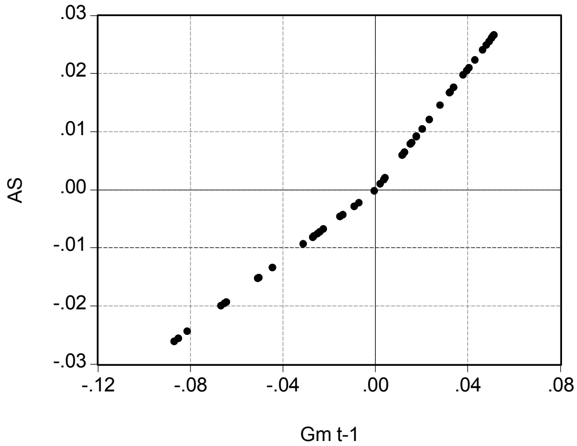
(c) The non-linear adjustment in prices (NL) is calculated as:

$$0.029PPP_{t-1} - 0.001PPP_{t-1}^{2} - 0.053PPP_{t-1}^{3}$$

Figure 5. Asymmetric price adjustment in response to market disequilibria

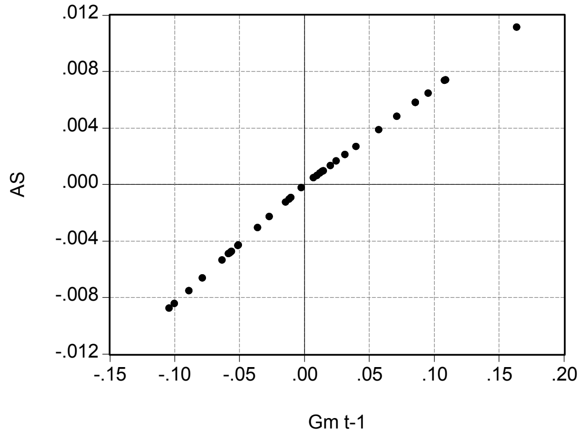
Before Central Bank Independence

(a)

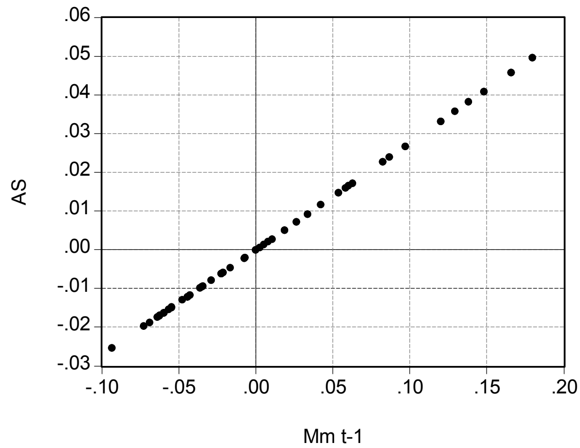


After Central Bank Independence

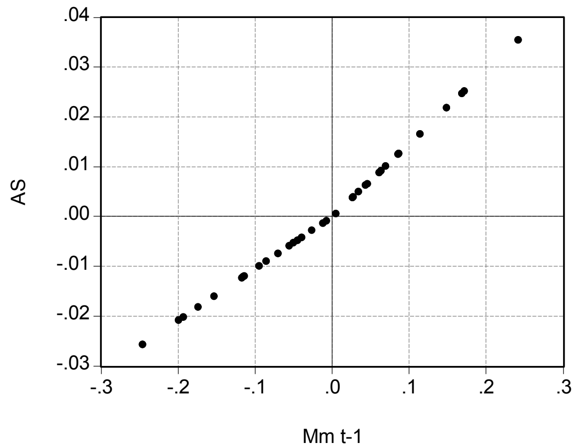
(b)



(c)



(d)



Notes:
The asymmetric (AS) adjustments in prices in Figures 5a,b are calculated as:

$0.521Gm_{t-1}^{+} + 0.299Gm_{t-1}^{-}$ (before Central Bank independence)

$0.068Gm_{t-1}^{+} + 0.084Gm_{t-1}^{-}$ (after Central Bank independence).

The asymmetric (AS) adjustments in prices in Figures 5c,d are calculated as:

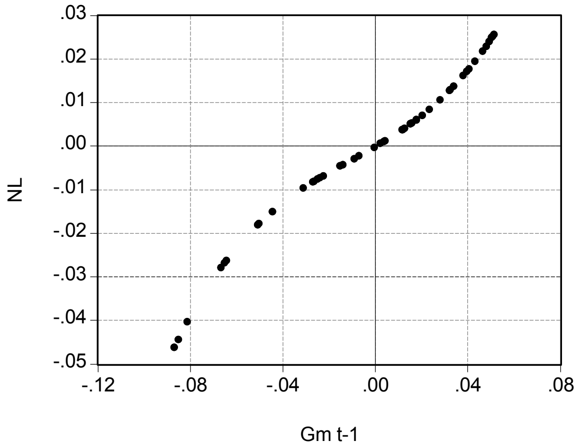
$0.277Mm_{t-1}^{+} + 0.270Mm_{t-1}^{-}$ (before Central Bank independence)

$0.147Mm_{t-1}^{+} + 0.104Mm_{t-1}^{-}$ (after Central Bank independence).

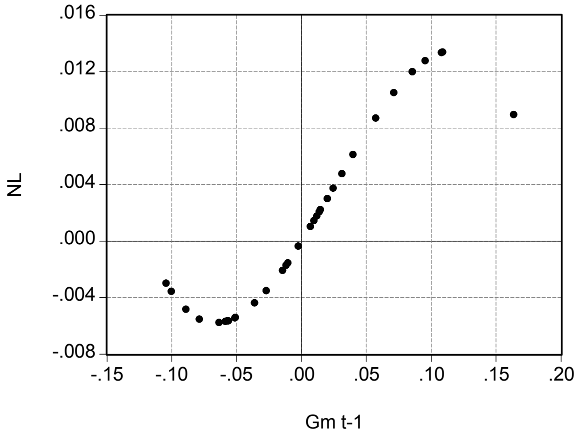
Figure 6. Non-linear price adjustment in response to market disequilibria

Before Central Bank Independence | After Central Bank Independence

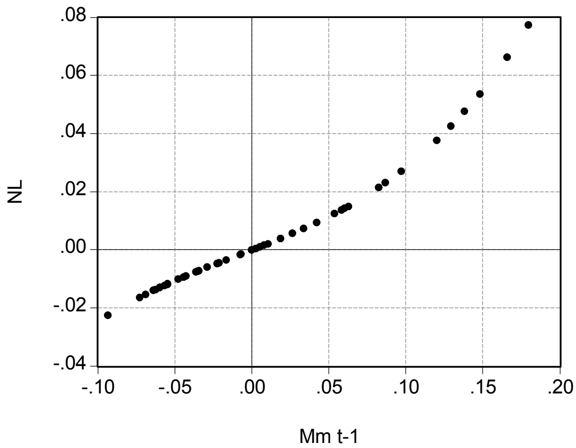
(a)



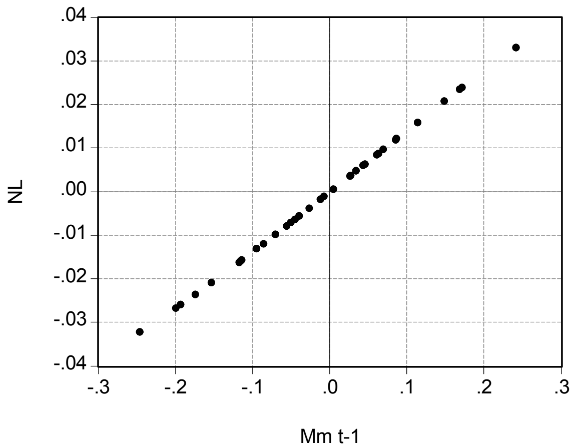
(b)



(c)



(d)



Notes:

The non-linear (NL) adjustments in prices in Figures 6a,b are calculated as:

$0.308Gm_{t-1} + 1.451Gm_{t-1}^2 + 45.942Gm_{t-1}^3$ (before Central Bank independence)

$0.146Gm_{t-1} + 0.472Gm_{t-1}^2 - 6.306Gm_{t-1}^3$ (after Central Bank independence).

The non-linear (NL) adjustments in prices in Figures 6c,d are calculated as:

$0.206Mm_{t-1} + 0.190Mm_{t-1}^2 + 5.924Mm_{t-1}^3$ (before Central Bank independence)

$0.140Mm_{t-1} + 0.014Mm_{t-1}^2 - 0.107Mm_{t-1}^3$ (after Central Bank independence).

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Data Appendix and Sources

M 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.

USP 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.

Figure A1. Recursive eigenvalues for the monetary sector model

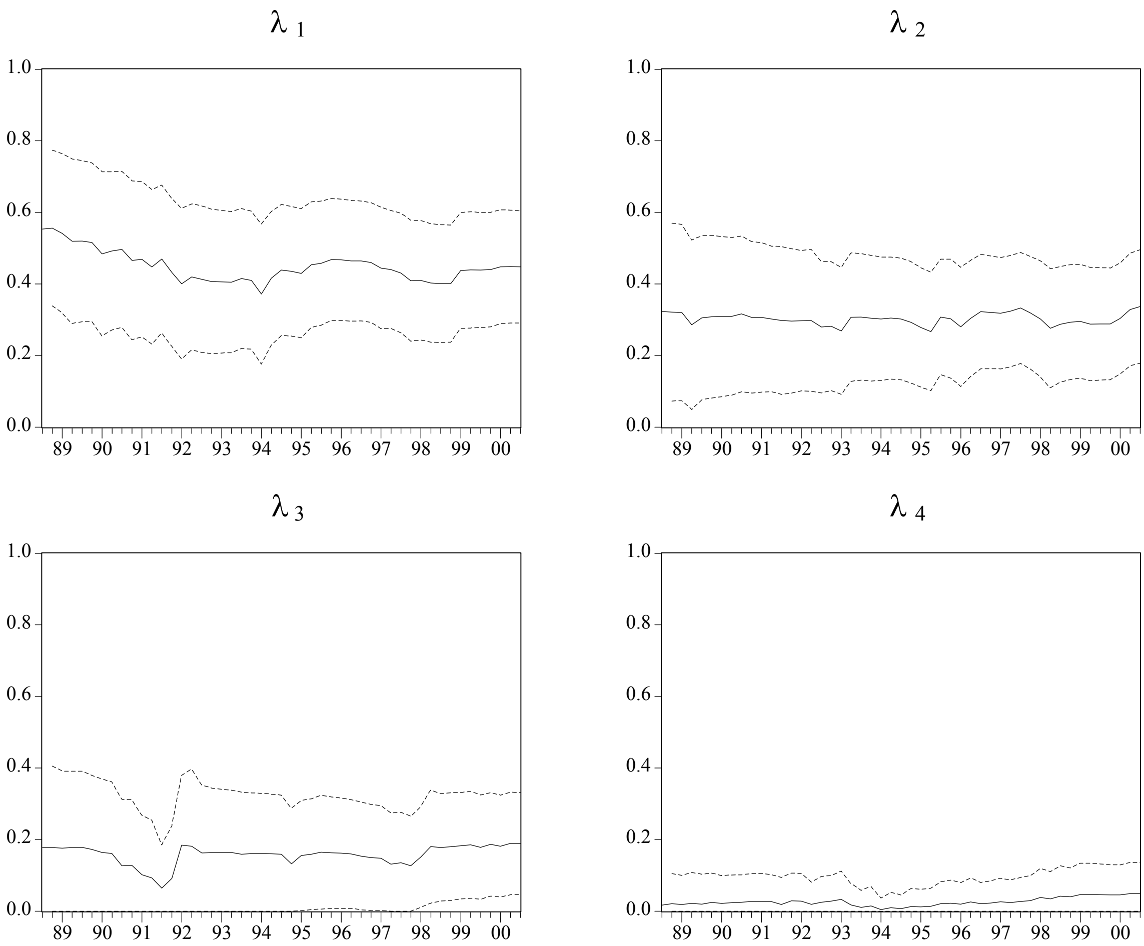
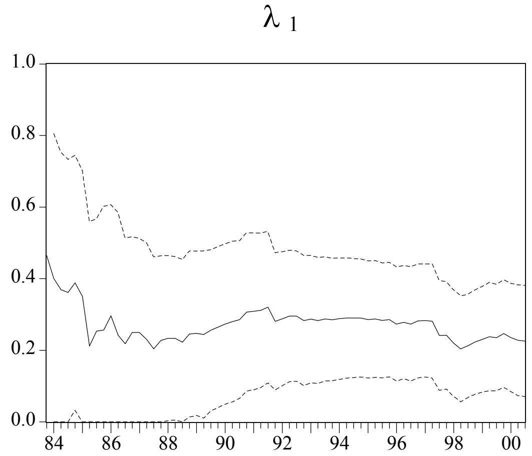


Figure A2. Recursive eigenvalues for the external sector model



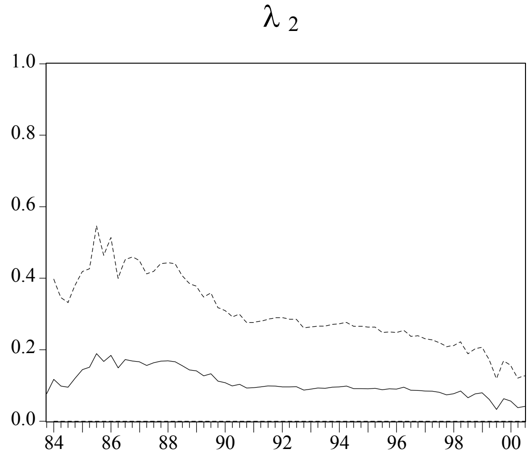


Figure A3. Recursive eigenvalues for the labour sector model