Inflation and Exchange Rate Regimes in Mexico

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Abstract

The paper presents a version of the exchange-rate-regime model of inflation. Quarterly data from Mexico from 1946 to 1995 are used to estimate and test a simultaneous-equation model for wage inflation, price inflation and industrial production, taking account of the Lucas critique and the statistical properties of the data. The main finding is that, after the fall of the fixed-exchange-rate regime in 1976, there was a Barro–Gordon type inflation bias owing to the inability of policymakers to commit to low inflation. There is no significant evidence of political business cycles in inflation.

1. Introduction

Persistent high inflation has been a common phenomenon in Latin America for decades. One of the possible reasons is the limited ability of governments to commit themselves to low inflation. In a monetary policy game between wage-setters and policymakers, Barro and Gordon (1983) have shown how this inability might result in an inefficient equilibrium characterized by high inflation and relatively low employment. An extension of the Barro–Gordon model is the exchange-rate regime model of Giavazzi and Giovannini (1987) and Giavazzi and Pagano (1988).1 In this model, participation in a regime of fixed exchange rates, in which monetary policy is determined by an inflation-averse center country, “ties the hands” of inflation-prone domestic policymakers. This is reflected in wage-setters’ expectations, and so the domestic economy ends up with the same average inflation as the center country.2

This paper presents a version of the exchange-rate-regime model of inflation, and investigates its applicability to the Mexican economy since 1946. Mexico is particularly suitable for such an investigation because it has had long experience with both fixed and independent managed floating exchange rates. Hence, it is interesting to examine whether the institutional constraints of the nominal exchange-rate regime matter for inflation, and possibly employment. The paper investigates the joint determination of wage inflation, price inflation, and industrial production by imposing the cross-equation restrictions of the Barro–Gordon model and the exchange-rate-regime model. By doing so, we respect the Lucas critique, and obtain estimates of the structural parameters of the model.3 As far as we know, this is the first attempt to estimate such a model for a Latin American economy.

The theoretical model is as follows. Under fixed exchange rates, there is no monetary policy independence; hence inflation is exogenously determined. By contrast, under managed floating exchange rates, there is room for policy independence. In
particular, under managed floating, inflation is endogenously determined via a Barro–Gordon game between policymakers and wage-setters.

The model leads to structural equations for wage inflation, price inflation, and employment. We test and estimate the model by using quarterly data from Mexico for the period 1946Q1–1995Q1. Our econometric work follows three steps. First, we introduce appropriate dummy variables to capture policy-induced parameter changes. Second, we investigate the statistical properties of the data. In particular, by following Perron (1989) and using the Zivot and Andrews (1992) unit-root test, we look for possible structural breaks. Univariate analysis indicates that price inflation displays a break around 1978 (almost two years after the official switch from fixed to managed floating exchange rates), while the cointegrating (Phillips-curve type) relationship between price inflation and industrial production displays a break around 1990 (when the government started to implement a series of stabilization policies to combat inflation). Third, given the statistical properties of the data, we test and estimate the model by using GMM. The theoretical cross-equation restrictions are not rejected by the data, and most estimated parameters are statistically significant and have plausible values.

The main result is that the exchange-rate regime matters for wage and price inflation. After the switch to managed floating in 1976, there is a significant Barro–Gordon type inflation bias owing to the inability of policymakers to commit to low inflation. As a result, wage and price inflation are higher, and significantly more persistent over time, during managed floating than during fixed exchange rates. The policy implication is that membership of an exchange-rate mechanism cannot make disinflation harder, since it can break inflation persistence (i.e., inertia). We also report that there is no significant evidence of electoral cycles in inflation, or differences in inflation performance across different administrations of the ruling political party.

2. Overview of the Mexican Economy

From the 1930s until the early 1970s, Mexico experienced a favorable position compared with other developing countries. Land reforms, import substitution industrialization, and investment in basic industries and infrastructure set the basis for economic growth. Government policy managed to reconcile the interests of the heterogeneous elite, to secure social support and a disciplined labor force, to protect domestic producers, and to make socioeconomic changes featuring close government ties with the industrial elite. As a result, during 1954–70, there was an average annual GDP growth of 7%, while average annual price inflation was only 3.5%. However, problems in the current account led to contractionary macroeconomic policies that caused a mild recession in 1971. This, added to a political crisis (erosion of government legitimacy and the guerrilla movement in Latin America), resulted in the implementation of “populist” macroeconomic policies in 1972.4

Mexico maintained a fixed peso–dollar exchange rate even after the collapse of the Bretton Woods system in 1971 and the world oil price shock in 1973. However, the distortions caused by import substitution policies, the public expenditure-led growth financed mainly by borrowing from the central bank (Banco de Mexico), the accumulation of foreign debt, high inflation, the overvaluation of the peso, and a large capital flight led finally to the devaluation of the peso in 1976. This basically ended the regime of fixed exchange rates.

The switch to managed floating exchange rates in 1976 gave more flexibility for discretionary economic policy. It was (wrongly) believed that demand management, in combination with price controls, could be used to cope with rising unemployment (the
average annual rate of urban unemployment was 8% in 1977, while underemployment was also strongly widespread. Nevertheless, the economic situation worsened: inflation (which was 17% on average during 1973–76) rose to 24% in the next five years; the public deficit to GDP ratio (which was 3% in 1977) increased to 6% in 1981; and the balance of payments further weakened (the current account deficit increased from 3.2 billion US dollars in 1978 to 16.1 billion US dollars in 1981). In 1982, the public deficit to GDP ratio was 15%, and inflation jumped to 59%. The economic collapse was precipitated by changes in the international environment. The sharp decline in oil prices, the US tight monetary policy, and the increase in international real interest rates pushed Mexico into a moratorium on repayments of foreign debt.

In response to the 1982 crisis, several stabilization and structural adjustment policies were implemented (there was also an earlier, but not successful, IMF stabilization plan in the mid-1970s). However, there were no signs of economic recovery and the situation was aggravated by an earthquake in 1985 and the collapse of oil prices. In 1986, inflation increased to 86% and continued to rise, the public deficit to GDP ratio rose again to 13%, while both real per capita GDP and real wages fell by 6%. Although GDP growth turned positive in 1987, the run of the peso (following the collapse of the New York stock market) and the persistency of high inflation forced policymakers to refocus more seriously on the fight against inflation.

In 1988, the government implemented “heterodox” policies to combat inflation. The Pacto de Solidaridad Económica was negotiated among workers, farmers, the industry, and the government. Wage settlements were gradually brought under control, prices of public goods and services remained frozen, and exchange-rate devaluations were reduced. The government agreed to cut its deficit, and promised tight monetary policy, privatization of public enterprises, and liberalization of the economy. Mexico negotiated a Brady Plan debt agreement and re-entered the international credit market in the early 1990s. Inflation was decreased, the fiscal stance was under control, and economic growth was restored. Nevertheless, current account deficits, negative shocks, inadequate economic policy, and a financial panic resulted in the collapse of the peso in the first quarter of 1995.

Finally, on the political front, up until the 1970s Mexico had one of the strongest civilian governments in Latin America. The success of the ruling political party (PRI) can be explained by its wide social representation, its system of clientilism and co-optation, and various political “rewards” (usually in the form of subsidies). Nevertheless, the accumulation of economic and social problems has recently weakened the traditional alliances of the Mexican political system.

3. The Theoretical Model

This section will set up a model consistent with the above stylized facts. The way the price level (or, equivalently, price inflation) is determined depends crucially on the nominal exchange-rate regime (McKinnon, 1993). Under fixed exchange rates, we assume that inflation is exogenously determined. By contrast, under managed floating exchange rates, we assume that policymakers are free to choose price inflation by playing a Barro–Gordon game with wage-setters. The sequence of events within each time-period is as follows: first, the current exogenous shocks are realized; second, wage-setters sign one-period nominal wage contracts; third, price inflation is determined; finally, firms set employment. Given this sequence, we assume (without any loss of generality) that the model is deterministic. Following the relevant literature, we use a log–linear model.
The Labour Market

Output, $y_t$, is produced by using labour, $\ell_t$, via a Cobb–Douglas production function, so that $y_t = \delta \ell_t + \mu_t$, where $0 < \delta < 1$. Productivity, $\mu_t$, follows the exogenous process $\mu_t = g + \mu_{t-1}$, where $g$ is a constant. Profit maximization by competitive firms leads to demand for labour:

$$\ell_t^d = -\beta(w_t - p_t - \mu_t),$$  \hspace{1cm} (1)

where $\beta \equiv 1/(1 - \delta) > 0$, $w_t$ is the nominal wage rate, and $p_t$ is the price level at time $t$.

Nominal wages are set by a group of insiders $\bar{n}_t$, where $\bar{n}_t$ follows an exogenous process (defined in section 4). On the other hand, policymakers care about the whole labour force $n_t$, where $n_t > \bar{n}_t$. Insiders, $\bar{n}_t$, choose $w_t$ to solve:

$$\min E_t(\ell_t^d - \bar{n}_t)^2,$$  \hspace{1cm} (2)

where the minimization is subject to equation (1). Here, $E_t$ is the rational expectations operator. Since the model is deterministic, and moves are sequential, it is only $p_t$ that is not observable by wage-setters when they choose $w_t$.

Using the first-order condition of (2) into (1), employment, $\ell_t$, is:

$$\ell_t = \bar{n}_t + \beta (\Delta p_t - E_t \Delta p_t),$$  \hspace{1cm} (3)

while it is convenient to write wage inflation, $\Delta w_t \equiv (w_t - w_{t-1})$, as:

$$\Delta w_t = g + E_t \Delta p_t - (1/\beta)(\bar{n}_t - \ell_{t-1}),$$  \hspace{1cm} (4)

where $\Delta p_t \equiv (p_t - p_{t-1})$ is price inflation and $E_t \Delta p_t \equiv (E_t p_t - p_{t-1})$ is expected price inflation. Equation (3) is a standard aggregate supply function saying that deviations of actual employment from its natural level, $\bar{n}_t$, are due to inflation surprises. Equation (4) is a Phillips-curve type expression saying that wage inflation decreases when employment has been below its natural level, i.e., $\ell_{t-1} < \bar{n}_t$.

Solution under Fixed Exchange Rates

Under fixed exchange rates (denoted by the superscript $f$), we assume that price inflation is exogenously determined. Intuitively, under fixed exchange rates, the domestic authorities do not have the independence of determining the average domestic inflation rate. This was especially true under the rules of the Bretton Woods system during which the average OECD inflation rate was determined by the monetary policy of the Federal Reserve System in the USA. Mexico maintained a fixed exchange rate against the US dollar until 1976Q3.

Therefore, we assume $\Delta p_t^f = \pi_t^f$, where $\pi_t^f$ is an exogenous process (defined in section 4). Then, by using (3) and (4), we simply have for price inflation, wage inflation, and employment, respectively:

$$\Delta p_t^f = \pi_t^f,$$  \hspace{1cm} (5a)

$$\Delta w_t^f = g + \pi_t^f - (1/\beta)(\bar{n}_t - \ell_{t-1}),$$  \hspace{1cm} (5b)

$$\ell_t^f = \bar{n}_t.$$  \hspace{1cm} (5c)

Solution under Managed Floating Exchange Rates

Under managed floating exchange rates (denoted by the superscript $m$), we assume that price inflation is endogenously determined via a Barro–Gordon game between policymakers and wage-setters.

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Policymakers choose $\Delta p^m_t$ to solve:

$$\min \left[ (\Delta p^m_t - \pi^m_t)^2 + \alpha (\ell_t - n_t)^2 \right],$$

where the minimization is subject to equation (3). Here, as in Canzoneri (1985), $\alpha > 0$ is the weight given to employment relative to inflation, and $n_t = \kappa \bar{n}_t$ where $\kappa > 1$. Comparison of (2) with (6) implies that the employment level sought by wage-setters is too small from the point of view of policymakers; this is why $\kappa > 1$. Also, $\pi^m_t$ is the policymakers’ exogenous target rate of price inflation under managed floating (defined in section 4).

It is well known that the equilibrium of this one-shot sequential-move game is the Nash equilibrium. Then, we have for price inflation, wage inflation, and employment, respectively:

$$\Delta p^m_t = \pi^m_t + \alpha \beta (\kappa - 1) \bar{n}_t,$$  

$$\Delta w^m_t = g + \pi^m_t + \alpha \beta (\kappa - 1) \bar{n}_t - (1/\beta)(\bar{n}_t - \ell_{t-1}),$$  

$$\ell^m_t = \bar{n}_t.$$  

To proceed, we make the (testable) assumption $\pi^m_t > \pi^f_t$; i.e., the target of price inflation under managed floating is higher than the exogenously determined price inflation under fixed exchange rates. Then, equations (5a)–(5c) and (7a)–(7c) imply the following:

**Remark 1:** Both price and wage inflation are lower under fixed exchange rates than under managed floating exchange rates.

**Remark 2:** Employment is always at its natural level.

The first remark follows from the assumption that there is a Barro–Gordon game only under managed floating, and the assumption $\pi^m_t > \pi^f_t$. The second remark follows from the assumption that there is no uncertainty about the exchange-rate regime, and so the model has classical features. Although classical neutrality is a strong result especially for a country like Mexico, we prefer to keep the model simple so as to concentrate on the main issue; i.e., the role of exchange-rate policy, and hence the role of monetary policy independence, for wage and price inflation.

4. **Empirical Specification and Results**

This section presents the econometric specification of the model and discusses testing and estimation by using quarterly data from Mexico over 1946Q1–1995Q1. We will work in three steps. First, we specify the exogenous targets, and introduce dummy variables to capture policy-induced parameter changes. Second, we investigate the statistical properties of the data, and look for possible structural breaks. Third, we estimate and test the econometric model.

**Exogenous Targets and Policy-Induced Parameter Changes**

This subsection specifies the exogenous, unobservable variables ($\bar{n}_t, \pi^f_t, \pi^m_t$). It also introduces appropriate dummies to capture changes in monetary policy across exchange-rate regimes.

To model the employment target of wage-setters ($\bar{n}_t$), we assume that $\bar{n}_t$ is a linear function of lagged-once employment ($\ell_{t-1}$). Intuitively, it is those who are employed at the time the wage contract is signed (called insiders) that matter for wage-setting. This
is as in Blanchard and Summers (1986) and Alogoskoufis and Manning (1988). Thus, we assume:\(^10\)

\[
\hat{n}_t = \lambda \ell_{t-1},
\]

(8)

where \(0 \leq \lambda \leq 1\) measures the power of insiders in wage-setting.

To model the exogenously determined price inflation under fixed exchange rates (\(\pi_f^t\)) and the policy target for price inflation under managed floating (\(\pi_m^t\)), we assume that \(\pi_f^t\) and \(\pi_m^t\) are linear functions (with a constant) of lagged-once inflation. Intuitively, the higher the inherited inflation rate, the more costly the policy adjustment required to reduce it to an arbitrarily low level (in equations (9a) and (9b), these levels are \(\pi_f^t\) and \(\pi_m^t\), respectively). In other words, policymakers modify their targets to take into account these adjustment costs. Thus, we assume:

\[
\begin{align*}
\pi_f^t &= \pi_f^t + \theta_f^t \Delta p_{t-1}, \\
\pi_m^t &= \pi_m^t + \theta_m^t \Delta p_{t-1},
\end{align*}
\]

(9a)  (9b)

where we make the (testable) assumptions \(\pi_f^t < \pi_m^t\) and \(0 \leq \theta_f^t < \theta_m^t \leq 1\). These assumptions are consistent with Remark 1.

In order to capture policy-induced parameter changes, we will use dummy variables. Define \(d_f^t\) to be a dummy for fixed exchange rates that takes the value of 1 during 1946Q1–1976Q3 and during 1988, and zero otherwise. Similarly, define \(d_m^t\) as \(1 - d_f^t\); i.e., \(d_m^t\) takes the value of 1 during managed floating exchange rates, and zero otherwise.

By substituting (8), (9a)–(9b), and the above dummy variables into (5) and (7), the wage, price, and employment equations during all exchange rate regimes can be summarized by:

\[
\begin{align*}
\Delta \omega_t &= g_{w1}d_f^t + g_{w2}d_m^t + g_{w3}d_f^t \Delta p_{t-1} + g_{w4}d_m^t \Delta p_{t-1} + g_{w5} \ell_{t-1} + g_{w6} d_f^t \ell_{t-1}, \\
\Delta p_t &= g_{p1}d_f^t + g_{p2}d_m^t + g_{p3}d_f^t \Delta p_{t-1} + g_{p4}d_m^t \Delta p_{t-1} + g_{p5} d_m^t \ell_{t-1}, \\
\ell_t &= g_{\ell1} \ell_{t-1},
\end{align*}
\]

(10a)  (10b)  (10c)

where the theoretical restrictions are:

\[
\begin{align*}
g_{w1} &= (g + \pi_f^t), & g_{w2} &= (g + \pi_m^t), & g_{w3} &= \theta_f^t, & g_{w4} &= \theta_m^t, \\
g_{w5} &= (1 - \lambda)/\beta, & g_{w6} &= \alpha \beta (\kappa - 1) \lambda, \\
g_{p1} &= \pi_f^t, & g_{p2} &= \pi_m^t, & g_{p3} &= \theta_f^t, & g_{p4} &= \theta_m^t, & g_{p5} &= \alpha \beta (\kappa - 1) \lambda, & g_{\ell1} &= \lambda.
\end{align*}
\]

(10d)

Data Sources, Integration, Cointegration, and Structural Breaks

Our data source is Estadísticas Históricas de México published by the Instituto Nacional de Estadística, Geografía e Informática (INEGI, Mexico). The data are quarterly. For \(\ell_t\), we will use data on real industrial production (\(y_t\)).\(^{11}\) Concerning \(\Delta \omega_t\), since there are no consistent time-series on urban wages for the period under study, we will use the average minimum wage of Baja California, California Sur, Chihuahua, Distrito Federal, and urban areas of the States of Sonora, Tamaulipas, and Veracruz.\(^{12}\) Finally, \(\Delta p_t\) will be based on the consumption price index (CPI).

Augmented Dickey–Fuller (ADF) \(t\)-statistics testing for unit roots in \(\Delta \omega_t\), \(\Delta p_t\), and \(y_t\), (allowing for four lags, and a constant or trend when necessary) have values of \(-4.14\), \(-2.84\), and \(-0.28\), respectively. The last two values are higher than the 5% ADF critical value of \(-2.86\), so that the null hypothesis of a unit root cannot be rejected for \(\Delta p_t\) and \(y_t\). When we test for second-order integration, the null of a unit root is rejected...
when $\Delta p_t$ and $y_t$ are first-differenced. Taken together, these results suggest that $\Delta w_t$ is $I(0)$, while $\Delta p_t$ and $y_t$ are $I(1)$.

However, it is well known that ADF tests tend to favor the null of unit roots when the true process is characterized by structural breaks (Perron, 1989). To circumvent this problem, we carry out the Zivot and Andrews (1992) sequential ADF test, which allows a possible break to be endogenously determined by the data. Under the null, the process is assumed to be structurally stable with a unit root and trend. Under the alternative, the process is assumed to be stationary with breaks in its mean and/or trend. The break point is selected to correspond to the minimum value of the relevant, one-sided unit-root $t$-statistic denoted as $\text{Inf ADF}$. The Zivot–Andrews results are reported in Table 1.

The test shows that $\Delta p_t$ and $y_t$ are $I(1)$, even if we take account of structural breaks in the deterministic components of the data. In particular, concerning $\Delta p_t$, there is evidence of a break in intercept where the break point is estimated to be at 1978Q3. In Mexico, this date almost coincides with the switch from fixed exchange rates to managed floating exchange rates. Although this is two years after the official regime switch which took place in 1976Q3, the delayed response implied by the data can be attributed to various adjustment mechanisms. Concerning $y_t$, there is weak evidence of a break in intercept at 1974Q1 (based on the 5% critical value of tests for the intercept in unit-root regressions). There is also evidence of a time-trend without break. Therefore, the Zivot–Andrews tests confirm the presence of unit roots in both $\Delta p_t$ and $y_t$.

Since each of $\Delta p_t$ and $y_t$ taken individually is $I(1)$, we now look for a cointegrating (long-run) relationship between them. The theoretical model, in particular equation (10b), implies a long-run relationship between $\Delta p_t$ and $y_t$ only during the managed floating-exchange-rate regime. Therefore, we look for a cointegrating relationship between $\Delta p_t$ and $d_m^t y_t$, where recall that $d_m^t$ is a zero-one dummy for managed floating. To investigate whether there is a break in the intercept of the cointegrating regression, we use the Gregory and Hansen (1996) test. The break point is selected in

### Table 1. Zivot–Andrews Unit-Root Tests for Mexico, 1946Q1–1995Q1

| $\Delta x_t = \alpha_0 + \alpha_1 d(\lambda t) + \alpha_2 t + \phi x_{t-1} + \sum_{i=1}^{k} y_i \Delta x_{t-i} + v_t$, where $x_t = (\Delta p_t, y_t)$ |
|---|---|---|
| $\Delta p_t$ | $y_t$ |
| $\alpha_0$ | 0.05 | 0.00 |
| | (1.56) | (0.06) |
| $\alpha_1$ | 0.02 | 0.04 |
| | (2.94) | (2.28) |
| $\alpha_2$ | — | 0.001 |
| | — | (3.20) |
| $\phi$ | $-0.23$ | $-0.11$ |
| | (4.12) | (3.14) |
| $\text{Inf ADF}$ | $-4.12$ | $-3.45$ |
| Date of break | 1978Q3 | 1974Q1 |

**Notes:** The date of break is the date of break in the intercept. The tests involved four lags, $k = 4$. The 5% critical value for the $\text{Inf ADF}$ test is $-4.80$ for the regression without the trend, and $-4.42$ for the regression with the trend. Absolute values of $t$-statistics are reported in parentheses.
a similar way to the Zivot–Andrews test; i.e. we choose the smallest value of the ADF test of the residuals of the cointegrating regression over all possible break points. Table 2 reports OLS estimates without, and with, a structural break.

The results show that while a cointegrating relationship between $D_p t$ and $d m_t y_t$ can be only marginally accepted in the absence of a structural break, it is easily accepted once we allow for a structural break (Table 2). Further support for the cointegrating regression with break is provided by the increased value of the $R^2$. After allowing for a break, the results imply that the cointegrating vector is $(1, \ -0.048)$, normalized at $D_p t$. They also imply that the break in intercept happened around 1990Q1. We will therefore include a dummy variable (denoted by $d_t 1990s$) in equation (10b) when we estimate the model. The dummy $d_t 1990s$ takes the value of 1 after 1990Q1, and zero otherwise.

Recall from the overview of the Mexican economy in section 2 that this break-point coincides with a number of important events. For instance, the government has implemented various stabilization programs to combat inflation since 1990.

The Econometric Model and Estimation Results

Having investigated the long-run properties of the data, we now estimate the theoretical model (10a)–(10c) subject to (10d). In the light of the results in the previous subsection, equation (10a) constitutes a structural relationship between $D w_t$, which is $I(0)$, and the cointegrating relationship between $D p_t$ and $y_t$. Next consider equation (10b). Subtracting $D p_{t-1}$ from both sides of (10b), and taking into account the cointegrating relationship between $D p_t$ and $y_t$, we can rewrite (10b) as an error-correction type equation between the change in inflation ($\Delta p_t - \Delta p_{t-1}$), which is $I(0)$, and the residuals of the cointegrating relationship between $D p_{t-1}$ and $y_{t-1}$. Finally, consider equation (10c). According to the theoretical model, industrial production depends only on its lagged-once value. Also, the univariate statistical analysis in Table 1 showed that $y_t$ has a unit root (with an intercept with break, and a trend). Taken together, these results essentially mean that $\lambda = 1$ (see the restrictions in (10d)). Hence,
to be consistent with the unit-root analysis, we impose $\lambda = 1$ upon the whole system (10a)–(10d).\(^{14}\)

Then, we estimate the following econometric model:

\[
\Delta w_t = g_{w1} d_1^f + g_{w2} d_1^m + g_{w3} d_1^f \Delta p_{t-1} + g_{w4} d_1^m \Delta p_{t-1} + g_{w5} d_1^m y_{t-1},
\]

\[
(\Delta p_t - \Delta p_{t-1}) = g_\rho d_1^f + g_\rho d_1^m + g_\rho d_1^f \Delta p_{t-1} + g_\rho d_1^m \Delta p_{t-1} + g_\rho d_1^m y_{t-1} + g_\rho d_1^{1990s} + g_\rho d_1^{debt} + g_\rho d_1^{oil} + g_\rho d_1^{liber} + g_\rho d_1^{peso},
\]

\[
\Delta y_t = \text{const} + g_{y1}^{rend} + g_{y2}^{debt},
\]

where the cross-equation restrictions are now:

\[
g_{w1} = (g + \pi^f), \quad g_{w2} = (g + \pi^m), \quad g_{w3} = \theta^f, \quad g_{w4} = \theta^m, \quad g_{w5} = \alpha\beta(\kappa - 1),
\]

\[
g_{\rho1} = \pi^f, \quad g_{\rho2} = \pi^m, \quad g_{\rho3} = (\theta^f - 1), \quad g_{\rho4} = (\theta^m - 1), \quad g_{\rho5} = \alpha\beta(\kappa - 1).
\]

Observe that in this econometric model, we have also included a number of unrestrictive additive dummies that influence the dynamics of (10a)–(10c). The inclusion of these dummies does not affect our main results. However, it improves the fit of the regressions, and is also consistent with the main (exogenous to the theoretical model) events that have characterized the Mexican economy (section 2). In particular, in the regression for price inflation (11b), we have included (in addition to $d_1^{1990s}$ which follows from the cointegrating regression) a dummy for the 1982 debt crisis (called $d_1^{debt}$), a dummy for the oil price shocks (called $d_1^{oil}$), a dummy for various liberalization policy attempts (called $d_1^{liber}$), and a dummy to account for those short time-periods characterized by peso-type expectations about exchange rate depreciations (called $d_1^{peso}$). In the regression for industrial production (11c), we have added $d_1^{debt}$ to capture the one-shot effect of the debt crisis.\(^{15}\)

The econometric model (11a)–(11c), subject to the cross-equation restrictions (11d), is estimated by using GMM. This enables us to identify the structural parameters $g$, $\pi^f$, $\pi^m$, $\theta^f$, $\theta^m$, $A$ (notice that since $A \equiv \alpha\beta(\kappa - 1)$, the coefficients $\alpha$, $\beta$, and $\kappa$ cannot be identified). As instruments, we use the explanatory variables in (11a)–(11c). As it is known, GMM estimates are consistent and asymptotically normally distributed. Thus, although the normality test rejects the null of normal errors in all regressions, the test statistics have the right asymptotic properties. The estimation results are reported in Table 3. The $LM (4)$ and $ARCH (1)$ misspecification tests show that the standard errors of the parameters have been corrected for serial correlation of order 4 and heteroskedasticity.

The results clearly support the theoretical model. In particular, the data cannot reject the over-identified restrictions in (11d) at the 9% significance level. Note that the number of the over-identifying restrictions implied by (11d) is 22, which equals the number of orthogonality conditions between (11a)–(11c) and the instrumental variables minus the number of parameters to be estimated.

Most of the estimated parameters have signs and magnitudes consistent with the theory and are statistically significant. In particular, $\pi^f = 0.009 < \pi^m = 0.025$ (although the fixed component of the quarterly rate of inflation under managed floating, $\pi^m$, is not significant at the 5% level). There is clear evidence that $\theta^f = 0.385 < \theta^m = 0.818$, which means that price and wage inflation persist more over time during managed floating than during fixed exchange rates. These results are consistent with the theoretical predictions, and are also similar to the estimates of Alogoskoufis et al. (1992) for the UK and Alogoskoufis et al. (1998) for Greece. The only crucial parameter that has sign opposite from that predicted by the theory is $A$, which is insignificant. In par-
ticular, recall that $A \int_{ab} (k - 1)$ should be positive. The negative sign of $A$ can be due to negative Phillips-curve type effects from unemployment to current inflation which more than outweigh the positive Barro–Gordon type effects.16

It is worth noting that the GMM estimate of the long-run elasticity of price inflation with respect to output in (11a)–(11d) is consistent with the analogous OLS estimate in the cointegrating regression with break of Table 2. This can be easily seen from the GMM estimates in Table 3. If we set the left-hand side of (11b) equal to zero and solve for $D_p$, then the resulting long-run coefficient on $d_p' y_t$ is equal to 0.049, which is very close to the OLS estimate of $\beta = 0.048$ in the regression with break of Table 2. Notice that the cointegrating regression in Table 2 gives a significant $\beta$, which in turn implies a significant $A$. By contrast, the GMM estimates in Table 3 give an insignificant $A$ at 5% level. However, the insignificance of $A$ in Table 3 has to be interpreted cautiously because of the small sample problems of the GMM estimates. By contrast, the estimate of the cointegrating coefficient $\beta$ is super-consistent.

Finally, notice that the unrestricted dummy variables have the expected effects. For instance, note the detrimental effects of the 1982 debt crisis on price inflation and output, and the inflationary effects of liberalization policies and peso-type expectations. Also note the strong disinflationary effect of recent stabilization policies being captured by the 1990s dummy (recall that the inclusion of this dummy follows from the cointegrating regression with break of Table 2).

Before closing this section, we would like to report some additional results concerning political business cycles (available upon request).17 First, there is no

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Estimate</th>
<th>$t$-statistic</th>
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<tr>
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<tr>
<td>$\pi_f$</td>
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<tr>
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</tr>
<tr>
<td>$\theta_f$</td>
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</tr>
<tr>
<td>$\theta_m$</td>
<td>0.818</td>
<td>14.12</td>
</tr>
<tr>
<td>$A$</td>
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<td>-0.61</td>
</tr>
<tr>
<td>$g_{D_p}$</td>
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<td>-3.25</td>
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</tr>
<tr>
<td>$g_{D}$</td>
<td>-0.061</td>
<td>-1.82</td>
</tr>
</tbody>
</table>

$J (22) = 31.05$

$L M_{(11a)} (4) = 33.23$

$L M_{(11b)} (4) = 29.59$

$L M_{(11c)} (4) = 87.69$

$ARCH_{(11a)} (1) = 16.39$

$ARCH_{(11b)} (1) = 46.63$

$ARCH_{(11c)} (1) = 32.75$

Notes: The estimated equations were (11a)–11(c) subject to (11d). See text for definitions of regressors. $J (22)$ is the over-identified restrictions test-statistic distributed as $\chi^2 (22)$. $LM_{(11a)} (4)$, $LM_{(11b)} (4)$, and $LM_{(11c)} (4)$ are the $LM$ test-statistics for serial correlation of order four distributed as $\chi^2 (4)$ for equations (11a), (11b), and (11c) respectively. $ARCH_{(11a)} (1)$, $ARCH_{(11b)} (1)$, and $ARCH_{(11c)} (1)$ are the ARCH test-statistics for conditional heteroskedasticity distributed as $\chi^2 (1)$ for equations (11a), (11b), and (11c) respectively.
statistically significant evidence of electoral cycles. In particular, when we tested for differences in inflation over the electoral cycle, there was only weak evidence that inflation was higher immediately after elections than in other periods of the electoral cycle. Second, there was no significant evidence of differences in inflation performance across different administrations of the ruling political party (PRI). In particular, various possible “partisan” differences were tested, but they were not found to be statistically significant.

5. Conclusions, Policy Implications, and Extensions

In this paper we have presented a version of the exchange-rate-regime model of inflation. When we tested and estimated the model for the Mexican economy during 1946Q1–1995Q1, the data could not reject its theoretical restrictions and give plausible estimates for the structural parameters.

Although our results are consistent with the widely accepted view that exchange rate policy affects the inflation process, we want to emphasize that they do not imply that low inflation can be attributed only to exchange rate policy. A country can clearly disinflrate on its own. However, the interesting question is not whether a country can disinflrate on its own or not, but whether exchange rate discipline can make the disinflation process less costly. To the degree it can provide a form of incomes policy or an anchor for expectations and so break inflation inertia (which—according to the estimate for $q^m$—seems to be very strong in Mexico), it cannot make disinflation more costly. If, however, nominal wages are driven without regard for inflation and exchange rate targets, any use of exchange rate commitments will result, sooner or later, in overvaluation and speculative attacks.

We close with some possible extensions. First, it would be interesting to include a public sector, and so examine how wages and employment in this sector affect inflation and employment in the private sector. Second, it would be interesting to introduce uncertainty about the exchange rate regime, and examine the implications of expected regime switches for inflation and employment (Hamilton, 1994). We leave these extensions for future work.

References


Instituto Nacional de Estadística, Geografía e Informática (INEGI), Series Históricas de México, several issues.


Notes

1. Participation in a regime of exchange-rate commitments is only one possible institutional solution to the Barro–Gordon outcome. Appointing an inflation-averse central banker is another institutional solution (Rogoff, 1985a). For a recent survey of this literature, see Svenson (1997).
2. Coles and Philippopoulos (1997) have extended this imported credibility argument to a target-zone model of exchange rates. The crucial assumption is that exchange rate commitments (e.g., fixed exchange rates and target zones) are more credible than domestic anti-inflationary policies.
3. Our estimation methodology is similar to that in Alesina and Sachs (1988) and Alogoskoufis et al. (1998). Specifically, Alesina and Sachs (1988) test and estimate the partisan model of Alesina (1987) for the USA. Alogoskoufis et al. (1998) combine the exchange-rate-regime model with the partisan model, and then test and estimate it for Greece. Here, we focus on exchange-rate regimes.
5. Ibarra and Alberro (1989) discuss the economic policy during this period.
8. For simplicity, the constant term ln\(d\) is omitted.
9. Notice that exchange rates (actual and expected) can be easily introduced. For instance, we can use a small open economy model with purchasing power parity. Then, our solution remains unchanged, if we replace domestic prices with exchange rates (Horn and Persson, 1988). This is especially true in a country like Mexico in which the traded sector is particularly important, and hence there is a strong link between price inflation and exchange rate depreciation. Alternatively, we can use a more general model with traded and nontraded goods. Again, this would not change our qualitative results (Rogoff, 1985b).
10. We could use more complicated targets, e.g., \(\bar{\eta}_t = \lambda \ell_{t-1} + (1 - \lambda)\eta_t\). This would not change our main results.
11. We could find consistent quarterly data on GDP only since 1980. Our results do not change much when we construct real quarterly data on GDP by using real annual GDP with weights provided by real quarterly industrial production.
12. See Ross (1980) for the correlation between the average manufacture wage and the minimum urban wage in Mexico.
13. We can report that our results do not change when we look for a cointegrating relationship between \(\Delta p_t\) and \(y_t\); i.e. during the whole sample period.
14. We can report that the restriction \(\lambda = 1\) is not rejected by the data.
15. The debt crisis dummy (\(d^{debt}\)) takes the value of 1 during 1982, and 0 otherwise. The oil price shock dummy (\(d^{oil}\)) takes the value of 1 during 1973–74 and 1979–80, and 0 otherwise. The trade liberalization dummy (\(d^{liber}\)) takes the value of 1 from 1985.Q3 onwards (the first serious attempt to reduce trade restrictions took place in July 1985), and 0 otherwise. The peso-type expectations dummy (\(d^{peso}\)) takes the value of 1 in 1987.Q4 and 1994.Q4, and 0 otherwise. We have experimented with various combinations of dummy variables in all three regressions (11c), (11b), (11c), but here we include only those which are significant. Also recall from Table 1 that there was weak evidence of a break in intercept at 1974Q1 in the regression for industrial production (11c). When we estimated the model (11a)–(11d) including a dummy that takes the value of 1 after 1974Q1 and zero otherwise into (11c), this was again found to be insignificant. We therefore omit it from the reported results in Table 3.
16. In the Barro–Gordon model of inflation in which policymakers have direct control over price inflation, and hence use it as a policy instrument, there is always a positive effect from unemployment to current price inflation. See equation (7a).
17. For details see Li et al. (1997) where we develop and test a richer version of the model than the one we are using here that includes electoral effects. See also Alogoskoufis et al. (1992) for the UK and Alogoskoufis et al. (1998) for Greece. Since there is only insignificant evidence of political business cycles, we have decided just to report the results, so as to keep the model simple.
and economize on space. However, it is worth noting that most of the empirical evidence on political business cycles involves democracies in industrial countries. There are few exceptions. For instance, Edwards (1994) nests the predictions of the naive and rational political business cycles models in an inflation equation, and tests it for Chile in the period 1952–73. He finds evidence of opportunistic (i.e., electoral) and partisan effects, but no support for any particular model. Remmer (1993) finds limited evidence of electoral effects on a number of macroeconomic variables. His results are based on single-equation regressions for a sample of Latin American countries. By contrast, our results reported here are based on a fully specified model with cross-equation restrictions.