Inflation and Imported Input Prices in Some Inflationary Latin American Economies
INFLATION AND IMPORTED INPUT PRICES IN SOME INFLATIONARY LATIN AMERICAN ECONOMIES

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Received September 1983, final version received January 1984

This paper first updates to 1980 the estimates of the Harberger inflation model for Argentina, Brazil, Chile, Colombia and Uruguay and shows that all the hypotheses of this model hold and the reaction to monetary growth is fairly rapid. An aggregate supply function including imported inputs is then hypothesized, following recent theoretical work. This yields two testable propositions on the reduced form estimates of inflation: (a) the relation between inflation and monetary growth should be less than proportional and the income coefficient should be insignificant contrary to the Harberger model, and (b) inflation should be a weighted average of monetary growth and imported inflation, assuming the underlying aggregate production function is homogenous of degree zero in prices. Tests on the same data base support both hypotheses. These results also explain the similar results of many of the previous estimates of the Harberger inflation equation, which included the rate of devaluation in an ad hoc fashion.

1. Introduction

Latin America has been a fertile ground for empirical studies of inflation. The countries of the Southern Cone and Brazil provided practically the only evidence on the effects of sustained high rates of inflation and Colombia, Peru and Mexico yielded some inferences on the effects of medium rates. In recent years inflation has picked up in both sets of countries, so that the first group experienced some inflation rates of 100% or more per year, while in the medium range countries inflation has moved up to levels approximating those formerly prevailing in the Southern Cone. Econometric investigations of these inflations usually were based on the demand function for real money balances. Following Harberger, a long line of researchers regressed inflation rates on money, income growth, and changes in proxies for the cost of holding money balances.

The purpose of this paper is to systematically introduce an additional variable into the empirical explanations of Latin American inflation — the

*The author wishes to thank Anne Case for research assistance and George Borts, Jim DeMelo, Jorge Garcia, and participants in seminars at the IMF, International Finance Division of the Board of Governors of the Federal Reserve, and the III Latin American Econometric Society Meetings for helpful comments, but accepts responsibility for all remaining errors and opinions. In particular, the views and interpretations expressed in this article should not be attributed to the World Bank, its affiliated organizations, or to any individual acting in their behalf.

local cost of imports, which reflects both the rate of exchange and the dollar prices of imports. Since most of the countries in question have pursued policies of import substitution for long periods, their imports are mainly inputs, capital goods, and non-competitive food products. Based on this stylized fact the cost of imports is treated as an input price. By using an implicit cost function approach the aggregate supply function can then be specified in terms of final goods prices and local prices of imported inputs.

The macroeconomic importance of imported inputs has been recognized theoretically for some time so such an approach is not wholly new. [See, for example, Schmid (1976), Findlay and Rodriguez (1978), and Bruno (1979).] Moreover, some of the estimators of the Harberger model introduced the exchange rate into their regression equations in an ad hoc fashion. [See, for example, Behrman (1973), Diaz Alejandro (1965, 1970, 1976), Musalem (1971), Nugent and Glezakos (1979), and Sheehey (1976, 1980).] The novelty of this paper lies in using the properties of the implicit cost function to derive a reduced form equation for inflation which is comparable to Harberger's single equation monetarist model but which implies two very different testable hypotheses: (1) inflation is less than proportional to money growth, and (2) inflation can be written as weighted sum of money growth and inflation in import prices under the assumption that the underlying aggregate production function is homogenous of degree one.

Section 2 of the paper outlines the basic Harberger model and presents some estimates for the 1954 to 1980 period for Argentina, Brazil, Chile, Colombia and Uruguay. These results update the earlier estimates of the Harberger model, most of which end around 1970. In general, the results confirm the model and are similar to those obtained by previous investigators, except that inflation appears to adjust fairly rapidly to money growth. Section 3 briefly develops the model of the supply side and then tests the new model's hypothesis on a reduced form comparable to Harberger's equation in the same five countries. In general, this model performs much better than the basic Harberger model and the hypotheses are an acceptable approximation of reality. Section 4 presents a summary and some concluding remarks.

2. The Harberger model

The basic Harberger model explains inflation by assuming that prices adjust to absorb excess nominal balances. Real income is assumed to be exogenous and, following the simplest version of Cagan's approach, lagged inflation is taken as the measure of the cost of holding narrowly-defined money, i.e., either portfolio substitutability is between goods and money balances or interest rates on alternatives to money mainly reflect inflation. Interest rates typically are not used as a cost of holding money because of lack of data and because legal ceilings on interest rates distort the data.
For estimation the money demand function is transformed so that inflation \((DP)\) is written as a function of current and lagged values of money growth \((DM)\); income growth \((DY)\); the variation in the cost of holding money \((DC)\), often taken to be the change in lagged inflation \(D(DP,-1)\); and an error term with the usual properties \((U)\),

\[
DP_t = a + \sum_{i=0}^{m} b_i DM_{t-i} - \sum_{j=0}^{n} c_j DY_{t-j} + d D(DP_{t-1}) + U_t, \tag{1}
\]

where the \(D\)'s refer to differences in the logs. In this formulation, the constant would pick up any trend in velocity and any average inflation unexplained by the average values of the independent variables. One common test of the validity of the model is that the coefficients of money growth sum to one.

Harberger used inflation as the dependent variable to emphasize that it is prices that adjust to excess quantities of money;\(^1\) because inflation, rather than the price level, is the important variable in policy decisions; and because first differencing tends to reduce spurious correlation. Differencing also may resolve some statistical problems if errors are non-stationary, although this involves sacrificing some efficiency, as Plosser and Schwert (1978) have discussed. Finally, writing the equation in this form also reduces the impact of periods of price control, although these may still show up in an increased significance of the coefficients of lagged exogenous variables.

Variants of Harberger's model have been applied to sixteen Latin American countries by Vogel (1974), Betancourt (1976) and Nugent and Glezakos (1979); to Chile by Behrman (1973); to Argentine, Brazil and Chile by Wachter (1976) and Sheehy (1980); to Argentina by Diz (1970), Diaz Alejandro (1965, 1970), Sheehy (1976) and Mallon and Sourrouille (1975); to Colombia by Musalem (1971) and Diaz Alejandro (1976); and to Brazil, Colombia, Mexico, Peru and Venezuela by Rosas (1972). Table 1 presents estimates of a simple version of the Harberger model for Argentina, Brazil, Chile and Colombia, over the period 1954–1980 and for Uruguay over the period 1961–1980.\(^2\) These estimates extend by eight to ten years the coverage of the aforementioned studies, all of which terminate before 1973.

The results strongly confirm the Harberger model and in general are similar to those obtained by other authors. The fits, as measured by the \(R^2\),

\(^1\)This approach is somewhat different from the now customary procedure of estimating the real money demand function as a partial stock adjustment process. The Harberger approach emphasizes the role of price as the variable adjusting to clear the money market. There are two principal differences empirically, when the two models are set up comparably, with inflation rather than real money as the dependent variable: (1) the implied sign of lagged money growth — positive in the Harberger model, negative in the stock adjustment model, and (2) the use of \(D(DP_{t-1})\) in the Harberger model, \(DP_{t-1}\) in the stock adjustment model.

\(^2\)Reasonable data for Uruguayan exchange rates are not available before 1959. Since the main purpose of the estimates in table 1 is to provide a comparison for a more complex model involving exchange rates, the estimation for Uruguay was limited to the 1961–1980 period.
J.A. Hanson, Inflation and imported input prices

Table I

<table>
<thead>
<tr>
<th>Coefficient of:</th>
<th>Argentina b</th>
<th>Brazil</th>
<th>Chile c</th>
<th>Colombia</th>
<th>Uruguay d</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>0.121 §</td>
<td>0.082</td>
<td>−0.015</td>
<td>0.064</td>
<td>0.012</td>
</tr>
<tr>
<td></td>
<td>(0.051)</td>
<td>(0.081)</td>
<td>(0.083)</td>
<td>(0.045)</td>
<td>(0.062)</td>
</tr>
<tr>
<td>$M$</td>
<td>0.88 f</td>
<td>0.93 f</td>
<td>1.27 f</td>
<td>0.90 f</td>
<td>1.11 f</td>
</tr>
<tr>
<td></td>
<td>(0.08)</td>
<td>(0.16)</td>
<td>(0.11)</td>
<td>(0.20)</td>
<td>(0.15)</td>
</tr>
<tr>
<td>$DY$</td>
<td>−1.87 f</td>
<td>−1.31 f</td>
<td>−4.08 f</td>
<td>−2.00 f</td>
<td>−2.44 f</td>
</tr>
<tr>
<td></td>
<td>(0.60)</td>
<td>(0.62)</td>
<td>(1.00)</td>
<td>(0.79)</td>
<td>(0.78)</td>
</tr>
<tr>
<td>$DR$</td>
<td>n.s.</td>
<td>0.34 f</td>
<td>2.91 f</td>
<td>1.05 b</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.16)</td>
<td>(0.89)</td>
<td>(0.54)</td>
<td></td>
</tr>
<tr>
<td>SSR</td>
<td>0.4383</td>
<td>0.1971</td>
<td>0.7024</td>
<td>0.0768 2</td>
<td>0.1597</td>
</tr>
<tr>
<td>SEE</td>
<td>0.1380</td>
<td>0.0906</td>
<td>0.1787</td>
<td>0.0578</td>
<td>0.1000</td>
</tr>
<tr>
<td>$R^2$ (adjusted)</td>
<td>0.89</td>
<td>0.66</td>
<td>0.92</td>
<td>0.52</td>
<td>0.75</td>
</tr>
<tr>
<td>$DW$</td>
<td>1.75</td>
<td>1.73</td>
<td>1.57</td>
<td>1.93</td>
<td>2.63 i</td>
</tr>
<tr>
<td>SSR with lags</td>
<td>0.4321</td>
<td>0.1953</td>
<td>0.2659</td>
<td>0.05983</td>
<td>0.1534</td>
</tr>
<tr>
<td>(F statistic, d.f.)</td>
<td>(0.31, 2, 22)</td>
<td>(0.10, 2, 22)</td>
<td>(10.9, 3, 20)</td>
<td>(1.9, 3, 20)</td>
<td>(0.19, 3, 13)</td>
</tr>
</tbody>
</table>

* Standard errors of coefficients or F statistics in parentheses. SSR = sum of squared residuals, SEE = standard error of estimate; $DW$ = Durbin–Watson statistic; n.s. = not significant at 90% level, regression results exclude this variable; d.f. = degrees of freedom.

b Includes price control dummy, 1974 = −3, 1975 = 1, 1976 = 2, coefficient = 0.13(0.04).

c Includes price control dummy, 1972 = −1, 1973 = −3, 1974 = 4, coefficient = 0.18(0.04).


f Significant at 95% level.

g Significant at 99% level.

h Regression results exclude this variable. See text.

i Significant at 90% level.

j First order autocorrelation is not significant.

k $F$ statistic less than critical value — accept hypothesis.

The coefficients of money (defined as currency plus demand deposits) and income growth all have the correct sign and are significant at the 95% level. The Durbin–Watson statistics generally indicate a low probability of serial correlation, except in the case of Uruguay and there first order autocorrelation is not significant. The constant is not significantly different from zero at the 90% level, except for the case of Argentina. This result is an improvement over many of the previous studies, which often contained significantly positive constants, i.e., some inflation was unexplained or velocity was steadily declining. The smaller constant also might be explained either by more observations or by higher average rates of inflation, either of which might tilt the regression plane toward the origin.³

³ Nugent and Glezakos discuss the problems with the inclusion of a constant and apparently omit it in their empirical work. However, they do include a smoothed income growth variable, which varies little and thus performs empirically much like a constant.
or better. In all cases, except Chile, the coefficient of money growth is not significantly different from one at any reasonable level of significance. The coefficients of income are negative as hypothesized. They are also somewhat larger and more significant than those obtained in many of the previously mentioned studies, with the principal exception of Sheehey (1980).

Regressions using changes in lagged inflation to proxy the opportunity of holding money yielded insignificant coefficients for that variable. As an alternative, I tried both the changes in monetary growth \([D(DM_i)]\) and the changes in interest rates on easily available assets \((DIR)\). These proxies for the cost of holding money avoid the problems inherent in using some form of inflation — either the assumption of unreasonably long lags in adjustment of expectations if \(D(DP_{t-1})\) is used, or the statistical problem of running the dependent variable on the right hand side if \(D(DP_t)\) is used. Changes in money growth are exogenous to the same degree as money growth and can be thought of as a reduced form for the acceleration of inflation. The interest rate data present more problems. The data mainly consist of infrequently changed ceiling rates, with some freely set rates in periods when interest rates were liberalized. The result is that the interest rate variable looks somewhat like a dummy variable for the pre- and post-1974 period.

The results using these opportunity cost variables are mixed, but generally better than the results which the aforementioned studies obtained using the change in lagged inflation. The change in the rate of monetary growth was never significant. However, the interest rate variable was of the right sign and very significant in the cases of Colombia and Chile, less significant in the case of Uruguay. In the case of Argentina, the interest rate variable was insignificant and in the case of Brazil it was negative but significant at the 90% level. The latter result probably reflects the 1965–1974 period, when inflation slowed and interest rates rose because of financial reform. [See Syvrud (1974).] The reported results for Brazil exclude the interest rate variable.

One difference from the results of most of the earlier studies, except Sheehey (1980), is the apparent rapid reaction of inflation to monetary growth. In all cases, except Chile, the lagged independent variables were not jointly significant. This is shown in the last two lines of table 1 which reports the SSR obtained in a regression including the lagged variables, and the corresponding F statistic and its significance. This result may reflect the use

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4 This result casts some doubt on the usefulness of a stock adjustment model. See footnote 1. The low significance of the lagged inflation variable sometimes occurred in previous work, although Vogel (1976), Sheehey and Betancourt obtain reasonable results with lagged inflation and Nugent and Glezakos with a more complicated lag structure of past inflation. However, the significance of such cost variables seems to decline when the exchange rate is included and/or countries are grouped by average inflation rates.

5 Only in Colombia is one of the variables, lagged money growth, significant separately and this result disappears in the model described in section 3.
of annual rather than quarterly data. It also may reflect the data base because, as Sheehey (1980) points out, the significance of lagged money growth in Vogel's work may be explained by his use of IFS money data, which actually refers to end of period figures. This problem occurs in most studies using IFS data directly. Both Sheehey (1980) and this study use the annual average money stock.

As pointed out in table 1, the regressions for Argentina and Chile include a dummy for a period of price control and decontrol — 1974–1976 for Argentina and 1972–1974 for Chile, respectively.6 These variables imply a reduction in inflation during the price control period and a compensating rise in inflation during the decontrol period, an approach which also fits the data best. Both dummies were quite significant; their omission would increase the SSR about 50% in the case of Argentina, about 80% in the case of Chile. It is worth noting that the significance of the lagged independent variables is not affected greatly by the omission or inclusion of the dummies.

3. A model of inflation including costs of imported inputs

As mentioned above, the Harberger model assumes prices adjust to absorb excess money balances, with output exogenous. An alternative approach would be to assume that output is determined by a comparison of prices and costs. In such a framework the price level would be determined jointly by demand and supply factors. Within a completely closed economy, such a complication might not add much, depending on the relation between costs and prices — see the discussion of the role of wages in Harberger (1963) and in Vogel's (1976) reply to Shehhey (1976). However, open economies are another matter. In particular, the inflationary Latin American economies are highly dependent on imported inputs for their manufacturing sector. In fact, their imports are composed of such inputs, some capital goods, some non-competitive agricultural products, and only a few consumer goods, which vary with the ups and downs of protection. In such an economy costs and final prices could move divergently for some time.

One simple way to incorporate costs into an econometric model of inflation is to use the non-familiar implicit cost function, which permits writing the supply function in the form

\[ Y = F(P_i, P), \]  

where \( P_i \) = price of input \( i, i = 1, 2, \ldots, n \) and the signs below the variables indicate the signs of the partial derivatives.7 A strong, testable hypothesis is

6See De Pablo (1980) for Argentina and El Mercurio (1975) and De Vylder (1967) for Chile.

7Note that strictly speaking \( Y \) should be interpreted as gross output rather than income in (2). In most of the discussion which follows \( Y \) can be either gross output or value added and the regressions use value added.
that this supply function is homogeneous of degree zero in its arguments, i.e.,
a doubling of both input and output prices would leave real supply unaffected, i.e., a doubling of both input and output prices would leave real supply unaffected. This would occur, for example, if the underlying aggregate production function is homogeneous of degree one.\(^8\)

For purposes of estimation, it is assumed that this supply function can be approximated using logarithmic first differences, and that prices of the non-imported inputs are functions of final goods prices, and output. These assumptions yield a supply function of the form

\[
DY = s_0 + s_1 DP - s_2 DPI + U_3,
\]

where \(DP_I=\) change in the log of price of imported inputs (exchange rate multiplied by a dollar price index of imports) and \(s_1 - s_2 = 0\) if the cost function is homogeneous of degree zero. In this formulation \(s_0\) is a time trend which could proxy neutral changes in technology.

Substituting (3) into (1) and ignoring lags so \(b_0 = 1\), yields an estimating equation for the inflation rate which is similar to the Harberger model, except for the inclusion of \(DP_I\).

\[
(1 + s_1 c_0)DP = (\alpha - s_0 c_0) + DM + s_2 c_0 DPI + DDC + U_4.
\]

The principal hypotheses which result from this equation, aside from a significant coefficient for \(DP_I\), are: (a) the estimated coefficient of \(DM\) should be significantly less than one; (b) inflation should be a weighted average of the growth of money and the inflation in the price of imported inputs, providing the underlying aggregated supply function is homogeneous of degree zero in price and the prices of imported inputs, i.e., \(s_1 = s_2\).

The reader may question the neglect of imported consumer goods and exports in the foregoing analysis; essentially the economy seems to be treated as closed, aside from its use of imported inputs. In fact these omissions are not that serious. Consider first the case of imported consumer goods. Assume that imports are imperfect substitutes for locally produced goods and that consequently prices of imported consumer goods should be included as an element of the price index that owners of non-imported inputs use in

\(^8\)Findley and Rodriguez use an aggregate supply function in their theoretical model but do not raise the question of homogeneity. Among the empirical studies of the Harberger model only Nugent and Glezakos considered a supply equation. They hypothesize a Phillips curve relation, based on the difference between unexpected and expected inflation. However, this equation is only used to generate instrumental variables for a two stage estimate (TSLS) of a Harberger-type money demand function, in which the exchange rate is included in an ad hoc fashion and the only hypotheses relate to the signs of the coefficients. According to their results, using TSLS rather than OLS reduces the statistical significance of the income variable, particularly its cyclical component, but does not affect the other estimated coefficients very much.
calculating their real returns. These real returns, in turn, can be assumed to
determine the quantity of factors supplied. Under these assumptions it is easy
to show that the changes in the price index — defined as a geometric
average of the prices of locally produced and imported goods — are a
weighted average of the (log) changes in money and prices of imported
goods. Thus the two hypotheses are not affected substantially by the
complication of imported consumer goods.

Turning to exports, theoretically they are already included because their
real value is part of the definition of real income. Nonetheless an estimate of
changes in the local price of exports \((DP_x = \text{first difference of the sum of logs of the exchange rate and dollar export prices})\) was tried as an additional
explanatory factor in the regression equations. However, the estimated
additional contribution of this variable to explaining inflation may in fact be
small since: (1) the estimating equation already includes the growth of money
as an independent variable, a growth which is correlated with local prices of
exports; (2) rises in the local prices of exports and imports also are correlated
owing to the effects of devaluations and to the worldwide inflation of the 70s;
(3) variations in imports, international reserves and foreign borrowings can
offset the inflationary effects of variations in export prices in the local
economy; (4) variations in taxes on exports and quantitative restrictions on
exporting may affect local export prices but not the proxy variable \((DP_x)\)
which was used in the regression.

A model in which the price of imported inputs influences inflation certainly
is not new. Without actually specifying a theoretical model, some of the
previously mentioned estimates of the Harberger model also included the
exchange rate as an exogenous variable. The theoretical model proposed here
is a simple variant of the models used by Schmid, Findley and Rodriguez,
and Bruno. The novelty of this paper lies in using the implicit cost model to
specify just why and how the price of imported inputs enters into the
inflation rate and the testing of the homogeneity property, to which we now
turn.

Table 2 presents estimates of eq. (4) — an equation comparable to
Harberger’s estimates — and some tests of the strength of the model. The
regression statistics — \(SSR, SEE, R^2, DW\) — indicate that eq. (4) fits the
data reasonably well. A comparison of these statistics with those in table 1
shows that the expanded model fits much better than the basic Harberger
model.

In all five cases, the coefficients of money growth and the inflation in
import prices have the correct sign and are significant at the 95% level or
better. The coefficient of the import price variable is quite significant, the first
confirmation of the model’s hypotheses. This result is similar to that
obtained by Behrman, Sheehy (1976, 1980), Musalem, Nugent and
Glezakos, and Díaz Alejandro (1965, 1970, 1976), who all found a significant
positive relationship between devaluation and the rate of inflation. The constant of the regression is not significant at the 95% level, implying that declines in velocity and neutral changes in technology were roughly offsetting. The interest rate variable remains significant only in the case of Chile and Colombia. As might be expected, the coefficient of interest rates is somewhat smaller than that obtained in table 1. The estimated local price of exports ($DP_x$) also was tried but in all cases, except Chile, it was insignificant and not included in the reported results. Finally, the price control dummy remains significant in the cases of Argentina and Chile.

The lower panel of table 2 reports some tests of the robustness of the simple regressions, using F tests based on the SSRs obtained by (1) adding lagged values of the independent variables, and (2) splitting the sample at the midpoint (1967) and testing for the stability of the coefficients. The results of these tests show: (1) The lagged independent variables are not jointly significant in any of the cases. Moreover additional regressions (not shown) indicate that the lagged independent variables also are not significant separately. For all practical purposes the inflation rate seems to adjust within one year to the variations in monetary growth and inflation in prices of imports. (2) In all cases one cannot reject the hypothesis that the pre- and post-1967 regressions are the same.

Regarding the two specific tests of the model, the coefficient of money growth is significantly less than one in all cases. This would imply rejection of the Harberger model but confirms one of the strong hypotheses of the cost model. Second, constraining the coefficients of money growth and inflation in import prices to sum to one — the homogeneity hypothesis — does not affect the fit of the regression very much. As shown in table 2, just below the regression statistics, an F test of this constraint would not allow rejection of the homogeneity hypothesis at any reasonable level of significance, except for Chile which is discussed below. Thus, the homogeneity hypothesis would be acceptable within this sample in all cases except Chile, confirming the model. The coefficient of the reduced form of the 'real' exchange rate in the constrained equation is reported in the next line of table 2 and, as might be expected, is almost exactly the same as the coefficient of $DP_t$ in eq. (4) and significant at the 99% level. In general, the coefficients of import prices are surprisingly large — much larger than would be expected given their weight in the wholesale prices index and the low ratio of foreign trade to GDP in these countries.9

Another comparison between the cost model and the typical Harberger, exchange-rate-augmented model can be made by re-running eq. (4) with an income variable, a form equivalent to eq. (1) with an import price variable

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9The coefficients for Colombia are the only ones which are close to the weights. Under the 1970 revision of the Colombian wholesale price index imports had a 10% weight and exports an 11% weight.
Table 2

<table>
<thead>
<tr>
<th>Coefficient of:</th>
<th>Country</th>
<th>Argentina</th>
<th>Brazil</th>
<th>Chile</th>
<th>Colombia</th>
<th>Uruguay</th>
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</thead>
<tbody>
<tr>
<td>Constant</td>
<td></td>
<td>-0.005</td>
<td>-0.027</td>
<td>-0.057</td>
<td>-0.018</td>
<td>0.076</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.034)</td>
<td>(0.043)</td>
<td>(0.028)</td>
<td>(0.030)</td>
<td>(0.055)</td>
</tr>
<tr>
<td>$DM$</td>
<td></td>
<td>0.58f</td>
<td>0.58f</td>
<td>0.26f</td>
<td>0.61f</td>
<td>0.44f</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.09)</td>
<td>(0.16)</td>
<td>(0.09)</td>
<td>(0.17)</td>
<td>(0.17)</td>
</tr>
<tr>
<td>$DP_I$</td>
<td></td>
<td>0.41f</td>
<td>0.38f</td>
<td>0.51f</td>
<td>0.25f</td>
<td>0.42f</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(0.08)</td>
<td>(0.08)</td>
<td>(0.09)</td>
<td>(0.06)</td>
<td>(0.10)</td>
</tr>
<tr>
<td>$DIR$</td>
<td></td>
<td>n.s.</td>
<td>n.s.</td>
<td>0.16f</td>
<td>1.94f</td>
<td>n.s.</td>
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<td></td>
<td>(0.07)</td>
<td>(0.79)</td>
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**Regression statistics**

<table>
<thead>
<tr>
<th>Statistic</th>
<th>Argentina</th>
<th>Brazil</th>
<th>Chile</th>
<th>Colombia</th>
<th>Uruguay</th>
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</thead>
<tbody>
<tr>
<td>SSR</td>
<td>0.2696</td>
<td>0.1260</td>
<td>0.1059</td>
<td>0.0589</td>
<td>0.1333</td>
</tr>
<tr>
<td>SEE</td>
<td>0.1083</td>
<td>0.0725</td>
<td>0.0710</td>
<td>0.0506</td>
<td>0.0886</td>
</tr>
<tr>
<td>$R^2$ (adjusted)</td>
<td>0.93</td>
<td>0.78</td>
<td>0.99</td>
<td>0.63</td>
<td>0.80</td>
</tr>
<tr>
<td>DW</td>
<td>1.46</td>
<td>2.01</td>
<td>1.92</td>
<td>1.76</td>
<td>2.45</td>
</tr>
</tbody>
</table>
### Statistics for tests of the model

<table>
<thead>
<tr>
<th>Test Description</th>
<th>SSR</th>
<th>F Statistics, d.f.</th>
<th>Coefficient of ( DP_1 - DM ) in the constrained regression</th>
<th>SSR with lags</th>
<th>F Statistic, d.f.</th>
<th>Sum of SSRs splitting samples at mid-point (1967)</th>
<th>Coefficient of ( DY ) and t statistic</th>
</tr>
</thead>
<tbody>
<tr>
<td>SSR forcing homogeneity</td>
<td>0.2696</td>
<td>(0.0, 1.24)</td>
<td>0.41</td>
<td>0.2293</td>
<td>(1.93, 2.22)</td>
<td>0.2661</td>
<td>-0.76</td>
</tr>
<tr>
<td>(F statistics, d.f.)</td>
<td>0.1266</td>
<td>(0.01, 1.24)</td>
<td>0.38</td>
<td>0.1087</td>
<td>(1.75, 2.22)</td>
<td>0.1194</td>
<td>-1.19-1.</td>
</tr>
<tr>
<td></td>
<td>0.1273</td>
<td>(4.44, 1.22)</td>
<td>0.63</td>
<td>0.0671</td>
<td>(2.45, 4.17)</td>
<td>0.0682</td>
<td>-0.38</td>
</tr>
<tr>
<td></td>
<td>0.0606</td>
<td>(0.67, 1.23)</td>
<td>0.26</td>
<td>0.0476</td>
<td>(1.66, 3.31)</td>
<td>0.0455</td>
<td>-0.61</td>
</tr>
<tr>
<td></td>
<td>0.1422</td>
<td>(1.23, 1.18)</td>
<td>0.40</td>
<td>0.1151</td>
<td>(1.19, 2.15)</td>
<td>0.1237</td>
<td></td>
</tr>
</tbody>
</table>

*aSee table 1 for abbreviations.*

*Includes dummy, 1974 = -3, 1975 = 1, 1976 = 2, with coefficient = 0.12 (0.03). See text.*

*Includes dummy, 1972 = -1, 1973 = -3, 1974 = 4, with coefficient 0.044 (0.017), and local export prices with coefficient = 0.37 (0.010). See text.*


*Significant at 90% level.*

*Significant at 99% level.*

*Significant at 95% level.*

*F statistic less than critical value — accept hypothesis.*

*Insignificant when a dummy is included for 1967–1974 or the exchange rate is used as \( P_r \). See text.*
added. This new regression produces the interesting result that the income variable generally drops to statistical insignificance, as reported in the last two lines of table 2.\textsuperscript{10} This result seems to indicate that in this sample income is far from exogenous, as the simple monetarist model tacitly assumes. Instead the effect of income on inflation can be adequately proxied by contemporaneous variations in import prices (and money), as implied by the cost model. To put it another way, the exchange-rate augmented, Harberger equation appears to be a reduced form rather than a money demand function.

The only ambiguous result occurs in the case of Chile, where strict homogeneity is rejected, although the sum of the coefficients does equal one if the coefficient of export prices is included. One explanation for the problem with the Chilean case may be that the Chilean price index includes only home and import goods prices and thus is not strictly the same as the other price indices. Moreover, Chile is the only country which experienced sharp decreases in protection along with its massive devaluation, changes which are not included in $P_I$. These differences may create a downward bias in the estimated effect of a devaluation.

A second explanation relates to the sensitivity of the Chilean results to the period chosen and to the inclusion of the price control dummy and the proxy for the change in the local price of exports. These two variables, and the interest rate variable, may simply be proxying the major economic changes between the last two years of the Allende administration and the Pinochet regime, which included a shift from price controls and rationing to market prices, a large increase in public enterprise prices, a collapse in the terms of trade, a trade liberalization and a significant fall in income. By comparison the coefficients of $DM$ and $DP_I$ for the shorter period 1954–1970 are both somewhat larger than those obtained for the 1954–1980 period, as shown in table 3, and the local export price and interest rate variables are not significant.\textsuperscript{11} While the hypothesis of homogeneity is rejected in the augmented regression shown in table 2, it would not be rejected for the period 1954–1970, as shown in table 3, nor can it be rejected in a regression involving only $DM$ and $DP_I$ for the period 1954–1980 (not shown). These additional results suggest that the hypothesis of homogeneity is not bad even for Chile, although strictly speaking, the results of table 2 imply rejection of the hypothesis for the 1954–1980 period at the 95\% level.

\textsuperscript{10}The coefficient of income is significant only for Brazil and it becomes insignificant either if a dummy is added for the 1967–1974 period of financial liberalization or if an exchange rate variable replaces the import price variable. In earlier studies the inclusion of the exchange rate always reduced the significance of the income coefficient. See for example, Musalem (1971) and Sheehey (1980). Nugent and Glezakos find that income generally is not a significant explanatory factor in the Harberger equation, when the exchange rate is included and countries are grouped by average inflation levels.

\textsuperscript{11}Although the hypothesis that the pre- and post-1967 regressions are the same cannot be rejected. See table 2.
Table 3
Alternative explanations of Chilean inflation, 1954–1970. a

<table>
<thead>
<tr>
<th>Coefficient of:</th>
<th>Regression comparable to table:</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>0.052 (0.103)  -0.040 (0.043)</td>
</tr>
<tr>
<td>DM</td>
<td>1.07b (0.25) - 0.38a (0.16)</td>
</tr>
<tr>
<td>DP,</td>
<td>-0.69b (0.09)</td>
</tr>
<tr>
<td>DY</td>
<td>-3.40c (1.27)</td>
</tr>
<tr>
<td>SSR</td>
<td>0.1138 0.01138</td>
</tr>
<tr>
<td>SEE</td>
<td>0.0902 0.0462</td>
</tr>
<tr>
<td>R² (adjusted)</td>
<td>0.66 0.91</td>
</tr>
<tr>
<td>DW</td>
<td>1.43 1.94</td>
</tr>
<tr>
<td>SSR with lags</td>
<td>0.1136 0.0218</td>
</tr>
<tr>
<td>(F statistics, d.f.)</td>
<td>(0.01, 2,12) (2.26, 2,12)</td>
</tr>
<tr>
<td>SSR forcing homogeneity</td>
<td>-0.0365</td>
</tr>
<tr>
<td>(F statistics, d.f.)</td>
<td>(3.03, 1,14)</td>
</tr>
<tr>
<td>Coefficient of DP, -DM in the constrained regression</td>
<td>0.70b (0.08)</td>
</tr>
<tr>
<td>Coefficient of DY and statistic</td>
<td>-1.04 (0.72)</td>
</tr>
</tbody>
</table>

aOmits DIR and DPx which were insignificant. See table 1 for abbreviations.
bSignificant at 99% level.
cSignificant at 95% level.
dF statistic less than critical value — accept hypothesis.

4. Summary and conclusions

The principal result of this study is that in the period 1954–1980 changes in local costs of imports played an important role in the determination of inflation in five Latin American inflationary economies. As a standard of comparison for demonstrating this result, the paper first presents some estimates for the 1954–1980 period of the well-known Harberger model of inflation, a model that is based on the demand function for money and that assumes output is exogenous. These estimates yield broadly similar results to existing studies, all of which end around 1970 or earlier. In particular, the fits are reasonably good and the coefficient of monetary growth is not significantly different from one, as the Harberger model predicts. The principal difference from previous work, except Sheehey (1980), is the greater rapidity with which inflation apparently adjusts to monetary growth.
The paper then assumes an aggregate production function which, using the well-known implicit cost function analysis, permits supply to be written as a function of the prices of final output relative to the costs of imported inputs. A simple version of such an aggregate supply function is combined with a money demand function to yield a reduced form comparable to Harberger's single equation with one exception — inflation is a function of both monetary growth and changes in the prices of imported inputs. Thus one test of the cost model is simply the significance of the price of imported inputs. However, two stronger, testable propositions emerge from the addition of the cost function: (a) the empirical relationship between inflation and monetary growth is no longer one of proportionality, and (b) inflation is a weighted average of monetary growth and changes in the price of imported inputs if the underlying production function is homogeneous of degree one in local and imported inputs.

Regressions of the annual inflation rate on growth in money (currency plus demand deposits) and a proxy for inflation in the price of imported inputs (the sum of the log of the exchange rate and import prices in dollars) in Argentina, Brazil, Chile, Colombia and Uruguay show that the second variable contributes significantly to the explanation of inflation, as predicted by the supply augmented model.\(^2\) The new regressions fit about twice as well (as measured by the SSR) as those based on the simple model and the coefficients are stable when the sample is split in half. The addition of lagged independent variables yields no significant improvement in the fit; for all practical purposes the adjustment of inflation to monetary growth or changes in the local prices of imports takes place within one year. One surprising result is the size of the coefficients for the import prices. They generally are much larger than the share of imports in the wholesale price indices. This suggests that prices of numerous goods, including government-set prices and labor costs, seem to move closely with the cost of imported inputs. Finally, the coefficient of monetary growth becomes significantly less than one in all cases. The homogeneity hypothesis — that inflation can be explained by a weighted sum of monetary growth and inflation in the price of imported inputs — is very acceptable in four of the five cases, tolerably acceptable in the Chilean case.

The idea that costs of imported inputs are an important macroeconomic variable is not a new one; for example, Schmid, Findlay and Rodriguez, and, more recently, Bruno have all developed theoretical models emphasizing the importance of this variable. Some of the previous estimators of the Harberger inflation model included the exchange rate on an ad hoc basis. The contribution of this paper lies in the development of an implicit cost model

\(^2\)Similar results are obtained using only the rate of devaluation, but the fits are somewhat worse when the dollar price of imports is not included.
which explains how these costs should enter, i.e., it yields strong testable hypotheses. These hypotheses in general are confirmed by the data. In addition, it is worth noting that in much of the existing work using the Harberger model, which includes the rate of devaluation as an ad hoc explanatory variable, the standard errors of the coefficients suggest that the hypothesis of homogeneity could not be rejected.\(^3\)

The size and significance of the estimated coefficient of the local price of imports suggests that future studies of inflation should pay close attention to this variable, particularly if the recent worldwide inflation continues into the 80s. For countries which are more open or have engaged in less import substituting industrialization and are correspondingly less dependent on imported inputs than those considered here, the inflationary impact may be less severe, but nonetheless important.

The empirical results of this study also shed some light on the question of whether exchange rate movements affect relative prices. The regression coefficient of import prices, while surprisingly large, generally is significantly less than one, a result supported by Connally and Taylor (1976) using different methodology. This indicates that, in the short run, 'slow' devaluation may repress inflation for a given rate of monetary growth by allowing 'cheap' imports. However, once a devaluation is forced upon a country for balance of payments reasons, inflation will exceed monetary growth and the repression will be corrected.

Two open issues are whether this relative price effect can be sustained in the longer run; and if it cannot, then by what mechanisms are relative prices adjusted? One of the hypothesized mechanisms that does not seem to work in this sample is a lagged adjustment to devaluation through a lagged rise in money growth and inflation. Some experiments with Sims tests of the direction of causation indicate a basically contemporaneous relationship between changes in local prices of imports and money growth.\(^4\) However, there does seem to be enough independent variation between money growth and the local prices of imports to permit separation of their inflationary effects — the estimated coefficients are three to five times their standard errors, indicating multi-collinearity is not a problem.

Finally, this study, like the original Harberger study, does not explain why the policies under consideration — money growth and devaluation — occur. Perhaps, as Connally and Taylor argue, it is inflation that causes devalu-

\(^3\)See Diaz Alejandro (1970), Musalem, Nugent and Glezakos, and Sheehey (1976, 1980). In Diaz Alejandro (1965) the sum of the coefficients exceeds one but this result may reflect the peculiarities of the 1945–1962 period or the use of 1962 as the end point. Nugent and Glezakos also explicitly test for the impact of current and lagged monetary growth on inflation and reject the hypothesis that the coefficients sum to one when the exchange rate is included.

\(^4\)Sims regressions involved leads and lags of two years and were run on untransformed data from Argentina, Brazil, Chile and Colombia. The Uruguayan series was too short to warrant this analysis.
ation, rather than the reverse. This may be a point worth considering by future investigators, especially in view of the quasi-indexed systems which were used in the five countries, but it remains beyond the scope of this study.

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