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Short-Run Price Dynamics in Mexico

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Abstract

A theoretical model of inflationary dynamics is developed for Mexico. Estimation is accomplished using transfer function ARIMA analysis. Monthly data for the January 1982 - December 1998 sample period are utilized in the empirical analysis. Overall econometric characteristics of the nonlinear regression output comply with standard statistical criteria. Results indicate that prices respond relatively quickly to changes in money supply, interest rate, wage, and peso/dollar exchange rate variables.

INTRODUCTION

Similar to other Latin American economies, monetary authorities in Mexico have waged a fairly intense battle to control prices in recent years. Although inflation has been successfully brought down from the levels observed in 1995 aftermath of the "tequila effect devaluation," prospects for additional near-term improvements remain uncertain (Lara and Beltrán del Río, 1999). As a means for examining the prospects for likely month-to-month movements in the consumer price index, a theoretical modeling framework is developed. Short-term dynamic characteristics are then examined utilizing transfer function ARIMA techniques (Liu, 1987; Fullerton 1993).

In attempts to slow price movements, recent governments in Mexico have introduced a variety of new policy measures. These include import liberalization, fiscal austerity, and slower rates of currency depreciation. By reducing price pressures, the government hopes to improve economic welfare by enabling the economy to operate more efficiently. This argument is similar to those being aired in advanced economies such as the United States (Motley, 1998) and being analyzed in other developing nations (Zind, 1993). Subsequent sections of the study present a review of the literature, theoretical model, and empirical results. Suggestions for future research are summarized in the conclusion.

PREVIOUS RESEARCH

Early research on inflationary dynamics in developing countries was conducted using Chilean data (Harberger, 1963). That paper employs percentage rates of change of the variables of interest in a linear regression framework based on the quantity theory of money. What became known as the “Harberger” framework incorporates real income, current and lagged values of the money supply, and the opportunity cost of holding cash balances. The success of this initial effort conducted on Chilean data spurred a series of replicated studies for other developing countries.

Hanson (1985) extends the Harberger framework in a systematic fashion to incorporate an important missing component, import costs. An implicit cost function is utilized to derive an aggregate supply curve which includes local prices of imported inputs. When the underlying production function is homogeneous of degree one, inflation becomes a weighted sum of money supply changes and import prices. This is important for studies using higher frequency data if the problem of measurement bias engendered by interpolated values of real output, generally published on either a quarterly or annual basis in developing countries, is to be avoided (see Bomberger and Makinen, 1979). The model also implies the elasticity of inflation with respect to money growth is less than one.

Subsequent research has provided additional evidence in favor of the augmented Harberger-Hanson approach wherein the effect of import prices on inflation is modeled. Koch, Rosensweig, and Witt (1988) and Fullerton, Hirth, and Smith (1991) uncover positive linkages between the trade-weighted currency indexes and consumer prices in the United States. Brajer (1992) provides statistical evidence that omission of exchange rate or import deflators will generally result in specification errors for developing country inflationary models. Fullerton (1993) examines Colombian price stabilization efforts utilizing monthly data that include both monetary and exchange rate variables. Recent research for Mexico uncovers similar results using annual data, but does not examine short-term price trends (Fullerton and Calderón, 1999).

THEORETICAL MODEL

Harberger's (1963) model is based on the traditional quantity theory of money macroeconomic equation of exchange:

(1) $MV = PQ,$

where M represents some measure of the money stock, V is the velocity of circulation, P is the nominal price level, and Q is real output. Velocity is assumed to be a predictable function of other macroeconomic variables such as the implicit cost of holding idle cash balances.

Transformation by natural logarithms and first differences creates a version of Equation 1 calculated in percentage changes. Introduction of a time subscript and rearrangement yields the basic Harberger equation:

(2) $DLP_t = DLM_t - DLQ_t + DLI_{t-1}$,

where the last term results from substituting for velocity with a foregone interest earnings measure and D represents a difference or backshift lag operator. Usage of a lagged change in the inflation rate as an interest proxy may be necessary to calculate the implicit cost of holding money in cases when government regulations on interest rates cause savings and loan rates to become negative in real terms. Unadjusted interest rates from such periods would fail to provide accurate estimates of the opportunity cost of holding idle cash. To ensure cointegration, whichever interest measure selected can be multiplied with national savings or the money supply to construct the foregone interest earnings estimate. Such a step is recommended since consumer prices are unbounded from above, but nominal interest rates generally oscillate between 0 and 100 percent (for an alternative strategy, see Elder and Kennedy, 2001).

Equation 2 indicates that inflation will vary positively with respect to the money supply and inversely with respect to real output. A statistically significant intercept term will enter the estimated equation if there is a trend in the velocity of circulation. If only contemporaneous lags of LM and LQ enter in the equation, the parameters for both variables are hypothesized to be unitary. This possibility can be empirically examined using the following specification:

(3) $DLP_t = a_0 + a_1DLM_t - a_2DLQ_t + a_3DLI_{t-1} + u_{3t}$,

where a_1 and a_3 are hypothesized to be positive, and the absolute values of a_1 and a_2 should both be statistically indistinguishable from one. The last argument in the expression represents a stochastic disturbance term.

Hanson (1985) and Fullerton (1999) utilize an implicit cost function dual of an aggregate production function that is homogeneous of degree one. Derived output supply functions from this framework will be homogeneous of degree zero in input and output prices. Equation 4 represents this relationship using logarithmic first differences:

(4) $DLQ_t = b_0 + b_1DLP_t - b_2DLW_t - b_3DLPI_t + u_{4t}$,

where W stands for wages paid to labor and PI represents imported input prices. Increases in the relative prices of imported inputs are assumed to lead to declines in domestic output. The standard homogeneity assumptions for production and derived supply relations imply that $b_1 - b_2 - b_3 = 0$.

Substitution of Equation 4 into Equation 3 eliminates the output term from the expression to be estimated. Because monthly estimates of gross domestic product do not exist for Mexico, this step is useful. The resulting equation can be written as follows:

(5) $(1 + a_2b_1)DLP_t = a_0 - a_2b_0 + a_1DLM_t + a_2b_2DLW_t + a_2b_3DLPI_t + a_3DLI_{t-1} + u_{5t}$.

Equation 5 can be simplified prior to estimation. Dividing through by the left-hand side constant term such that the aggregate price index remains as the dependent variable yields the following relation:

(6) $DLP_t = c_0 + c_1DLM_t + c_2DLW_t + c_3DLPI_t + c_4DLI_{t-1} + u_{6t}$,

which also has testable properties. The most important change in terms of model characteristics is that the coefficient on the monetary variable, c_1, is now hypothesized to be smaller than one in the statistical sense. With the possible exception of the intercept, all of the regression parameters in Equation 6 are expected to exceed zero.

As discussed in the literature review, Equation 6 has provided a useful framework for analyzing both quarterly and annual inflation rates. Because the lag structure in this model is fairly short, it may require additional modification prior to estimation. This possibility does not reflect inherent deficiencies in the theoretical model as such, but arises due to the fact that this study relies upon monthly data. As a result, if the inflationary impact of a change in the money supply is felt over the course of one calendar year, the implied lag structure for a model estimated with data published at a monthly frequency would potentially range up to 12 months in length. Equation 7 reflects this potential empirical issue and also substitutes for the import price deflator with an exchange rate measure. The latter change is mandated by the availability of the exchange rate at a monthly frequency in Mexico. Said modification should not lead to empirical difficulties because import price deflators usually are highly correlated with currency values. Additionally, the interplay between rates of depreciation and inflation in Mexico is fairly well established (Fullerton and Calderón, 1999; Ibarra, 1999; Li, Philippopoulos, and Tzavalis, 2000).

(7) $DLP_t = c_0 + Sum(c_{1i}DLM_{t-i}) + Sum\ (c_{2j}DLW_{t-j}) + Sum(c_{3k}DLX_{t-k}) + Sum(c_{4h}DLI_{t-1-h}) + u_{7t}$.

The above model provides an attractive starting point for examining inflationary trends in an economy. It is not, however, without potential problems for analyzing price movements in a relatively high inflation country such as Mexico. The principal concern arises from utilizing first differenced, log-transformed time series data in the equation to be estimated. If the resulting series are stationary, the equation can be estimated with reduced risk of obtaining spurious correlation in the results. As shown in studies of hyperinflationary economies, however, higher order differencing may be required to induce stationarity during periods in which prices increase very rapidly (Engsted, 1993). Because Mexico has never experienced hyperinflation, first differences will likely remove nonstationary trends from the first moments of the variables in question. If this assumption is valid, the transfer function version of the equation is shown below, inclusive of univariate autoregressive and moving average components:

$$(8) \quad DLP_t = c_0 + Sum(c_{1i}DLM_{t-i}) + Sum(c_{2j}DLW_{t-j}) + Sum(c_{3k}DLX_{t-k}) + Sum(c_{4h}DLI_{t-1-h}) + Sum(c_{5p}DLP_{t-p}) + Sum(c_{6q}u_{t-q}) + v_{8t}.$$

A secondary concern arises that is also related to the manner in which the theoretical model is specified. Namely, the equation form shown above assumes unidirectional causality from the independent regressors to the left-hand side variable. Granger F-tests are utilized to assess whether such an assumption is warranted. Should the results of the causality tests indicate simultaneity exists between the dependent and independent variables, transfer ARIMA techniques would be inappropriate. Potential alternative candidate approaches include atheoretic constructs (Leiderman, 1984; Hamilton, 1994) as well as multi-equation structural models (Fullerton and Araki, 1996).

Finally, it should be noted that the above model has its roots in the traditionally utilized monetary approach that has long supported aggregate price trends research and policy analysis. In recent years, this basic framework has periodically come under attack (Meltzer, 1998). Empirical research has generally supported such theoretical treatment of nominal price variations (Rolnick and Weber, 1997; Dwyer and Hafer, 1999). Ultimately, the empirical applicability of this or other inflationary models is an issue that can only be resolved on a case-by-case basis. Such verification is attempted with respect to Mexico in the material that follows.

EMPIRICAL ANALYSIS

Table 1 lists variable names and data sources. In order to examine whether the working series included in Equation 8 are stationary, a battery of unit root tests are utilized. Estimation is conducted for the January 1982 - December 1998 sample period for which complete data are available. Applying unit-root tests to a relatively short time span of this nature may be risky due to the fact that these tests typically have low power unless long-run data sets are used (Hakkio and Rush, 1991). Because this is also the period for which both interest rates and the exchange rate are flexible in Mexico, there is little that can be done to circumvent this potential problem.

TABLE 1 APPROXIMATELY HERE

Augmented Dickey-Fuller t-statistics were estimated for equations without intercepts; with constant terms, but without trend variables; and with both intercepts and trends (the latter results are reported in Table 2). In all cases, tests for unit roots in the first differenced log transformed series for consumer prices, money, wages, the exchange rate, and the foregone interest earnings proxy are rejected at the 1-

percent level. Based on this evidence, the first-order differenced series used to estimate Equation 8 are assumed to be stationary.

TABLE 2 APPROXIMATELY HERE

As specified above, the model is explicitly built around a set of unidirectional causality relations from movements in the regressors to consumer prices. To examine whether the absence of simultaneity in the model is plausible, a set of Granger causality F-tests were calculated for the stationary components of the series of interest for lags of 6, 12, 18, and 24 months. Results associated with these tests rendered surprisingly strong evidence in favor of movements in the independent variables preceding subsequent changes in the CPI in a statistically significant manner (outcomes for 24 lags appear in Table 3). This is in contrast to the mixed evidence uncovered for the 1971-1997 sample period in Mexico (Fullerton and Calderón, 1999) and possibly reflects greater autonomy on the part of the central bank during the past decade and a half.

TABLE 3 APPROXIMATELY HERE

Initial lag length selection was conducted on the basis of two sets of cross correlation functions. The first set was calculated for the stationary components of the series while the second were estimated using the residuals from univariate ARIMA equations estimated for each respective variable (Mills, 1990). Table 4 summarizes the univariate models estimated for each of the variables utilized. As might be expected, all of the univariate equations included statistically significant positive intercepts, implying that these series contain strong deterministic components over the sample period in question. Final lag selection was determined using likelihood ratio tests estimated from the various specifications considered.

TABLE 4 APPROXIMATELY HERE

Nonlinear regression results for Equation 8 are summarized in Table 5. All of input variables are found to affect consumer prices in Mexico within fairly short time periods of 12 months or less. For the cost of holding cash balances regressor, lags of 1 and 3 months are included in the estimated equation. Lags of the M1 money supply measure appear at lags 2, 3, and 10. Movements in the exchange rate are found to affect price levels after lags of 1, 5, 6, and 12 months. The rapid transmission of currency valuation changes to aggregate prices is in line with the "whiplash scenario" hypothesis explored in recent inflationary targeting studies (Svensson, 2000). Nominal wage coefficients are estimated at lags of 4 and 9 months. All of the coefficients carry the hypothesized algebraic signs discussed above.

TABLE 5 APPROXIMATELY HERE

Because the data have been differenced prior to estimation, the coefficient of determination for the dependent variable used in modeling is lower than might be expected for macroeconomic time series (R-squared = 0.664143). Adjusting the fitted data back to level form, however, allows calculating a “pseudo” coefficient of determination. This measure indicates that the model explains more than 99 percent of the variation in the consumer price index over the sample period in question (January 1982 - December 1998).

Although many of the econometric results in Table 5 exhibit good statistical traits, five of the computed t-statistics do not satisfy the 5-percent type-I error criterion. Exclusion of those regressors from the equation causes the log-likelihood ratio to decline substantially. Given the high level of significance associated with the F-statistic for the model, plus the fact that it includes sixteen separate regression coefficients, the low t-statistics reflect the presence of multicollinearity. All of the input variable categories include at least one lag that is statistically significant. Encouragingly, the equation does not overlook any systematic movements in consumer prices for the sample period in question. The Q-statistic for white noise, random movements in the residuals, at 24 lags is 12.322. The latter measure follows a chi-square distribution.

Because the model is estimated using differenced data, the positive sign on the intercept confirms that a general upward trend exists in the Mexican consumer price index. As hypothesized by Hanson (1985), the sum of the coefficients for the lagged monetary series is significantly less than one. Similar to Fullerton (1993), but unlike Kamas (1995), the exchange rate appears to play an important role in determining aggregate price trends. That result is also in contrast to theoretical arguments previously developed for Latin American economies (Herrera, 1985). As with Fullerton and Calderón (1999), at least partial evidence is also shown that money supply, velocity of circulation, and wage variables help explain price trends in Mexico. Lag structures shown in Table 5 also match well with earlier studies for Latin America and Mexico (Leiderman, 1984; Fullerton, 1993; Kamas, 1995).

Despite the potential pitfalls identified in earlier sections of the paper, the empirical characteristics shown in Table 5 are generally in line with what might be expected from a well-specified underlying theoretical model. Several advantages arise from the fact that empirical regularities earlier identified with both quarterly and annual frequency data also emerge from the monthly data utilized herein. Much economic policy design in Mexico is implemented with the goal of obtaining short-range objectives (Ortiz Martínez, 1999). The model developed herein provides one means for assessing whether attaining those goals under the stated time frame is, in fact, feasible. In an environment where inflation targeting (Bernanke and Mishkin, 1997) is becoming widespread, such a tool is potentially very helpful. Because monetary policy potentially carries with it sharp output reductions (Filardo, 1998), accurate insight regarding aggregate price dynamics is complementary to effective central bank decision making.

CONCLUSION

An empirical model of monthly consumer price inflation in Mexico is developed and estimated by incorporating both monetary and input cost effects in a theoretically plausible manner. Because the model does not pose stringent data requirements, it is also applicable to other Latin American economies where inflation remains a problem. Examples include Costa Rica, Ecuador, and Venezuela where authorities continue to face short-term inflationary risks. The methodology is also attractive from the perspective of providing a means of examining the rapidity with which the various policy instruments are likely to influence nominal prices.

Additional econometric testing will undoubtedly prove useful. The Mexican economy is in the midst of ongoing structural change and policy experimentation. Parameter heterogeneity, therefore, remains a distinct possibility. It may also be useful to expand the scope of the model to include additional equations that allow for simultaneous treatment of inflation, liquidity, wages, and exchange rates. Results obtained in this paper indicate that the latter are not required for parameter estimation consistency, but they could enrich subsequent policy simulation analyses.

These suggested changes represent avenues for refinement to the basic framework outlined above. They are not likely to result in wholesale alterations to the general approach. Consequently, it is not anticipated that the conclusions reached above will change markedly due to expanding the scope of the empirical analysis. Given the breadth of economic conditions prevailing across Latin America, steps in these directions may prove helpful in future econometric research of this nature. The spread of macroeconomic price index targeting will also require the development of tools such as those discussed herein in order to gauge the intensity with which anti-inflationary measures must be imposed. Finally, the adoption of formal inflation targets will logically raise questions about the implied costs in terms of foregone output and employment. The latter suggests a valuable new track for short-range policy research and analysis in developing economies worldwide.

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TABLE 1
Variable Names and Data Description

Series	Definition
DLP	First Difference of the Natural Logarithm of the Consumer Price Index, Monthly Average, 1995 = 100
DLM	First Difference of the Natural Logarithm of the M1 Money Supply, Millions of Pesos
DLX	First Difference of the Natural Logarithm of the Nominal Exchange Rate, Monthly Average, Pesos per Dollar
DLW	First Difference of Natural Logarithm of Nominal Wage Index, Pesos Per Month, Monthly Average, 1995 = 100
DLI	First Difference of Natural Logarithm of Foregone Interest Earnings Approximated by Product of M1 Money Supply with Average Commercial Bank System Deposit Rate

Notes
1. Sample Period: January 1982 - December 1998.
2. Source: International Monetary Fund, *International Financial Statistics*, various issues, with missing interest rate estimates for four months obtained from Banco de México, Ciudad Juárez Branch Office Research Department.

TABLE 2
Augmented Dickey-Fuller Unit Root Stationarity Tests

Series	ADF Test Statistic	1% MacKinnon Critical Value
DLP	-13.620	-3.465
DLM	-12.391	
DLX	-12.090	
DLW	-13.935	
DLI	-9.686	

Notes
1. Sample period: January 1982 - December 1998.
2. Test form: Unit root equations used for this table included four lags, constant, and time trend regressors.

TABLE 3
Pairwise Granger Causality Tests

Null Hypothesis	Observations	F-Statistic	Probability
DLM does not precede DLP	179	1.782	0.022
DLP does not precede DLM	179	0.968	0.512
DLX does not precede DLP	179	2.139	0.004
DLP does not precede DLX	179	0.316	0.999
DLW does not precede DLP	179	2.080	0.005
DLP does not precede DLW	179	1.482	0.085
DLI does not precede DLP	179	1.855	0.015
DLP does not precede DLI	179	0.760	0.779

Notes
1. Sample period: January 1982 - December 1998.
2. Test form: Granger causality equations for this table included 24 lags.

TABLE 4
Univariate ARIMA Model Specifications

Series		Specifications	
DLP	Consumer Price Index	AR(8, 12)	MA(1, 12)
DLM	M1 Money Supply	AR(12)	MA(1, 6, 12)
DLX	Nominal Exchange Rate	AR(1, 5, 9)	
DLW	Nominal Wage Index	AR(12)	MA(1, 5)
DLI	Foregone Interest Earnings	AR(4, 12)	MA(1)

Notes
1. Sample period: January 1982 - December 1998.
2. Univariate ARIMA equation residuals utilized in cross correlation function selection of lag structure employed in transfer function shown in Table 5.

TABLE 5
Transfer ARIMA Empirical Output

Variable	Coefficient	Std. Error	t-Statistic	Probability
C	0.009162	0.002246	4.079885	0.0001
DLI (-1)	0.045995	0.011267	4.082450	0.0001
DLI (-3)	0.013410	0.012071	1.110937	0.2682
DLM (-2)	0.057358	0.018567	3.089184	0.0024
DLM (-3)	0.038974	0.021113	1.845960	0.0667
DLM (-10)	0.017517	0.016733	1.046843	0.2967
DLX (-1)	0.102198	0.035401	2.886833	0.0044
DLX (-5)	0.143079	0.029795	4.802089	0.0000
DLX (-6)	0.113781	0.030626	3.715180	0.0003
DLX (-12)	0.158434	0.031749	4.990254	0.0000
DLW (-4)	0.113563	0.025824	4.397619	0.0000
DLW (-9)	0.104733	0.024825	4.218929	0.0000
AR (8)	0.009890	0.075187	0.131533	0.8955
AR (12)	0.079177	0.046814	1.691309	0.0926
MA (1)	-0.739459	0.044898	-16.46974	0.0000
MA (12)	-0.224123	0.044986	-4.982102	0.0000

R-squared	0.664143	**Mean dependent var**	0.027076
Pseudo R-squared	0.998330	**S.D. dependent var**	0.025735
S.E. of regression	0.015586	**Akaike criterion**	-8.237669
Sum squared resid	0.039595	**Schwarz criterion**	-7.952763
Log likelihood	499.2814	**F-statistic**	21.48834
Durbin-Watson	1.941454	**F-stat. Probability**	0.000000
Iterations	12	**Q(24) white noise**	12.32200
Observations	191	**Q-statistic Probability**	0.950000

Notes
1. Sample period: January 1982 - December 1998.
2. Dependent variable: DLP, the first difference of the natural logarithm of the consumer price index in Mexico.

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