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Brazilian Economic Studies

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Brazilian Economic Studies

IPEA/INPES. Rio de Janeiro, Brasil. 1984.

ISSN 0100 — 2910

Brazilian economic studies. — Rio de Janeiro (Caixa Postal 2.672) : Instituto de Planejamento Econômico e Social, Instituto de Pesquisas, 1975 (n.º 1)-23 cm.

Brazilian Economic Studies is published by IPEA under the responsibility of its research institute (INPES). Opinions expressed in this publication are those of the authors and do not reflect the views of the Institute.

Inflation and anti-inflationary policies in Brazil *

Mario Henrique Simonsen **

1 — Historical background

Inflation has already been recognized as part of the Brazilian way of life. Prices have been rising uninterruptely since 1933, which is to say that people under fifty or even perhaps sixty of age have not experienced life with stable prices. Meanwhile real output has been expanding at significantly fast rates, a trend of 7% a year having been sustained since the early fifties. There is no evidence that the Brazilian lack of price discipline contributed to this long term growth. Yet, the Brazilian case dismisses the old parable that chronic inflation is not consistent with sustained economic development.

Reliable series of price indices extend back only to 1939. Among them, the most popular indicator of inflation is the general price index estimated by the Getulio Vargas Foundation. It is a curious blend of a national wholesale price index (weight 6) with two regional indices for the city of Rio de Janeiro, the cost of living index (weight 3) and the construction cost index (weight 1) Tradition is the only explanation why this weighted average is still labeled the general price index. The Government has recently tried to divert attention from this indicator to the national consumer price index of the IBGE. (National Institute of Geography and Statistics). This campaign seems to be inspired on a pragmatic motivation. In the twelve month period ending in october 1980, the inflation rate, which stands at 87% according to the national

* Paper presented to the Conference on World Inflation and Inflation in Brazil, Rio de Janeiro, December 15-16, 1980. The author is indebted to Rudiger Dornbusch for his comments on this paper.

** Fundação Getulio Vargas, Rio de Janeiro.

consumer price index, leaps to 109% when measured by the general price index, the difference being explained by the explosive behavior of the wholesale price index in the last twelve months. From the technical point of view, the national CPI is probably less of a mixed bag than the general price index. Yet, since there are no reports on the national CPI prior to 1979, we shall stick throughout this paper to the traditional measurement of inflation by the changes in the general price index of Getulio Vargas Foundation. [1]

Information on inflation rates prior to 1939 is at best imprecise. With some tolerance for statistical methods one might use a cost of living series for the city of Rio de Janeiro which extends back to 1912. For the period before 1912 we are led into the cloudy field of price archeology. An interesting book in this field is Mircea Buescu's "300 years of inflation", where it is shown that prolonged general price rises were not uncommon since the middle colonial times. In another curious book published in 1960, Oliver Onody produced a cost of living series tor the period 1829/1912. Onody's indices are not to be taken too seriously, since they are based on simple arithmetical averages of some selected prices. Yet, they do provide some indication on the general price movements during the nineteenth century.

According to Onody's indices, throughout the years of the Empire (1822/89) Brazil experienced a very moderate but steady general price rise, at an average annual rate of 1,5%. By modern standards this could be praised as a period of admirable price stability, but one should remember that in the nineteenth century, since prices were flexible enough, waves of deflation were almost as frequent as those of inflation. Price cycles, so frequently found in the economic records of other countries, were virtually unknown under the Brazilian Empire.

The early days of the Republic were accompanied by huge budget deficits financed by money printing. Economic policies were then oriented by Ruy Barbosa, who was a remarkable Brazilian writer and thinker but also a highly expansionary Minister of Finance. The result was an inflationary burst, prices more than doubling over a five-year period.

1 Implicit deflators, in spite of their theoretical advantages, will not be taken into account, since they accumulate two groups of measurement errors which should not be neglected: those in the nominal and those in the real output. Estimates of the implicit deflator of the gross domestic product, published by the Getulio Vargas Foundation, have in fact been subject to frequent revisions.

President Campos Salles (1898/1902) and his Minister of Finance Joaquim Murtinho decided to implement the most severe antiinflationary program ever experienced by the country. Public expenditures were dramatically cut, the money supply was contracted and prices actually fell. Virulent side-effccts have been reported in terms of noisy bankrupcies which probably caused a substantial decline in the levels of output and employment. As a political cost, Campos Salles ended his term as the most unpopular President of the Old Republic. It was recognized only later that he paved the way to the first Brazilian miracle, a twelve year period of prosperity with perfectly stable prices, from 1902 through 1914.

Inflation resumed its pace with the beginning of World War I, now at an average annual rate of 8%, from 1914 to 1927. A brief period of price stability was achieved during the Washington Luiz administration (1926/30). Prices then plunged some 15% between 1929 and 1933, as a result of the Great Depression.

The Old Republic was buried by the 1930 Revolution which led Getulio Vargas into power, first as a provisional President, then as a Constitutional one, and later on as a Dictator. Substantial economic changes were then introduced. The prices of coffee, the main export item of the country, had collapsed with the Great Depression and, as a result, Brazil had to face unprecedented trade deficits. Confidence in the efficiency of the market mechanisms was disrupted, and the Government tried to solve the problems by imposing a number of regulations. Import substitution was fostered by a new wave of protectionism. Institutes were created to control the production and exports of some key items, such as coffee and sugar. Labor laws were enacted. In terms of economic growth, the balance of such policies does not appear to have been unfavorable. According to Isaac Kerstenetsky's estimates, Brazil was able to sustain an average rate of growth of real GDP of 4% a year along the thirties, a remarkable record for that decade. Inflation, however, picked up again in 1934, at an average annual rate of 7%.

World War II imposed a shortage of Brazilian imports and favored the country's exports. Monetary and fiscal policies were unable to offset the expansionary effects of the balance of payments surpluses. Both, the increase in the aggregate demand and some import bottleneeks pushed the inflation rate of 15% a year.

In October 1945 Vargas was deposed. The first constitutional government of the new republic, under President Eurico Dutra (1946/50) was temporarily successful in bringing the inflation rates down through the elimination of the import restraints. Since the cruzeiro was clearly overlued for the postwar economy, the cost of such policies was the depletion of the foreign reserves accumulated during the war period.

From 1951 to 1958 the average annual rate of inflation jumped to 17%, with a minimum of 11% in 1952 and a maximum of 27% in 1954. The acceleration of the price increases can be ascribed to three main causes: i) the expansion of the money supply, at an annual average rate of 22%; ii) the need to realign the exchange rate, which was pegged to the Bretton Woods parity until 1953; iii) the minimum wage increases of 1952, 1954, 1956. Minimum wage increases exerted substantial pressures on prices in those days, since the average salaries were then quite close to the minimum official levels.

A new period of soaring rates of inflation was to begin in 1959, in the second half of the Kubitschek administration. The Government had launched an ambitious program of import substitution, of highway construction, of expansion of the hidroelectric capacity, and, as a very especial symbol, of construction of Brasilia, which was to become the new capital of the country. Economic growth was to be achieved at any cost, according to the official ideology, and monetary orthodoxy was duly rejected. In 1961, President Janio Quadros announced a well articulated economic program, the first phase of which was a strong devaluation of the cruzeiro, which had been once again overvalued. Nobody will ever know how the subsequent phases would have evolved, since Quadros resigned after seven months. He was succeeded by his populist Vice-President João Goulart who encouraged labor strikes, generously expanding both, nominal wages and budgetary deficits, under the heading of income redistribution. General price level increases, measured from December to December, soared to 39.2% in 1959, 30.5% in 1960, 47.7% in 1961, 51.3% in 1962, 81.9% in 1963 and 91.9% in 1964.

In March 31, 1964, Goulart was deposed. The new military government, under President Castello Branco (1964/67) implemented a comprehensive economic program not only designed to fight inflation and to adjust the balance of payments, but especially to lay the basis for a future period of accelerated growth. The first phase of the program consisted of a quick realignment of relative prices, including the exchange rates, which had been strongly distorted in the late fifties and early sixties. This was the so-called corrective inflation period, in which the Government had to bear the cost of inheriting an explosive and still repressed general price rise.

After the 1964 peak of 91.9%, the inflation rate plunged to 34.5% in 1965. Excellent crops might have favored this result (the agricultural production increased by 13.8% in 1965), but there is little doubt that inflation basically declined because of the reduction in aggregate demand. Evidence is to be found in the fact that the only prolonged industrial recession in Brazil, since World War II, occurred in 1965, when industrial output fell by 4.7% on a yearly average. Yet, there seems to be no room to explain the 1965 events on pure monetary grounds. The money supply had expanded by 81.6% in 1964 and continue to expand at 79.5% in 1965. In fact, the monetary expansion of 1965 was, in some sense, much healthier than that of the previous year. It was much less directed to the expansion of domestic credit, and much more to the purchase of foreign assets and agricultural surpluses. Yet, these sectoral considerations did not imply a reduction of the real cash balances held by the public. Aggregate demand and inflation rates fell in 1965 mainly because of the substantial budget cuts, and because of a newly implemented wage formula which will be described in Section 3. These facts were strong enough to change the inflationary expectations of most people, inducing them to accept a substantial increase in the real stock of money.

Anti-inflationary policies had no visible effects in 1966, when the general price index rose by 38.8% and when industrial production recovered by 9.9%. Yet, for the first time the newly created Central Bank set forth a truly tight monetary policy (by Brazilian standards, of cource), limiting to 13.8% the expansion of M_1. The liquidity squeeze again pushed down aggregate demand and industrial output in late 1966 and early 1967. As a result, the inflation rate fell to 24.3% in 1967.

Castello Branco's term ended on March 15, 1967. His period marks the most dramatic anti-inflationary achievement in the country since the Campos Salles administration, and much of the impressive results are to be attributed to Finance and Planning Ministers Octavio Bulhões and Roberto Campos. In fact a much more ambitious target was being pursued in those days, that of bringing the inflation rates down to 10% a year. This objective was never to be achieved. Yet the importance of the 1964/67 policies was not confined to the reduction of the inflation rates. The balance of payments was also properly adjusted. And a number of institutional reforms paved the way to a new era of accelerated growth.

A widespread system of indexation rules was the most interesting outcome of the above mentioned reforms. Problems created by overabundant escalator clauses will be discussed in

Section 7, where it will be suggested that widespread indexing might introduce inflation rigidities, with especially perverse effects in periods of adverse supply shocks. Life with inflation, however, becomes much easier. Until 1964 the Brazilian economy was subject to all the classical inflationary distortions listed in elementary textbooks, in fact those created either by unanticipated inflation or by money illusioned regulations, such as usury laws, rent freezes, etc. Since the middle sixties most of these distortions were duly eliminated by indexing. (1980 is to be regarded as an unfortunate exception, since the indexation rules were broken).

Since the average real wage in the manufacturing industry declined 24.8% from 1964 to 1967, it is often argued that the anti-inflationary policies of the Castello Branco Government could only be successful to the extent that they squeezed the wage earners. The fact the labor unions were weakened and that nominal wages were just partially indexed, according to a formula which will be described later on, obviously contributed to the effectiveness of the anti-inflationary policies. Yet, the decline of real wages is to be ascribed to other objectives of economic policy, those of realigning the relative prices, of reinforcing the domestic savings rate, and of adjusting the balance of payments. It is worth mentioning that approximately half of the 24.8% real wage decline in the manufacturing industry was due to the price increase of the cost of living basket in terms of industrial goods units, mainly the result of rent defreezes and of subsidy cuts. The other half was a procyclical wage decline, which finds its natural explanation in the more competitive conditions of labor supply.

The years 1968 through 1973 were the fortunate period of the so-called "Brazilian Miracle". Annual inflation rates gradually declined from 24.3% in 1967 to 15.7% in 1973. Real domestic product expanded at an average rate of 11.5% a year. Money expanded at an average annual rate of 36.8%. Real wages experienced a substantial growth, especially those of skilled workers. Foreign capital flowed into the country, which was able to show a 6.4 billion dollar reserve figure by the end of 1973.

The natural unemployment rate hypothesis provides a simple explanation of the miracle. The restrictive policies of the Castello Branco period had brought the actual industrial output substantially below its potencial level, presumably some 15% in 1967, leaving a large space to accelerated growth in the subsequent years. Moderately expansive monetary policies could then stimulate growth without speeding up the inflation rate. Moreover, the period was blessed by favorable supply shocks, both from agriculture and from external prices.

Strong inflationary pressures were built up in 1973, when M_1 expanded 47%, when output advanced far beyond its trend, and when OPEC quadrupled the oil prices. Yet, the Medici Government was strongly commited to a 12% a year inflation rate ceiling. With massive price controls and subsidies the general price rise increase was repressed at 15.7%.

President Geisel took office in March 15, 1974, inheriting repressed prices, a number of commodity shortages and a huge deficit in the trade balance. In his five year period a compromise solution was attempted for a number of objectives, namely: i) to adjust the balance of payments and to keep a good international creditstanding; ii) to keep real product growing at its historical pace; iii) to control the inflation rate within acceptable limits by Brazilian standards; iv) to promote export growth and to reduce the foreign dependence of the country through a new program of import substitution. Times were difficult enough and objectives to reconcile also conflicting enough.

The balance of payments and import substitution objectives were satisfactorily managed, the trade deficit, in nominal dollars, being reduced from 4.7 billion in 1974 to approximately one billion in 1978, and foreign reserves being raised to 12 billion dollars at the end of the period. In spite of annual swings, the average rate of growth of real GDP was kept at 7%, the historical trend rate.

Anti-inflationary policies were subject to varying degrees of priority, and inflation rates thus became extremely volatile. A price outburst, due to the previously repressed inflation, occurred in the first few weeks of the Geisel Government. Then, the expansion of money was restricted to 23% in the first twelve months of the new Government. Tight monetary policies limited the general price increase to 23% from may 1974 to may 1975. They also abated the industrial rate of growth. Emphasis was then shifted to growth targets. Money growth thus increased to 42.8% in 1975 and an ambitious program of public investment was implemented. Inflation soared again, reaching the 46.3% rate in 1976. Tight monetary and fiscal policies were then reestablished. Yet, inflation now appeared more rigid, perhaps because of adverse supply shocks, perhaps because of full wage indexing which had been implemented in January 1975. The general price level increased by 38.7% in 1977 and by 40,8% in 1978.

President Figueiredo, who took office in March, 15, 1979. formally announced that his main economic objective would be to reduce inflation to the pre-oil crisis rates. Yet, in August 1979 a number of unorthodox policies were introduced, under the principle that inflation was to be fought through accelerated growth. Administered prices were quickly adjusted, since policy-makers believed that a high rate of inflation in 1979 would automatically be followed by a low rate in 1980. This was called corrective inflation. The money supply expanded rapidly, especially to provide abundant and cheap credit to agriculture, since it was also announced that a super-crop would bring inflation to a halt. Interest rates were controlled, since policy makers preferred Tooke to Wicksell. Wages became indexed on a six-month basis, free negotiation of a real increase, labeled productivity gain, being superimposed on the automatic nominal adjustment. Moreover, a 10% real bonus on a six-month basis was granted on all salaries up to three minimum wages. In december, 7, 1979, the cruzeiro was maxidevalued, the dollar/cruzeiro rate being increased 30%. Oddly enough, the main impact of the maxidevaluation was just to break an eleven year tradition of crawling pegs. Effects on imports and exports were largely offset by the sudden withdrawal of the export subsidies and import prior deposits, which previously were being phased out under a gradual five year program agreed with GATT.

In january 1980 the Government decided that a strong psychological move would erase most of the recurring effects of the late 1979 moves, including the external oil price shocks. Massive price controls and subsidies were then implemented, and the Government announced that monetary correction and exchange devaluations in 1980 would be limited to 45% and to 40% respectively. These ceilings were lately increased by some ten points of percentage.

Inflation rates skyrocketed.

2 — The monetary system

A chronic inflation at the Brazilian rates could never be sustained without highly expansive monetary policies, a view which seems to be shared even by most non-monetarists. In fact, in the thirty year period from december 1949 to december 1979, M_1 increased almost fourteen thousand times. The Brazilian monetary system is a very peculiar one, and with a built-in bias toward the expansion of the money supply. It has been long recognized that inflation is hard to prevent whenever the Federal Government has the authority to print money to finance its deficits, especially if this authority can be exercised independently of Congressional approval. This is the reason why a number of countries try to keep their Central Banks within some sort of apolitical shrine.

Inflation is still harder to prevent if the Government can create money not only to finance its deficits, but also to extend subsidized loans to the private sector. This is fundamentally the Brazilian case.

A unique role in the system is played by the "Banco do Brasil", a commercial bank controlled by the Federal Government, but with a minority fraction of its equity capital largely dispersed among private shareholders. At least in part, Banco do Brasil is a profit oriented financial institution, and its shares have been rated for long as stock-exchange blue-chips.

Until 1964 Brazil had no Central Bank. Monetary regulations were issued by SUMOC (Currency and Credit Superintendency). Legal money was issued by the Treasury whenever requested by the Banco do Brasil. And the Banco do Brasil acted simultaneously as a commercial bank, as a development bank, as the Government's bank, also providing rediscount facilities to the commercial banks. Since the board of the Banco do Brasil was appointed by the President of the Republic, this was just a roundabout method by which the Government could create money to finance its budget dificit (federal bonds were largely discredited until 1964), to extend commercial bank credit, to accumulate foreign reserves (a rare event in the years 1947 through 1964) and to expand Banco do Brasil loans to the private sector. The latter form of money creation increased both the profits and the political prestige of the Banco do Brasil, so that the official bank always kept strong vested interests in the expansion of the money supply.

The Central Bank was created in 1965. Legal currency issues, Federal Government financing, the provision of rediscount facilities to commercial banks and the control of foreign reserves were all ascribed to the Central Bank. Yet, the 1965 monetary reform was largely incomplete. Far from standing as an independent institution, the Central Bank was put under the control of the Ministry of Finance. (The same happened, incidentally, with the Banco do Brasil). And the Banco do Brasil still kept an unique operational system which will be described below.

Money and foreign exchange regulations are issued by the National Monetary Council, which is composed of five Ministers of State, eight Chairmen of federal financial institutions, including the Central Bank and the Banco do Brasil and by eight private experts. The Council is chaired by the Minister of Finance. Ceilings on the annual rates of growth of the monetary base, of the money supply and of the assets of Banco do Brasil are set by the Monetary Council, in an approved Monetary Budget.

Besides usual commercial bank liabilities, a very peculiar system of funding has been provided to Banco do Brasil: an open-end rediscount facility at the Central Bank, with a symbolic nominal interest rate of 1% a year. Thus the Banco do Brasil is "de facto" a second Central Bank. It can indirectly print money to finance the expansion of its assets. In theory, the Banco do Brasil can only use this rediscount facility to the extent that its assets do not exceed the monetary budget ceilings. For a period of time, when due respect was paid to monetary policy, excesses were penalized with heavy rediscount rates, which strongly deterred the expansionary tendencies of the Banco do Brasil. This rediscount rate penalties were abolished in August 1979.

A strong and determined Minister of Finance can manage the monetary budget with some success. Fine tuning of the monetary targets never proved to be feasible, since many of the budget items are subject to heavy forecast errors, such as foreign reserve holdings. Yet, a great part of these errors can be offset by open-market operations, reserve requirement regulations etc., provided the monetary accounts are subject to a careful follow-up and provided the Banco do Brasil is brought under control. Summing up, for a determined Minister of Finance, backed enough by the President of the Republic, the money supply can be managed as a quasi-exogenous variable.

For a less determined Minister of Finance, the monetary budget is a mere piece of formality, since it can be revised at any moment by the National Monetary Council, and since the majority of the members of the Council are Federal Government officials. Moreover, even if the budget is not revised and if its ceilings are largely disrespected, nobody gets penalized. In pratice, a number of public expenditures and subsidies can run through the monetary accounts, thus escaping the fiscal budget which is subject to Congress approval. This incidentally explains the paradoxal coexistence of huge rates of monetary expansion with systematic surpluses in the fiscal budget.

In a word the system is too flexible and too much exposed to political pressures. In the fifteen year period preceeding the creation of the Central Bank, the money supply expanded at an average annual rate of 35%. In the following fifteen years this average rate rose to 40%. This is the best evidence that the Brazilian monetary system never in fact experienced a meaningful reform.

3 — Indexation rules

Indexing, as previously noted, has played a major role in the working of the Brazilian price system since 1964. Different rules of indexation have been adopted for Treasury Bonds, income tax brackets, rents, mortgages and saving accounts, fixed assets, for wages and for the exchange rate. Without entering too much into details, it is important to describe the essentials of such rules, which, incidentally, were subject to a number of changes since 1964.

Indexed Treasury Bonds (the ORTN) were created in july 1964. Their face value was formerly adjusted every quarter, and lately every month. Except in 1973 and 1980, the nominal value of the ORTN has always been determined by a three-month moving average of the general wholesale price index of the Getulio Vargas Foundation. For a number of years, the face value V_t of the ORTN in month t was determined by the formula:

$$V_t = K(P_{t-4} + P_{t-5} + P_{t-6}) \qquad \text{(a)}$$

K indicating a constant, P_{t-i} the wholesale price index for month i.

Lags were lately reduced, since the Getulio Vargas Foundation became able to produce each month's index in the first ten days of the following month, and the following chain-rule was substituted for the above formula:

$$\frac{V_t}{V_{t-1}} = \frac{P_{t-2} + P_{t-3} + P_{t-4}}{P_{t-3} + P_{t-4} + P_{t-5}} \qquad \text{(b)}$$

In July 1975 it was suggested by former Minister of Finance Octavio Bulhões that indexation rules should exclude the effects of supply shocks. For some months the Getulio Vargas Foundation attempted to produce a series of wholesale price indices where supply shocks were duly offset, the so called "accidentality adjustment". The idea was technically sound and was implemented for an interim period, but calculations proved to be rather difficult and discretionary. Thus, in July 1976 the accidentality was put aside and a new chain rule was adopted for the monthly correction of the value of the ORTN:

$$\frac{V_t}{V_{t-1}} = 0{,}8\,\frac{P_{t-2} + P_{t-3} + P_{t-4}}{P_{t-3} + P_{t-4} + P_{t-5}} + 0{,}2023 \qquad \text{(c)}$$

Since $0{,}2023 = 0{,}2\ (1{,}15)^{1/12}$, the formula actually meant partial indexation whenever the inflation rate exceeded 15% a year, presumably on account of systematically adverse supply shocks. For an annual inflation rate of 40%, indexation would be reduced to approximately 35%, etc.

In late 1979 the above formula was coupled with substantial discounts for accidentality, so that the ORTN adjustment lagged 30 percentage points behind the actual inflation rate. For 1980, as previously remarked, the monetary correction of the ORTN has been prefixed far below the general price rise. A similar, although less distorted procedure, had been adopted in 1973, when the Government insistently sticked to the 12% inflation rate target. A new rule, linking the ORTN to the national CPI is now under study by the Government.

Income tax brackets are annually adjusted according to a coefficient at the discretion of the Minister of Finance. Traditionally this coefficient corresponds to some rounded up approximation of the ORTN percentual change.

Saving accounts and mortgages are indexed according to the ORTN, except that the nominal adjustments are made on a quarterly instead of monthly basis. Fiscal debts are also indexed along with the ORTN. Fixed assets, until 1977 were adjusted once a year, according to the annual averages of the wholesale price index. Since 1978 they have been quarterly adjusted with the ORTN.

From August 1968 to December 1979, foreign exchange rates were indirectly indexed by the minidevaluations of the cruzeiro. As a basic guideline, the dollar/cruzeiro rate was changed by small percentages, and at short and irregular intervals (10 to 50 days) according to the inflation rate differential between Brazil and the United States. Domestic inflation, for this purpose, was measured by the industrial wholesale price index, which appropriately reflected the costs of manufactured exports. Slight adjustments were superimposed to this basic rule, taking into account a number of factors, namely: i) the fluctuations of the dollar relative to other major currencies; ii) the inflation rate differentials between the major OECD countries; iii) changes in the terms of trade; iv) domestic balance of payments problems. As previously remarked, this basic rule was broken by the 30% maxidevaluation of December, 7, 1979 and by the predetermination of the cruzeiro devaluation for 1980. A return to the traditional minidevaluation rule was recently announced for 1981, except that domestic inflation will now be measured by the national CPI.

Rents have been subject to various indexation rules. In 1964, law 4494 established two regimes of rent adjustments, one for the so called new rents, the ones under contract as well as the new ones to be contracted, another for the so called old rents. The latter were the result of tenancy laws which indefinitely extended rent contracts after their maturities, at the option of the tenant. Rent adjustments in new contracts should follow the minimum wage increases. The same rule applied to the old rents, plus a supplementary increase intended to gradually phase out the real gap accumulated in the past. In 1967, free choice of escalator clauses was authorized in new rent contracts. As of 1979, rent payments can be adjusted only once a year, proportionally to the change of the ORTN.

Incomes policies found their most powerful weapon in the wage adjustment rules which were legally enforced from 1965 through 1979. They legally bound all collective wage negotiations, leaving no degree of freedom neither for the employers nor for the employees. The market could only work over and above the wage formula in individual negotiations.

The three wage laws, that of 1965, that of 1968 and that of 1975 established that nominal wages should be fixed for periods of twelve months. In order to simplify the presentation of the wage adjustment rules, let us indicate by w_t the logarithm of the nominal wage in year t; p_t will represent the log of the (geometric) average cost of living index in year t; $\tilde{p}_t$ the log of the cost of living index at the end of year t. "Year" is here to be understood as a twelve month period which might begin at any calender month, since different groups of workers obtain wage adjustments in different months of the calendar year. Obviously none of the wage laws referred to logs, although their authors had them clearly in mind. Thus, the actual rules were just very good proxies to the ones presented below.

The 1965 wage law established that nominal wages should be adjusted so that, taking into account prospective inflation, their average purchasing power in the following twelve months would be equal to the average real wage of the past twenty-for months, plus a productivity gain z_t, *i. e.*

$$w_t - p_t^e = 0{,}5\ (w_{t-1} - p_{t-1} + w_{t-2} - p_{t-2}) + z_t \qquad \text{(d)}$$

p_t^e standing for the expected average cost of living index in year t; p_t^e was calculated as

$$p_t^e = \tilde{p}_{t-1} + 0{,}5\ \pi_t^e \qquad \text{(e)}$$

$\pi_t^e = \tilde{p}_t^e - \tilde{p}_{t-1}$ indicating the anticipated inflation rate for the twelve month period during which the nominal wage would be kept fixed. Thus, the 1965 wage rule actually read:

$$w_t = 0{,}5\ (w_{t-1} + \tilde{p}_{t-1} - p_{t-1}) + 0{,}5\ (w_{t-2} + \\ + \tilde{p}_{t-1} - p_{t-2}) + 0{,}5\ \pi_t^e + z_t \qquad \text{(f)}$$

The cost of living indices were calculated by the Ministry of Labor. The productivity gain was set by the Ministry of Planning, at an uniform rate for all working groups. The anticipated rate of inflation (the so called inflationary residual) was determined by the National Monetary Council. So the formula left no room for collective bargaining. It was obviously prohibited to strike against the formula.

Except for market adjustments in individual negotiations, which were never prohibited by law, the 1965 wage formula would actually squeeze the real wages whenever future inflation rates were underestimated. The problem was felt in 1965, 1966 and 1967, when the cost of living increased 45.5%, 41.2% and 24.1%, respectively, compared to prospective inflation rates of 25%, 10% and 15%.

Public complaints against the wage squeeze led the Government to revise the wage adjustment rule in 1968. According to the new law, nominal wages in the previous twelve months should enter into the formula not by their actual values, but by those which would have prevailed if inflation rates were properly foreseen. Summing up, the 1968 formula read as

$$w_t = 0{,}5\ (\hat{w}_{t-1} + \bar{p}_{t-1} - p_{t-1}) + 0.5\ (w_{t-2} + \bar{p}_{t-1} - p_{t-2}) + 0{,}5\ \pi_t^e + z_t \tag{g}$$

where

$$\hat{w}_{t-1} = w_{t-1} + 0{,}5\ (\bar{p}_{t-1} - \bar{p}_{t-2} - \pi_{t-1}^e) \tag{h}$$

Oddly enough, very few people realized that the new wage formula only corrected half-way the inflation underestimation, that of year $t-1$, but not that of year $t-2$, so that, on average, real wages would still be squeezed by one-fourth of the unanticipated cost of living increase. Also very few people remerked that the productivity gain z_t should refer to an eighteen, instead of twelve month period. Anyhow, the formula was accepted until December, 1974. It did not survive by its own merits, but for three other reasons: a) inflation was not that considerably underestimated from 1968 to early 1974; one fourth of the underestimated rate lay in the range of 1 to 2 percentual points, and this was largely offset by the productivity coefficient z_t; b) since the economy was growing at exceptionally high rates, individual negotiation were able to raise the wage substantially above the official adjustment rules, especially those of skilled workers; c) a tough political regime was introduced in the country in December, 13, 1968, under the Institutional Act 5.

The 1974 inflationary outburst brought into discussion the shortcomings of the 1968 wage formula, which was then succeeded by the one established in law 6.147. Accordingly, collective wage adjustments as of January 1st, 1975, were calculated by the formula:

$$w_t = \bar{p}_{t-1} + (w_{t-1} - p_{t-1}) + 0{,}5\ \pi_t^e + \\ + 0{,}5\ (\bar{p}_{t-1} - \bar{p}_{t-2} - \pi_{t-1}^e) + z_t \qquad \text{(i)}$$

which reduced to the past twelve month period the real wage basis, and which introduced full compensation for past inflation underestimation. The National Monetary Council decided that the inflationary residual should enter into the formula at the highly idealized rate of 15% a year, so that π_t^e and π_{t-1}^e were duly cancelled, the formula being applied as:

$$w_t = 1{,}5\bar{p}_{t-1} - 0{,}5\bar{p}_{t-2} - p_{t-1} + w_{t-1} + z_t \qquad \text{(j)}$$

As a good proxy, one can assume $P_{t-1} = 0{,}5\ (\bar{p}_{t-1} + \bar{p}_{t-2})$, so that the new wage formula practically corresponded to

$$w_t = w_{t-1} + \bar{p}_{t-1} - \bar{p}_{t-2} + z_t \qquad \text{(k)}$$

which is a simple wage indexing rule with an average six month lag. From July 1976 to July 1979, the productivity coefficient z_t was adjusted for supply shocks.

In late 1979 a new wage law was enforced, quite different in substance from the former ones. It still can be summarized in the formula:

$$w_t = w_{t-1} + \bar{p}_{t-1} - \bar{p}_{t-2} + f\,(w_{t-1} - p_{t-1}) + z_t \qquad \text{(l)}$$

but the period of adjustment was reduced from twelve to six months; the productivity gain z_t is to be freely bargained between employers and employees, so that in fact the formula sets a floor but no ceiling to collective wage negotiations; there is, moreover, a redistributive percentage $f\,(w_{t-1} - p_{t-1})$, which is positive for the low wages (ten percent up to three minimum wages) and negative for the high ones. This superimposition of mandatory indexing with free bargaining, plus the redistributive coefficient which was brought into the formula, is one of the most cost pushing devices ever invented in the country.

4 — The demand for money

Let us now analyze the behavior of the demand for money in the years 1950 through 1978. Money will be here understood as M_1. A number of empirical investigations have been made on the subject and I will limit myself to focus on two key issues. The first one is the estimation of the income and interest rates elasticities of the demand for money. The second one deals with the shifts of the money demand schedule as a result of financial innovations.

According to a number of empirical studies, one additional percentage point in the nominal interest rate reduces by 0,3% to 0,5% the demand for real cash balances, a well behaved result by international standards. Estimates of the income — elasticity are much more controversial. Some authors believe that the income elasticity of the demand for money has been consistently inferior to one. Many others, however, assert that two oppositive effects have often been garbled in least squares calculations. On the one hand, the demand for money would have increased, as a result of economic growth, with the property of a luxury good. But, on the other hand, financial innovations shifted the money demand function to the left. This view is far from unreasonable, since the Brazilian financial markets experienced remarkable changes throughout the period.

One difficulty must be faced from the outset. Nominal interest rates have often been controlled in the country. Usury laws prohibited, until 1964, nominal interest rates above twelve per cent a year, in spite of the soaring inflation rates. Interest rates have also been controled from 1972 to early 1976, and as of September 1979. Artificial interest rate controls induce households and firms to substitute inventories for money and interest bearing assets. Foreign currencies sold in the black exchange market also provide a substitute for the cash balances held by a selected group of individuals. Thus, in the Brazilian environment, the demand for real cash balances should be specified as a function of three variables, namely the real GDP, the real interest rate and the expected rate of inflation. Unfortunately for econometricians and fortunately for the economic system, interest rate ceilings, whenever artificially set, have been loopholed by a number of ingenuous financial devices. This is to say that it is impossible to produce a reliable interest rate series for the period 1950/78. Moreover, some hypothesis must be introduced or how inflation rate expectations are formed. Since cash holdings can be quickly adjusted, we shall adopt the simplifying assumption of perfect foresight.

Summing up, we estimate the demand for real cash balances in the years 1950 through 1978 as a function of the real GDP and of the actual inflation rate. In order to escape from seasonalities and short term inbalances between the demand and the supply of money we shall limit our calculations to annual average figures. As a first exercise, ordinary least squares produce the following equation (parentheses indicate the t-statistics):

$$m_t - p_t = \underset{(14.9678)}{4.3286} + \underset{(14.4035)}{0{,}1574 y_t} - \underset{(1.9430)}{0{,}5618 (p_t - p_{t-1})} + \qquad R^2 = 0.8900 \quad DW = 1.0016$$

where m_t stands for the log of the money supply in billion cruzeiros; p_t for the log of the general price index and y_t represents the log of the real GDP index. All the coefficients in the above equation have the signs predicted by the theory and it is suggested that the income elasticity of the demand for money is inferior to one. Yet, the above regression equation is unserviceable, because of the poor Durbin-Watson statistics.

Things are not improved by the introduction of time as an additional explanatory variable, on account of financial innovations. As one might naturally suspect, the shift and the income effects are improperly estimated because of multicollinearity. In fact, ordinary least squares lead to a shift coefficient with the wrong sign and to an incredibly low estimate of the income elasticity. None of these results nor the t-statistics are to be taken seriously, because of the miserable Durbin-Watson statistics.

$$m_t - p_t = \underset{(12.6608)}{4.3286} + \underset{(1.5161)}{0{,}1574 y_t} - \underset{(3.3109)}{0{,}5618 (p_t - p_{t-1})} + \underset{(4.8838)}{4{,}1359\,(t - 1950)} \qquad R^2 = 0.9438 \quad DW = 0.5214$$

The serial correlation of the residuals in the above regression equations suggests that financial innovations would shift the money demand function according to a trend component plus a random walk. This hypothesis can be tested using OLS to the first differences, *i. e.*, estimating the rate of expansion of the real demand for money as a function of the rate of growth of real GDP and of the changes in the inflation rate:

$$\mu_t - \pi_t = \underset{(1.6525)}{-3.9475} + \underset{(4.1523)}{1.2980\, \eta_t} - \underset{(3.8544)}{0.4441\, (\pi_t - \pi_{t-1})} \qquad R^2 = 0.5695 \quad DW = 2.1751$$

where $\mu_t = 100\,(m_t - m_{t-1})$, $\eta_t = 100\,(y_t - y_{t-1})$, $\pi_t = 100\,(p_t - p_{t-1})$.

The explanation coefficient 0,57 is quite acceptable for a first difference equation, and all the remaining estimates appear to confirm the predictions of the theory. As a result of the financial innovations, the demand for money probably moved to the left, at an annual average rate of approximately 4%. An increase in the inflation rate, which can be used as a proxy to the nominal interest rate, contracts the real cash balance with the same intensity that has been indicated in most empirical studies. And, as a theoretical relief, money now does no more appear to be subject to Engel's law.

Open-market operations, with short term Federal Bonds and repurchase agreements by dealers, were the most outstanding financial innovation of the last decade. It is natural to question whether they have been more or less important than the preceeding financial innovations of the fifties and of the sixties. A test is provided by the introduction of a dummy variable d as an additional explanatory variable in the previous regression equation. We shall take $d = 0$ until 1967 and $d = 1$ thereafter:

$$\mu_t - \pi_t = \underset{(1.6013)}{-3.9681} + \underset{(3.4851)}{1.3072\,\eta_t} - \underset{(3.7772)}{0.4441\,(\pi_t - \pi_{t-1})} - \underset{(0.0466)}{0.1106\,d}$$

$$R^2 = 0.5699$$
$$\mathrm{DW} = 2.1744$$

The dummy has no statistical significance, supporting the hypothesis that financial innovations since 1968 have been as important as the ones of the preceeding eighteen years.

Up to now we have been assuming that supply and demand are always in equilibrium in the money market. This is not an intolerable simplification, since we have only dealt with annual average figures. Yet, one might improve the explanation coefficient of our regression equations by assuming an adjustment function of the form:

$$m_t - \bar{m}_t = k\,(m_t - m_{t-1})$$

where m_t indicates the money supply and $\bar{m}_t$ the money demand, both in logs. We are led to introduce $\mu_t - \mu_{t-1}$ as an additional explanatory variable for the changes $\mu_t - \pi_t$ in the real cash balances:

$$\mu_t - \pi_{t-1} = \underset{(1.4252)}{-3.0182} + \underset{(4.1380)}{1.1512\,\eta_t} - \underset{(5.1068)}{0.5418\,(\pi_t - \pi_{t-1})} + \underset{(2.9425)}{0.3295\,(\mu_t - \mu_{t-1})}$$

$$R^2 = 0.6831$$
$$\mathrm{DW} = 1.6387$$

5 — The Phillips relation

Most empirical investigations on inflation-output trade-offs in Brazil have been carried out by Antonio Carlos Lemgruber.[2] The results of such investigations are usually understood as supporting the natural unemployment rate hypothesis. Since unemployment figures have never been regularly estimated for Brazil until late 1979, Lemgruber fits a geometric trend to the series of real industrial output; and correlates deviations from the trend with his estimates of the unanticipated inflation rate. Lemgruber has properly limited his trade-off studies to industrial output. Deviations from trend in agricultural production basically depend on climatic events. They are cause but not effect of unanticipated inflation.

The assumption that industrial capacity expands as a geometric progression is to be considered as an archeological simplification intended to circumvent the absence of unemployment statistics. Data on the average utilized capacity, however, are available since 1969. The series is too short to provide the appropriate inputs to a Phillips curve. Yet, it can be used to test the adequacy of the hypotheses on the industrial trend, since one should expect deviations from trend to be strongly correlated with the average utilization of industrial capacity.

The log of Lemgruber's industrial trend for the period 1950/78 is a straight line with a 8,13% slope. The correlation coefficient between his estimated deviations from trend and the average utilized capacity is no more than 0.46, for the period 1969/78. This result suggests that Lemgruber's findings should be confronted with the ones emerging from some modified series of potential industrial output. Modified trends will be introduced later on.

One of the best Phillips relations estimated by Lemgruber for the period 1950/78, using annual data, has been the following:

$$\begin{array}{llll} h_t^I = -0.1133 + 0.7466 h_{t-1}^I + 0.2095\,(\pi_t - \pi_{t-1}) & & R^2 & = 0.7648 \\ \qquad\quad (0.15) \qquad\quad (6.79) \qquad\qquad\quad (2.02) & & DW & = 1.4875 \\ & \text{Durbin} & h & = 1.6672 \end{array}$$

where h_t^I indicates 100 times the deviation of the log of industrial output from trend; π_t, as previously, indicates the percentual rate of inflation. The constant term, as expected in a Phillips relation, is not statistically significant. Deviations from the trend, however, are strongly subject to serial correlation. The result is not surprising, since the business cycle is not a white noise, but the high regression

[2] Lemgruber (1979).

coefficient on h^I_{t-1} requires some explanation. Lemgruber just sends us back to the Lucas supply function, where serial correlation in output is introduced as an *ad hoc* assumption, and accepts $\pi_t - \pi_{t-1}$ or some fraction of it as a good proxy for the unanticipated part of the inflation rate. Adjustment costs as pointed out by Sargent, could be responsible for serially correlated deviations from the trend, but I think that an alternative explanation is more suitable to the Brazilian case.

As a point of departure, let us assume that the industrial supply function is described by

$$y^I_t = \bar{y}^I_t + h^I_t = a_t + b\,(p_t - w_t) + u_t \qquad (1)$$

where y^I_t and $\bar{y}^I_t$ indicate the logs of actual and potential industrial output, respectively, where p_t and w_t stand for the logs of the general price and wage levels and where u_t is a random walk.

Let us now assume that nominal wages are indexed with a one year time lag, plus a productivity gain z_t plus an adjustment factor to the deviations of industrial output from the trend, plus a serially independent random variable e_t:

$$w_t - w_{t-1} = p_{t-1} - p_{t-2} + z_t + ch_{t-1} + e_t \qquad (2)$$

Taking the first differences in equation (1) and introducing the above wage adjustment hypothesis yields:

$$h^I_t = \alpha_t + b\,(\pi_t - \pi_{t-1}) + (1 - bc)\,h^I_{t-1} + \varepsilon_t \qquad (3)$$

where $\alpha_t = a_t - a_{t-1} - (\bar{y}_t - \bar{y}_{t-1}) - bz_t$ should be zero under proper settlements for the productivity gain z_t, and $\varepsilon_t = u_t - - u_{t-1} - be_t$ is white noise. Equation (3) has the precise specification of the Phillips relation estimated by Lemgruber and is perfectly consistent with strong serially correlated deviations of industrial output from trend.

It should be noted that $c = 0$ in equation (2) would lead us to a non-accelerating Phillips relation, namely to a permanent trade-off between the inflation rate and the deviation of output from the trend. Lemgruber's findings support the natural unemployment rate hypothesis in the sense that they reject the hypothesis $c = 0$ at 98% confidence level.

Another interesting result pointed out by Lemgruber is the evidence of asymetrical impacts of inflation acceleration and deceleration on industrial output. The following equation, estimated for the period 1950/78 suggests that inflation can only decrease at the price of a temporary industrial slow-down, but that acceleration of the inflation rate has no real effects:

$$h_t^I = \underset{(2.06)}{2.1662} + \underset{(7.81)}{0.7765} h_{t-1}^I + \underset{(3.59)}{0.5988} (\pi_t - \pi_{t-1})^- -$$

$$- \underset{(0.81)}{0.1223} (\pi_t - \pi_{t-1})^+ \qquad R^2 = 0.8211 \quad DW = 1.6817$$

Lemgruber explains his kinked phillips curve findings as a result of a mix of rational expectations and inflation rigidity. Wages would be adjusted according to rational inflationary expectations if and only if inflation rates were accelerating. I have myself elsewhere suggested that indexation with a lag on reference prices could be responsible for this as asymetrical behavior. A more careful analysis of the statistical data, however, discards these sophisticated explanations. In the 1950/78 period, the inflation rate experienced six big leaps, the ones of 1954, of 1959, of 1962 and 1963, of 1974 and 1976. In the first two cases the acceleration of inflation coincided with generous minimum wage increases. In 1962 and 1963 nominal wages were strongly pushed up by the Goulart populist Government. Inflation accelerated in 1974 because of the oil shock and because prices had been strongly repressed throughout 1973. Only the 1976 leap cannot be mostly ascribed to unusual events, in spite of some adverse suply shocks.

Summing up, *I* prefer to stay with Lemgruber's symetrical Phillips relation. *I* still assume that wage indexing in the Brazilian way creates some sort of inflation rigidity, in a sense which will be explained in section 7. Yet, since wage adjustment rules were subject to dramatic changes throughout the period, and since full wage indexing was just introduced in 1975. *I* do not believe that rigidities can be properly identified by least squares exercises on the 1950/78 data.

As previously noted, Lemgruber findings depend on his simplifying assumption of a geometric industrial trend. It is important, therefore, to check whether his conclusions hold when other plausible hypothesis on the trend are adopted.

As alternative exercises, let us determine the trend equations by fitting a parabola and a cubic polynominal to the logs of the indices of real industrial production. Potencial industrial output is now determined by one of the following equations:

$$\hat{y}_t^I = 3.127563 + 0.068140\hat{t} + 0.000465\hat{t}^2$$

$$\hat{y}_t^I = 3.061284 + 0.099318\hat{t} - 0.002368\hat{t}^2 + 0.00006743\hat{t}^3$$

where $\bar{t} = t - 1950$, and with the understanding that the above equations should only be applied to the 1950/78 period, since it would be obviously absurd to extrapolate parabolic or cubic trends. With the parabolic trend, the correlation coeficient between deviations from trend and the average utilized industrial capacity, from 1969 to 1978, increases to 0.71. The cubic trend raises this correlation coefficient to 0.91.

With the above trend equations, the Phillips relations are now described by the following regression equations:

$$h_t^I = -\ 0.1517 + 0.7149\ h_{t-1}^I + 0.2119\ (\pi_t - \pi_{t-1})$$
$$(0.2003) \quad (6.0982) \quad (2.0606)$$
$$R^2 = 0.7382$$
$$DW = 1.3725$$
$$\text{Durbin } h = 2.1167$$

$$h_t^I = -0.5515 + 0.6602\ h_{t-1}^I + 0.2055\ (\pi_t - \pi_{t-1})$$
$$(0.7049) \quad (4.7814) \quad (1.9350)$$
$$R^2 = 0.6543$$
$$DW = 1.2834$$
$$\text{Durbin } h = 2.7768$$

which provide some additional support to Lemgruber's findings. In fact, the hypothesis that the log of the industrial trend is a straight line is highly arbitrary, and might lead to strong estimations errors in the h_t^I. Yet, the fact that deviations from trend are strong serially correlated, makes the Phillips relation somewhat insensitive to trend errors. (To get an extreme case, put $c = 0$ in equation (2). Equation (3) becomes a relation between the rate of industrial growth and the acceleration of the inflation rate, independently of the industrial trend).

6 — Inflation and industrial growth in the 1950/78 period — A retrospective analysis

The Phillips relation estimated by Lemgruber and the last money demand equation in section 4 provide a monetarist explanation of the Brazilian experience of inflation and industrial growth, in the sense that inflation rates, as well as deviations of industrial output from trend are determined by changes in the money supply combined with demand and supply shocks. Although we are not engaged in econometric modeling, a worthwhile exercise is to check to what extent these equations can explain the Brazilian rates of inflation and industrial growth throughout the 1950/78 period. (Some two-stage least squares tests suggest that the simultaneous equation biases can be tolerated).

Our system of equations, in the absence of shocks, reads as follows:

$$\mu_t - \pi_t = -3.0182 + 1.1512\,\eta_t - 0.5418\,(\pi_t - \pi_{t-1}) +$$
$$+ 0.3295\,(\mu_t - \mu_{t-1})$$
$$h_t^I = 0.7466 h_t I_1 + 0.2095\,(\pi_t - \pi_{t-1})$$

Where the constant term in the Phillips relation, with no statistical significance, has been conveniently deleted.

We still need a link between the real rate of growth of GDP and the deviations of industrial output from trend, which can be approximated by the following cousin of Okun's law:

$$\eta_t = 6.84 + 0{,}841\,(h_h^I - h_{t-1}^I)$$

Solving for η_t and h_t^I, we are led to the following reduced form equations:

$$\pi_t = -0.5128\pi_{t-1} + 0.3711 h_{t-1}^I + 1.0143\mu_t +$$
$$+ 0.4985\mu_{t-1} - 7.3460$$
$$h_t^I = -0.3169\pi_{t-1} + 0.8244 h_{t-1}^I + 0.2125\mu_t +$$
$$+ 0.1044\mu_{t-1} - 1.5390$$

Some interesting features of this system of diference equations should be pointed out.

The system is conveniently stable, since the eigenvalues of the matrix

$$\begin{bmatrix} -0.5128 & 0.3711 \\ -0.3169 & 0.8244 \end{bmatrix}$$

are both inferior to one in absolute value (—0.4165 and 0.7281). The negative eigenvalue is responsible for some unimportant cyclical effects. The positive root is large enough to produce a much more significant result: deviations of industrial output from trend only adjust quite slowly. This is a result of the strong serial correlation in industrial output deviations from trend.

The system provides a monetarist explanation for inflation and growth in one basic sense: if money is expanded at a constant rate μ, and if shocks are entirely absent, the inflation rate will converge to μ — 4,9 and the output gap will tend to zero. Yet, the convergence speed is quite different for inflation rates and for output deviations from trend.

A once for all price shock has some significant lagged effects one year later, but virtually phases out after the second year, according to the following interim multipliers.

Period	0	1	2	3	4
Δr_t	1	—0.5128	0.1454	– 0.1112	0.0097
Δh_t^f	0	—0 3169	– 0.0987	—0.1275	– 0.0699

A shock in output has a much more prolonged real effects, as shown below:

Year	0	1	2	3	4
Δr_t	0	0.3711	0.1156	0.1493	0.0818
Δh_t^f	1	0.8244	0.5620	0.4267	0.3045

A temporary change in the rate of expansion of the money supply has almost only contemporaneous effects on both, inflation rates and deviations of industrial output:

Year	0	1	2	3	4
$\Delta \mu_t$	1	0	0	0	0
$\Delta \pi_t$	1.0143	0.0572	—0.0448	0.0035	—0.0126
Δh_t^f	0.2125	—0.0418	—0.0526	—0.0291	—0.0251

A permanent change in the rate of expansion of the money supply accumulates the above interim effects. As a result, the inflation rate promptly reacts (there is an obvious lag which is obscured by our analysis with annual average data). The real side effects, however, persist for a number of years:

Year	0	1	2	3	4
Δr_t	1	1	1	1	1
$\Delta \pi_t$	1.0143	1.0715	1.0267	1.0301	1.0175
Δh_t^f	0.2125	0.1707	0.1181	0.0889	0.0637

The conclusion is that if inflation rates are quickly brought down by a shock treatment, a "miracle" is likely to be sustained for a number of following years. To take a concrete example, let us start in 1967, when the money supply expanded 31.4%, when the rate of inflation was 24.9% and when, according to Lemgruber, industrial output was 18.2% below trend. (This gap has been probably overestimated but, as we have previously remarked, errors in the trend do not significantly affect our results, because of the strong serial correlation of the gaps). Let us indicate by $\hat{\pi}_t$, $\hat{h}_t^I$, $\hat{\eta}$ the inflation, the industrial gap and the real GDP growth rates which would have prevailed if, from 1968 through 1971, the Brazilian economy actually behaved according to our equations, if no shocks had occurred and if money expanded as it actually did. As one can note below, the real world did not perform very differently from our imaginary one:

Year	68	69	70	71
u_t	34.5	28.0	25.1	26.6
$\hat{\pi}_t$ (simulated)	23.8	20.9	17.3	20.0
π_t (actual)	21.7	18.9	18.1	18.6
$\hat{h}_t^I$ (simulated)	—13.82	—10.92	—8.91	−6.09
h_t^I (actual)	—13.80	—10.50	—8.75	−3.51
$\hat{\eta}$ (simulated)	10.5	9.3	8.5	9.2
η (actual)	10.6	9.5	8.4	12.2

The miracle was that, in spite of moderately expansive monetary policies, inflation rates declined while industrial output was able to grow at an average annual rate 3.7% above trend. Industrial output would only have grown at is trend rate (8.13% per annum) if the same monetary policies were followed but if no output gap existed in 1967. And the inflation rate, although coming down to 21.5% in 1971, would have been substantially higher in the interim years, especially in 1968.

Table 1 below compares the actual inflation rates and industrial output deviations from trend with the ones resulting from the previous reduced form equations. For a simple monetarist model, with only two equations, results are not bad. Most of the prediction errors find some simple factual explanation. After all, in a highly regulated economy, with a proliferation of price, wage, exchange and interest rate controls, one should not expect annual inflation rates and output gaps to be solely explained by the rates of expansion of the money supply. Inflation rates are significantly

Table 1

Year	μ_t	π_t	$\hat{\pi}_t$	h_t^f	$\hat{h}_t^f$
1951	22.6	15.3	21.7	22.22	5.44
1952	13.7	11.2	10.8	—1.05	0.71
1953	16.6	13.8	—10.2	—0.84	—1.00
1954	20.9	23.9	14.7	—0.66	—0.43
1955	17.5	15.2	8.5	1.31	—3.76
1956	18.4	18.2	12.7	—0.14	0.46
1957	20.4	13.1	13.3	—2.72	—1.17
1958	29.1	12.2	24.5	4.15	0.32
1959	24.5	32.1	27.3	7.24	6.26
1960	33.8	25.6	25.4	8.29	4.00
1961	36.9	31.5	36.9	10.30	8.55
1962	44.0	41.6	43.3	9.67	10.17
1963	46.2	56.2	43.7	1.73	7.66
1964	61.4	64.4	49.8	—1.37	— 0.05
1965	60.7	45.0	51.3	—14.30	— 3.77
1966	30.3	32.3	25.2	— 13.10	—14.81
1967	31.4	24.9	18.2	—18.20	—12.74
1968	34.5	21.7	23.8	— 13.80	—13.82
1969	28.0	18.9	22.0	—10.50	—10.24
1970	25.1	18.1	18.5	—8.75	—7.93
1971	26.6	18.6	19.6	3.51	—6.22
1972	27.0	15.7	22.5	0.93	— 1.81
1973	38.5	14.1	37.5	7.49	5.25
1974	32.5	25.2	40.4	8.77	11.09
1975	28.5	24.5	28.1	6.70	7.15
1976	34.1	34.6	31.4	8.71	6.44
1977	32.0	35.6	27.6	4.38	5.04
1978	33.9	32.7	26.4	4.04	1.33

SOURCE: See Table 2.
NOTES: μ_t, π_t, h_t^f: actual percentages: money supply growth, inflation rate, industrial output deviation from trend; μ_t, h_t^f: predicted values based on previous year actual values.

above predictions in the years of strong wage pushes such as the 1954/56 period, the Goulart 1963, and the corrective inflation period extended throughout the Castello Branco administration. Inflation also exceeded the monetarist forecasts in the adverse supply-shocks period from 1976 to 1978. Prices increased much less than money would explain in years of tight controls, such as 1951, 1958, and, generally speaking, the 1972/74 three year period.

The biggest discrepancy in the industrial gap forecast occurs in 1965, when demand was strongly reduced by a fiscal and psychological shock. Industrial output had already fallen below projections in 1963, when the populist official policies stimulated every sort of adverse domestic supply shock.

7 — Notes on indexation

Escalator clauses are by no means a recent economic proposal. They were explicitly suggested in 1807 by an English writer on money, John Wheatley, and were enthusiastically defended by at least two great economists of the past, Alfred Marshall and Irving Fisher. The success of the practical experiences with widespread indexation is highly controversial, Hyperinflation epidemies were always accompanied by quickly adjustable escalator clauses, but indexation was much more the effect than the cause of the price virulence. Mandatory indexation of wage contracts was adopted by some Europen countries after the Second World War. In most cases the experience was soon abandoned because of the alledged feed-back effects on inflation rates. In the last fifteen years widespread indexation was introduced in a number of Latin-American countries, the most comprehensive experience being that of Brazil, where the indexation rules were described in section 3. There is no doubt that, at least in the Brazilian case, indexing removed most of the economic distortions traditionally caused by chronic inflation. In particular, it restored the functioning of a number of markets which had virtually disappeared because of the uncertainty in long term price forecasts: mortgages, long term bonds and long term rent contracts. Inflation accounting was made mandatory with a double benefit for corporate savings: illusory profits were neither distributed nor subject to income tax. In spite of all these advantages, indexation is usually understood as a fruitful device to live with inflation, but as the cost of almost perpetuating price increases.

Widespread indexation, or at least mandatory indexation was strongly criticized by John Maynard Keynes in Chapter 19 of *The General Theory,* from which the following passage is quoted:

> "If, as in Australia, an attempt were made of fix real wages by legislation, then there would be a certain level of employment corresponding to that level of real wages; and the actual level of employment would, in a closed system, oscillate violently between that level and no employment at all, according as the rate of investment was or was not below the rate compatible with that level; whilst prices would be in unstable equilibrium when investment was at the critical level, racing to zero whenever investment was below it, and to infinity whenever it was above it. The element of stability would have to be found, if at all, in the factors controlling the quantity of money, being so determined that there always existed some level of money-wages at which the quantity of

money would be such as to establish a relation between the rate of interest and the marginal efficiency of capital which would maintain investment at the critical level. In this event employment would be constant (at the level appropriate to the legal real wage) with money-wages and prices fluctuating rapidly in the degree just necessary to maintain this rate of investment at the appropriate figure. In the actual case of Australia, the escape was found, partly of course in the inevitable inefficacy of the legislation to achieve its object, and partly in Australia not being a closed system, so that the level of money wages was itself a determinant of the level of foreign investment, and hence of total investment, whilst the terms of trade were an important influence on real wages."

Yet, in a challenging essay published in 1974, Milton Friedman suggested that the adoption of widespread indexing clauses would not only neutralize the inflationary distortions, but also reduce the side effects of anti-inflationary policies. The central argument by Friedman is that these side effects fundamentally reflect distortions introduced in relative prices by unanticipated inflation or deflation. These distortions arise because contracts are entered into in terms of nominal prices under mistaken perceptions about the likely course of inflation. Widespread adoption of indexation clauses would minimize the effects of the expectational errors, easing the side effects of anti-inflation masures. Friedman neither takes into account supply shocks nor the possibility of inflation rigidities introduced by indexation. The second omission seems perfectly excusable, since Friedman never advocated mandatory indexation.

Let us make some theoretical comments on this controversial issue. For sake of simplicity let us confine our discussion to wage indexation.

As a point of departure, let us assume an aggregate supply function of the form:

$$y_t = a_t + b\ (p_t - w_t) + u_t \tag{4}$$

where y_t, p_t and w_t are the logs of real output, price and wage levels, and where u_t stands for a supply shock. This is essentially the same as equation (1), except that now y_t represents total (and not industrial) output, and the shock u_t is not bound to be a random walk.

A number of wage rules will be entered into this supply function. We shall first assume that wages are flexibly determined according to rational expectations on prices on the basis of information available at the end of period $t-1$

$$\hat{y}_t = a_t + b\,(E_{t-1}\,p_t - w_t) + E_{t-1}\,u_t \tag{5}$$

where $\hat{y}_t$ stands for the log of potential output and where the operator E_{t-1} indicates the orthogonal projection on the information subspace available in period $t-1$. [3]

Equations (4) and (5) lead to the abridged Lucas supply function:

$$y_t - \hat{y}_t = b\,(I - E_{t-1})\,p_t + (I - E_{t-1})\,u_t \tag{6}$$

Where the neutrality theorem holds, since $E_{t-1}\,(y_t - \hat{y}_t) = 0$, and since no policy rule can affect the unexpected component of $y_t - \hat{y}_t$. This is a highly comfortable hypothesis, since it excludes any short or long run trade-off between price and output stabilization policies. Actually, anti-inflationary policies would be painless under such a supply behavior.

As noted by Stanley Fischer, wages might be determined according to rational expectations on prices. Yet, since wage contracts usually extend over a number of periods, different information sets become involved in the supply function, which now assumes the form:

$$\begin{aligned} y - \hat{y}_t = {} & b_1\,(I - E_{t-1})\,p_t + b_2\,(I - E_{t-2})\,p_t + \dots + \\ & + b_n\,(I - E_{t-n})\,p_t + u_t \end{aligned} \tag{7}$$

where the supply shock u_t is orthogonal to the information set of period $t-n$, but not necessarily to that of period $t-1$.

Supply equation (7) brings us down to earth, since it is able to explain why a temporary slow down in the economy is the usual cost of anti-inflationary polices. It also introduces a difference between optimum policy rules for price and for output stabilization when supply shocks are serially correlated. The point now is that money-wages are temporarily rigid, so that real wages depend on

3 We shall treat rational expectation by vector methods. We first define a vector space H generated by the constant 1 and by a finite set of random variables with a multivariate normal distribution. The inner product is defined by the usual rule, $(x,y) = E(xy)$, where E indicates mathematical expectation. An information set L is a subspace including the constant 1. $E_L x$, i.e. the orthogonal projection of the vector x on L, is the conditional expectation of x to the information set L. $(I - E_L)\,x$ stands for the unexpected part of x, I indicating the identity operator. The square of the norm of this unexpected part, i.e. $|(I - E_L)\,x|^2$ is the conditional variance of x to the information set L.

inflation rates. The optimum price stabilization rule entirely transfers to output the effects of the identified supply shocks. The optimum output stabilization policy passes all these effects into prices.

Let us now introduce two alternative wage indexation rules:

$$w_t = w_0 + p_t + x_t \tag{8}$$

$$w_t = \hat{w}_0 + p_{t-1} + x_t \tag{9}$$

Formula (8) describes instantaneous or *ex post* indexing: nominal wages are automatically adjusted to changes in the general price level plus a productivity increase x_t. Formula (9) describes a more practical arrangement whereby nominal wages are adjusted to the general price level with a certain lag. Since wages are usually paid before the price index of the period has been published, there are good reasons to prefer this type of indexing arrangement. In this case, because of the adjustment lag, $\hat{w}_0$ is expected to incorporate an allowance for one period inflation.

In spite of formal similarities, wage rules (8) and (9) lead to substantially different supply curves:

$$y_t = a_t - b\,(w_0 + x_t) + u_t \tag{10}$$

$$y_t = a_t + b\,(p_t - p_{t-1}) - b\,(\hat{w}_0 + x_t)\ u_t \tag{11}$$

With instantaneous indexing, described by equation (10), aggregate supply becomes totally inelastic with respect to the general price level. This obviously eliminates any temporary or permanent conflict between price and output stabilization policies. In fact there is only one possible output stabilization device, and it belongs to the controversial field of incomes policies: to establish the productivity gain x_t as a stochastic process, subject to a rule $x_t = E_{t-1}x_t$ such that:

$$\bar{y}_t = a_t - b\,(w_0 + x_t) + E_{t-1}u_t \tag{12}$$

according to which the productivity coefficient should absorb all the supply shocks.

Equation (11), associated to lagged indexing, is a conventional and not an accelerating Phillips curve; output is an increasing function of the actual rate of inflation $p_t - p_{t-1}$ and not of its unanticipated component. This opens the possibility of stabilizing output through the perpetuation of the inflation rate. Surely, price and output stability could still be reconciled if the wage basis $\hat{w}_0$ and the productivity rule x_t were established so that:

$$\bar{y}_t = a_t - b\ (\hat{w}_0 + x_t) + E_{t-1}\ u_t \qquad (12.a)$$

which would mean that indexation would have been introduced in a stable price environment. I have never heard of such an experiment. Indexation is usually introduced after a prolonged inflation period. The wage basis $\hat{w}_0$ is seldom revised downward and combined with a productivity rule according to equation (12.a). The right side usually stands significantly below potential output, and governments are tempted to achieve full employment through accepting a chronic inflationary process. This seems to be the explanation for the conventional wisdom which considers indexation as an inflation perpetuating device.

Mandatory indexation apparently just differs from voluntary indexation in its legal content. Yet, there are some important economic consequences. An escalator labor contract involves two key elements, the wage basis w_0 and $\hat{w}_0$ and the productivity coefficient x_t rule. Errors can be commited when these elements are negotiated. The possibility of output stabilization without resort to accelerating inflation therefore depends on the possibility of periodically renegotiating the wage basis and the productivity rule. Such renegotiations might be relatively easy to conduct if indexing arrangements are a voluntary clause in labor contracts. Yet, they become a political nightmare when the indexation and productivity rules are imposed by law.

Even with mandatory indexation, market forces prevent output from indefinitely deviating from the trend. A possibility, is that actual nominal wage adjustments deviate from the indexation rule according to an increasing function of the output ga, *i. e.*:

$$w_t - w_{t-1} = p_{t-1} - p_{t-2} + z_t + f\ (y_{t-1} - \bar{y}_{t-1}) \qquad (13)$$

Where $z_t = x_t - x_{t-1}$ and $f\ (0) = 0$. The problem is that $f\ (y_{t-1} - \bar{y}_{t-1})$ is likely to be a strongly convex function. It might even present a kink at the origin. In fact, since indexing laws set floors but not ceilings to wage adjustments, nothing prevents the market to exceed the legal nominal wage increase, except perhaps, a little bit of inertia. Now, to actually adjust nominal wages below the legal requirements, there is only one costy procedure: to lay off workers and contract some cheaper substitutes. This might be the reason for a kink in our adjustment function. In a simplified version, one could assume $f\ (y_{t-1} - \bar{y}_{t-1}) = c\ (y_{t-1} - \bar{y}_{t-1})$ for negative deviations from trend and

$f(y_{t-1} - \hat{y}_{t-1}) = C(y_{t-1} - \hat{y}_{t-1})$ for positive deviations, with $c < C$. We would now be led to an accelerating Phillips curve similar to the one described in equation (3), section 5, but with a highly uncomfortable peculiarity. The regression coefficient $1 - bc$ or $1 - bC$ of the deviation of output from trend on its lagged value would be higher to the left than to the right.

Instantaneous indexing, has two important theoretical virtues. First, as we have already shown, it eliminates any short or long run conflict between price stabilization and output stabilization. Since aggregate output becomes insensitive to the price level, inflation can be fought without unpleasant side-effects on output and employment. Second, it also prevents unexpected shifts in output caused by demand shocks. To prove the latter proposition, let us apply the operator $(I - E_{t-1})$ to both sides of supply equation (10). Since a_t and w_0 are nonrandom, and since x_t is subject to some rule, *i.e.* $x_t = E_{t-1}x_t$:

$$(I - E_{t-1})\,y_t = (I - E_{t-1})\,u_t \tag{14}$$

or, taking the squar of the norm of both sides:

$$\mathrm{var}_{t-1}\,y_t = \mathrm{var}_{t-1}u_t \tag{15}$$

var_{t-1} indicating the conditional variance to the information set available in period $t - 1$.

Since demand shocks do not affect output, they must be entirely transfered to prices when wages are instantaneously indexed. To prove the above proposition let us assume that demand can be described by the velocity equation:

$$y_t = m_t - p_t + e_t \tag{16}$$

where e_t stands for the demand shock.

The money supply is supposed to follow some rule, *i.e.* $(I - E_{t-1})\,m_t = 0$. Hence:

$$(I - E_{t-1})\,y_t = (I - E_{t-1})\;(e_t - p_t) \tag{17}$$

Combining (14) and (17):

$$(I - E_{t-1})\;p_t = (I - E_{t-1})\;(e_t - u_t)$$

Assuming that the unexpected components of the supply and demand shocks are uncorrelated:

$$\mathrm{var}_{t-1}\;p_t = \mathrm{var}_{t-1}\;u_t + \mathrm{var}_{t-1}\;e_t \tag{18}$$

To have a standard for comparison, let us calculate the conditional variances of prices and output if wages, instead of instantaneously indexed, were flexibly negotiated according to equation (5). Insted of (14) we would now have:

$$(I - E_{t-1})\, y_t = b\,(I - E_{t-1})\, p_t + (I - E_{t-1})\, u_t \qquad (19)$$

Combining (17) and (19), simple calculations yield:

$$\mathrm{var}_{t-1}\, y_t = \frac{\mathrm{var}_{t-1}\, u_t + b^2 \mathrm{var}_{t-1}\, e_t}{(1+b)^2} \qquad (20)$$

$$\mathrm{var}_{t-1}\, p_t = \frac{\mathrm{var}_{t-1}\, u_t + \mathrm{var}_{t-1}\, e_t}{(1+b)^2} \qquad (21)$$

The same formulae apply to the Stanley Fischer supply equation (7), taking $b = b_1 + \ldots + b_n$.

We can now use the above results to identify the black face of instantaneous indexing. First, although it protects output against demand shocks, it overexposes it to impact of unexpected supply shocks. Second, it increases the sensitivity of prices to either type of shock. A more serious problem arises when supply shocks are serially correlated: output can only be stabilized at its potential level if the stochastic productivity rule (12) is adopted. Although some proxies to this rule have been occasionally implemented (for instance, in Brazil, from july 1976 through August 1979) it is hard to convince workers that their increase in physical productivity should be adjusted for changes in the sectoral terms of trade. Still more dangerous is an arbitrary productivity rule imposed by law, where errors might exceed supply shocks.

Lagged indexation, described by supply equation (11), poses other types of problems. Prices and real wages still keep some degree of flexibility, and in fact the conditional variances of prices and output are described by equations (20) and (21). The problem is that now errors in the wage basis or in the productivity rule are not necessarily transmited to output. They generate, as we have already mentioned, a very unfortunate Phillips relation.

Since instantaneous indexing has some virtues which are not shared by its lagged version, one might be tempted to recommend the shortening of the adjustment lag, to get the best possible proxy. Except if the wage basis $\bar{w}_0$ is properly revised downward, this is a toxic medicine. For it will simply accelerate the annual rate of inflation that will keep output at its potential level. Neglecting productivity and shocks in equation (11), the rate of inflation which will keep full employment is determined by:

$$p_t - p_{t-1} = \frac{\hat{y}_t - n_t}{b} + \bar{w}_0$$

The right side of the above equation is the annual rate of inflation if wages are adjusted once a year. It becomes the quarterly rate of inflation if wages are adjusted every three months.

Summing up, indexing is a powerful but a dangerous tool. It makes life easier with inflation and, in highly idealized circumstances, it might even reduce the unpleasant side-effects of anti-inflationary policies. Supply shocks are the first instance of malfunctioning of the system. Mandatory indexing might create rigidities which are hard to correct. And lags in indexation arrangements often induce policy makers to perpetuate inflation in order to achieve full employment.

Table 2

Years	P	π	M	μ	y	y^I	y^A	h	h^I	h^A
1950	1.1	10.5	61.4	19.9	28.7	22.4	48.2	3.52	4.16	—0.88
1951	1.3	15.3	77.0	22.6	30.4	23.8	48.5	2.43	2.22	—4.21
1952	1.4	11.2	88.3	13.7	33.0	25.0	52.9	3.03	—1.05	0.48
1953	1.6	13.8	104.2	16.6	33.9	27.2	53.0	—0.41	—0.84	—3.36
1954	2.1	23.9	128.4	20.9	37.3	29.5	57.2	2.38	—0.60	0.20
1955	2.4	15.2	152.9	17.5	30.9	32.7	61.6	2.19	1.31	3.61
1956	2.9	18.2	183.8	18.4	41.1	34.9	60.1	—1.52	—0.14	—2.82
1957	3.3	13.3	225.5	20.4	44.4	36.9	65.7	—0.61	—2.72	2.06
1958	3.7	12.2	301.6	29.1	47.9	42.9	61.1	—0.02	4.15	0.04
1959	5.1	32.1	385.4	24.5	50.5	48.0	70.6	—1.43	7.24	1.20
1960	6.6	25.6	540.3	33.8	55.4	52.6	74.1	1.00	8.29	1.97
1961	9.1	31.5	781.0	36.9	61.1	58.2	79.7	3.97	10.30	5.23
1962	13.8	41.6	1.213.5	44.0	64.4	62.8	84.0	2.25	0.67	6.55
1963	24.2	56.2	1.926.8	46.3	65.5	62.9	84.9	—2.77	1.73	3.54
1964	46.1	64.4	3.560.6	61.4	67.3	66.1	86.0	—7.00	—1.37	0.85
1965	72.3	45.0	6 533.3	60.7	69.1	63.0	97.9	—11.10	—14.30	9.75
1966	99.8	32.3	8 842.7	30.3	71.7	69.2	83.6	—14.30	—13.10	—10.10
1967	128.0	21.9	12 098.8	31.4	75.2	71.3	91.3	—10.40	—18.20	—5.27
1968	159.0	21.7	17 086.0	34.5	83.6	80.8	95.4	—12.00	—13.80	—1.00
1969	192.0	18.9	22 598.0	28.0	91.9	90.6	99.0	—9.98	—10.50	—5.21
1970	230.0	18.1	29 054.2	25.1	100.0	100.0	100.0	—8.36	—8.75	—8.23
1971	277.0	18.6	37 904.6	26.6	112.3	114.3	114.4	—2.71	—3.51	—1.45
1972	324.0	15.7	49 653.2	27.0	126.6	129.6	110.0	1.55	0.93	—1.42
1973	373.0	14.1	72 953.6	38.5	144.2	150.1	120.1	7.73	7.49	—1.97
1974	480.0	25.2	100 920.8	32.5	158.3	164.9	130.3	10.20	8.77	2.16
1975	613.0	24.5	134 245.3	28.5	167.3	175.2	134.7	8.92	6.70	1.46
1976	866.0	34.0	188 872.3	34.1	182.3	193.9	140.3	10.70	8.71	1.52
1977	1 236.0	35.0	260 169.5	32.0	190.8	201.4	153.8	8.39	4.38	6.68
1978	1 714.0	32.7	365 247.3	33.9	202.2	217.7	151.2	7.36	4.04	0.96

SOURCES: Central Bank (*Bulletin*) and Getulio Vargas Foundation (*Conjuntura Econômica*).

NOTATION: P: General Price Level Index (1965/67 = 100) (annual average); π Rate of Inflation measured as 100. log (P/P_{t-1}); M: Money Supply (Cr$ million) (annual average); μ: Rate of Money Growth measured as 100. log (M/M_{t-1}); y: Real Output Index (1970 = 100); — Real GDP; y^I: Industrial Output Index (1970 = 100); y^A: Agricultural Output Index (1970 = 100); h: 100. log (y/y^*) where log y^* is derived from a trend regression (slope: 0.0684); h^I: 100. log (y^I/y^{I*}) where log y^{I*} is derived from a trend regression (slope: 0.0813); h^A: 100. log (y^A/y^{A*}) where log y^{A*} is derived from a trend regression (slope: 0.0402).

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Real output – Inflation trade-offs, monetary growth and rational expectations in Brazil – 1950/79 *

Antonio Carlos Lemgruber **

1 — Introduction

Inflation and recession are certainly among the main themes of modern macroeconomics. As is well known, the recent economic literature has concentrated efforts to explain simultaneously inflationary and recessionary movements in market economies. The purpose of this paper is to contribute to the study of the inflation-recession dillema, by making use of data from Brazil for the period 1950/79. The Brazilian case is particularly interesting because of its high and variable inflation rates, ranging from 11% in 1950 to 90% in 1964 – with an average of 30%. In the same period, the rate of real growth has varied between 14% (1973) and 1.8% (1963) for an average of 7.2%, and the so-called real output "gap" varied between — 16% (1967) and + 11% (1976) for a zero average. [1]

In order to undertake this analysis of the relations between inflation and real output, as well as to study the effects of economic

* Originally published in *Revista Brasileira de Economia,* vol. 34, n.º 4, October/December, 1980.

** Fundação Getulio Vargas, Rio de Janeiro.

[1] Output gap here has the meaning of the relative deviation of actual output from normal output. For further details, see Sections 2 and 4. Notice that the numbers for inflation and growth here refer to percentage rates, in contrast to logarithmic rates used in the rest of the paper. Percentage rates are the more usual measures, but logarithmic rates present some properties that make them preferable for theoretical analysis (in particular, symmetry). The zero average for output gap is obviously a consequence of the fact that the measure of the gap is based on the residual of a least-squares trend regression. See Section 4 also.

policy on both, we will try to apply some recent models and ideas of modern macroeconomics. More specifically, we will consider two small-scale macromodels (Laidler-Parkin and Lucas), and we will also take into consideration rational expectations theory and some hipotheses related to downward inflation rigidity.

In Section 2, we consider the Laidler-Parkin model and the Lucas model. Both models concentrate their attention on inflation-real output trade-offs. In Section 3, we present a rapid review of rational expectations and discuss some of the criticisms that have been made against this approach, including the hypothesis of downward inflation rigidity, due to contracts and customer markets. In the next Section 4, estimates of the Laidler-Parkin model and the Lucas model are presented for Brazil (1950/79) with annual data. Finally, in Section 5, the inflation rigidity hypothesis is also tested for Brazil. A summary and some conclusions can be found in Section 6.

2 — Two popular modern models of inflation

In order to organize our discussion about real output-inflation trade-offs in Brazil, let us consider in this section the workings of two well-known models of inflation: the Laidler-Parkin model (hereafter, LP model) and the Lucas model (hereafter, L model) Both models are very good examples of one of the main purposes of modern macroeconomics, that is, to study inflation within the context of simple but complete macromodels that concentrate their attention on the problem generally characterized by Friedman's so-called "missing equation".[2] In Friedman's words, "the key need to remedy the defects common to (standard macroeconomic) models ... is a theory that will explain ... the short-run division of a change in nominal income between prices and output ...".[3]

Let us start by reviewing the LP model,[4] composed of 6 equations related to the endogenous variables h, Δh, ΔP, ΔP^*, Δy and Δy^*, that is, the level of excess demand measured by a "real output gap", the change in excess demand, the rate of inflation, the expected rate of inflation, the rate of real growth, and the rate of potential real growth.

One should notice that in one of the equations below h is implicitly being defined as $y - y^*$, or the difference between the log of real output and the log of potential or normal output —

[2] See Friedman (1971).

[3] *Ibid.*, p. 48.

[4] See Laidler and Parkin (1975) and Laidler (1975).

real output gap measure. Moreover, P is the log of the price level and P^* is the log of the expected price level. Notice also that, by definition, Δh equals $h - h_{t-1}$.

The model can be written as follows:

$$\Delta P = ah + \Delta P^* \tag{1}$$

$$\Delta P^* = b\Delta P_{t-1} + (1 - b)\,\Delta P^*_{t-1} \tag{2}$$

$$\Delta h = c\,(\Delta M - \Delta P) \tag{3}$$

$$h = \Delta h + h_{t-1} \tag{4}$$

$$\Delta y = \Delta y^* + \Delta h \tag{5}$$

$$\Delta y^* = \beta \tag{6}$$

where a, b and c are positive parameters (with b less or equal to 1), ΔM is the rate of monetary expansion, and β is a constant potential real growth rate. [5]

Equation (1) is a price formation equation with emphasis on excess demand and expectations, equation (2) is an adaptive expectations hypothesis, [6] and equation (3) is a simplified aggregate demand formulation. The other three equations have mainly a definitional content.

It will prove useful to reduce the model to 2 equations:

$$h = c\,(\Delta M - \Delta P) + h_{t-1} \tag{7}$$

$$\Delta P = \Delta P_{t-1} + ah - a\,(1 - b)\,h_{t-1} \tag{8}$$

Equation (7) — related to the aggregate demand side — implies a *negative* relationship between h and ΔP, and it contains two predetermined variables: ΔM and h_{t-1}. It could also be written as

$$\Delta h = c\,(\Delta M - \Delta P) \tag{7a}$$

Equation (8) corresponds to a "modern" Phillips Curve of the accelerationist type, leading to a *positive* relation between

[5] In the original formulation, some of these equations were only implicitly taken into account, but not explicitly. Another difference between the model composed of equations (1) to (6) and the original formulation relates to the presence of a lagged h in the price equation instead of contemporaneous h, but our preference for h is based purely on empirical reasons.

[6] But see the next section on rational expectations.

excess demand and inflation, and containing the following predetermined variables: ΔP_{t-1} and h_{t-1}. It could also be written as

$$\Delta^2 P = ah - a(1-b)\ h_{t-1} \tag{8a}$$

where $\Delta^2 P = \Delta P - \Delta P_{t-1}$, represents the acceleration of inflation or the second derivative of the log of the price level.

The reduced forms of this model are very significant, with ΔM, ΔP_{t-1} and h_{t-1} determining variables such as ΔP, $\Delta^2 P$, h and Δh. Let us show the reduced form coefficients for ΔP and h only:

$$\Delta P = \frac{1}{1+ac}\ \Delta P_{t-1} + \frac{ac}{1+ac}\ \Delta M + \frac{ab}{1+ac}\ h_{t-1} \tag{9}$$

$$h = \frac{-c}{1+ac}\ \Delta P_{t-1} + \frac{c}{1+ac}\ \Delta M + \frac{1+ac(1-b)}{1+ac}\ h_{t-1} \tag{10}$$

We will return to analyze these reduced forms of the LP model just after the presentation of the Lucas model (or L model).

Using the same notation as before, the formulation of the L model would be as follows: [7]

$$h = \gamma\ (P - P^*) + \lambda h_{t-1} \tag{11}$$

$$P = Y - y \tag{12}$$

$$P^* = E\ (P/I) \tag{13}$$

$$Y^* = Y_{-1} + \alpha \tag{14}$$

$$Y = Y^* + u \tag{15}$$

$$y = h + y^* \tag{16}$$

$$y^* = \delta + \beta T \tag{17}$$

Equation (11) is the well-known Lucas' aggregate supply formulation and it could also be written as:

$$y = y^* + \gamma(P - P^*) + \lambda(y_{t-1} - y^*_{t-1}) \tag{11a}$$

Equations (12) and (16) are definitions, but (12) should be regarded as aggregate demand for Y given. Equation (13)

[7] See Lucas (1972 and 1973).

embodies the rational expectation hypothesis,[8] where I represents the available information set at period t. In (14), one finds a very simple hypothesis about policy behavior. Notice that Y is log of nominal output and Y^* is predicted log of nominal output. People expect ΔY^* (the predicted rate of change of nominal output) to be on average equal to α.[9] But, according to (15), Y will differ from Y^* (or ΔY will differ from ΔY^*) by u. Therefore, $\Delta Y = \alpha + u$ – an anticipated and an unanticipated component. Finally, (17) identifies potential real output with a trend value. Parameters γ and λ are positive, with $\lambda < 1$.

It is certainly useful to attempt to present the L model with the same endogenous variables as the L-P model. We have then:

$$h = \gamma\,(\Delta P - \Delta P^*) + \lambda h_{t-1} \tag{18}$$

$$\Delta P = \alpha + u - \Delta y \tag{19}$$

$$\Delta P^* = E\,(P/I) - P_{t-1} \tag{20}$$

$$\Delta y = \Delta h + \Delta y^* \tag{21}$$

$$\Delta h = h - h_{t-1} \tag{22}$$

$$\Delta y^* = \beta \tag{23}$$

The aggregate supply equation (18) is the counterpart of the L-P Phillips Curve. Equation (19) becomes an aggregate demand relation, where ΔY is simply disaggregated between anticipated and unanticipated components. Moreover, (20) replaces the adaptive expectations formulation for a rational one.[10]

Let us now reduce the L model to 2 equations:

$$h = \gamma\,(\Delta P - \Delta P^*) + \lambda h_{t-1} \tag{24}$$

$$\Delta P = \Delta Y^* + (\Delta Y - \Delta Y^*) - (h - h_{t-1}) - \beta \tag{25}$$

with ΔP^* for the time being taken as exogenous (we are therefore neglecting (20) for a while).

The simultaneity between h and ΔP is similar to the one in the L-P model, with a negative link in (25) – *aggregate demand* – and a positive link in (24) – *aggregate supply*.

8 See the next section.

9 Notice that $\Delta Y^* = Y^* - Y_{t-1}$.

10 It must be emphasized that $P - P^* = \Delta P - \Delta P^*$, since $\Delta P^* = P^* - P_{t-1}$. This link between inflation expectations and price expectations is the only logical one.

For the sake of comparison with the L-P model, it will be interesting to consider the "non-rational" hypothesis $\Delta P^* = \Delta P_{t-1}$. Let us call the model formed by (24) and (25) with $\Delta P^* = \Delta P_{t-1}$ as the "L1 model".

But rationality implies that

$$\Delta P - \Delta P^* = \Delta Y - \Delta Y^* = \Delta Y - \alpha = u \tag{26}$$

this is derived from a "reduced form" for ΔP deduced from (24) and (25) with the hypothesis that $\Delta P^* = E(\Delta P)$. Therefore, one could have the "L2 model":

$$h = \gamma (\Delta Y - \Delta Y^*) + \lambda h_{t-1} \tag{27}$$

$$\Delta P = \Delta Y - (h - h_{t-1}) - \beta \tag{28}$$

All these models were estimated for Brazil, but with some minor modifications. For example, the excess demand variable that was used was related to the industrial sector (h^I), and in consequence we needed complementary equations to link h^I and h as well as Δh^I and Δh. Another modification was the inclusion of constants terms in some equations as well as a supply shock variable (measured by the deviation of agricultural output from trend, that is, h^A). Moreover, some additional lagged variables such as $(\Delta M - \Delta P)_{t-1}$ were introduced. Furthermore, in order to make the LP model and the L1 and L2 models more similar, we have assumed that it is possible and useful to replace ΔY by ΔM in the L models. As a matter of fact, the subsequent literature on rational expectations suggests that such a replacement is quite natural, if one is interested in studying policy effects.[11]

As a consequence of this last modification, of course, one should now regard α as the anticipated part of *money growth* and u as the unanticipated part, with $\Delta M = \alpha + u$. Let us further simplify the analysis and assume that $\alpha = \Delta M_{t-1}$. It could be argued that this hypothesis is not entirely rational in a context of strong monetary acceleration (or deceleration), but it seems to be a reasonably good empirical approximation for Lucas' hypothesis, where α corresponds to the "average rate of demand expansion".[12]

11 See Barro (1977) and Gordon (1977).

12 See Lucas (1973, p. 330).

The new formulations of the Lucas models would then be:

L1

$$h = \gamma\,(\Delta P - \Delta P_{t-1}) + \lambda h_{t-1} \tag{29}$$

$$\Delta P = \Delta M - (h - h_{t-1}) - \beta \tag{30}$$

L2

$$h = \gamma\,(\Delta M - \Delta M_{t-1}) + \lambda h_{t-1} \tag{31}$$

$$\Delta P = \Delta M - (h - h_{t-1}) - \beta \tag{32}$$

In the L1 model, reduced forms would include the same predetermined variables as in the LP model, that is, h_{t-1}, ΔP_{t-1}, and ΔM. In the L2 model, on the other hand, reduced forms would relate h and ΔP to ΔM, ΔM_{t-1}, and h_{t-1}. For example, here are the reduced forms for real output detrended (or h) in both models:

L1

$$h = \frac{\gamma}{1+\gamma}\,\Delta M - \frac{\gamma}{1+\gamma}\,\Delta P_{t-1} + \\ + \frac{\gamma+\lambda}{1+\gamma}\,h_{t-1} - \frac{\gamma}{1+\gamma}\,\beta \tag{33}$$

L2

$$h = \gamma\,\Delta M - \gamma\,\Delta M_{t-1} + \lambda\,h_{t-1} \tag{34}$$

In the following table, there is a complete list of short-run and long-run multiplier effects of ΔM on ΔP and h: [13]

		L-P Model	*L1 Model*	*L2 Model*
$\Delta P/\Delta M$	*S-R*	$ac/(1+ac)$	$1/(1+\gamma)$	$1-\gamma$ *or* 1
	L-R	1	1	1

(35)

[13] It is useful to recall here the relation between h and y. By definition, $h = \Delta y - \Delta y^{*} + h_{t-1}$ or $\Delta y = h - h_{t-1} + \Delta y^{*}$. Therefore, *in the short-run*, given $\Delta y^{*} = \beta$ and h_{t-1}, policy effects on h are equal to policy effects on Δy, the actual rate of real economic growth. The analysis in this paper could have been made all along with Δy instead of h as "the real variable", but we have chosen to stay close to the Phillips Curve tradition.

$$h/\Delta M \quad \begin{array}{lccc} S\text{-}R & c/(1+ac) & \gamma/(1+\gamma) & \gamma \text{ or } 0 \\ L\text{-}R & 0 & 0 & 0 \end{array} \tag{36}$$

All models present output effects in the short-run and only price effects in the long-run. It should be said, however, that in the L2 case the positive short-run output effect is a consequence of unanticipated movements in money growth, with $\Delta M - \Delta M_{t-1} = u$.

There is a great similarity between the LP model and the L1 model. In fact, the expected signs are the same for all the predetermined variables in both reduced forms. If one investigates the *final forms* [14] of both models, one will find relations such as

$$\Delta P = F_1\ (\Delta P_{t-1}, \Delta P_{t-2}, \Delta M, \Delta M_{t-1}) \tag{37}$$

$$h = F_2\ (h_{t-1}, h_{t-2}, \Delta M, \Delta M_{t-1}) \tag{38}$$

that is, second-order difference equations. In the LP model, the characteristic equation or auxiliary equation would be $X^2 - [2 + ac\,(1-b)]/(1+ac)\ X + 1 = 0$ and in the L1 model, it would be $X^2 - (2\gamma + \lambda)/(1+\gamma)\ X + \gamma = 0$. It can be shown that these auxiliary equations generally contain complex roots and in consequence the models present oscillatory movements or cyclical properties, depending particularly on the values of a or γ as well as b or λ. [15]

If $\lambda = 0$ and $b = 1$, the LP model and the L1 model become even more similar models. In the empirical applications for Brazil, it will become clear that these simplifications are quite valid. In this case, one could write:

$$\Delta^2 P = a\ h \quad \text{or} \quad h = \gamma\ \Delta^2 P \tag{39}$$

and

$$\Delta h = c\,(\Delta M - \Delta P) \quad or \quad \Delta P = \Delta M - \frac{1}{c}\,\Delta h \tag{40}$$

14 The concept of a *final form* is useful for cyclical analysis. One eliminates lagged endogenous variables in a reduced form if they are predetermined by other reduced forms.

15 The L2 model has a final form corresponding to a first-order difference equation for h and ΔP, and in consequence it does not present cyclical properties.

The reduced forms of these simpler models continue to contain ΔM, ΔP_{t-1} and h_{t-1}, and the final forms continue to lead to a second-order difference equation with ΔM and ΔM_{t-1} as the *moving* variables. There are now two basic parameters: a or $\gamma = 1/a$, and c. If c is supposed to be unitary, then it becomes clear that the overall cycle of ΔP and h is basically guided by a (or γ), that is, the slope of the short-run Phillips Curve.

One could say that models such as the ones presented above are the 70's counterpart of the well-known Samuelson multiplier — accelerator model. Forty years ago, Samuelson developed from the multiplier and the accelerator a second order difference equation to explain nominal income fluctuations due to autonomous expenditures.[16] During the 70's, Laidler, Parkin, and Lucas produced models that *simultaneously* explain prices and real output with second-order difference equations that are moved by monetary (as well as other policy) effects.

After reviewing rational expectations in the next section, we shall present estimates of the following equations for Brazil:

1) Laidler-Parkin Price Equation — see (1), (8) and (8a).

2) Laidler-Parkin Aggregate Demand Equation — see (3), (7) and (7a).

3) Complementary Equations.

4) Laidler-Parkin Reduced Forms[17] — see (9), (10), (33), (35) and (36).

5) Lucas Type 1 Aggregate Supply Equation — see (11), (18), (24) and (29).

6) Lucas Type 2 Aggregate Supply Equation — see (11), (18), (24), (27) and (31).

7) Lucas Aggregate Demand Equation — see (12), (25), (28), (30) and (32).

8) Lucas Type 2 Reduced Forms — see (34), (35) and (36).

Furthermore, we shall consider in Section 5 the "inflation rigidity hypothesis" and shall present tests for it using the L1 and L2 aggregate supply formulations.

16 See Samuelson (1939).

17 Also valid for L1 reduced forms.

3 — Rational expectations and inflation rigidity

At least up until the early seventies, the economic literature used to present dozens of examples of expectation hypotheses about some variable based on the past history of that variable. As far as inflation (or the price level) is concerned, one used to find so-called naive models, classical models, adaptive models, regressive models, and extrapolative models — all based on past inflation. One good example is precisely the L-P adaptive hypothesis shown in the previous section:

$$\Delta P^* = b \, \Delta P_{t-1} + (1 - b) \, \Delta P^*_{t-1} \tag{41}$$

Such models were applied either with respect to the price level or with respect to the rate of inflation, and they can be generalized in one of the following ways:

$$\Delta P^* = \alpha_1 \, \Delta P_{t-1} + \alpha_2 \, \Delta P_{t-2} + \ldots \tag{42}$$

$$P^* = \gamma_1 \, P_{t-1} + \gamma_2 \, P_{t-2} + \ldots \tag{43}$$

Essentially, these models suggest that expectations about inflation (or the price level) are based on past rates of inflation (or past price levels).

In 1961, Muth [18] presented some ideas about the formation of expectations which would become the basis of the "rational expectations" model:

> "Expectations, since they are informed predictions of future events, are essentially the same as the predictions of the relevant economic theory... Expectations of firms tend to be distributed, for the some information set, about the predictions of the theory." [19]

A model that does not incorporate such ideas should be considered as "irrational". Clearly, based on Muth's definition, those previous expectational models were not "rational". It becomes implicit that, in Muth's hypothesis, information is scarce and there is no waste of information, with expectations depending on the complete structure of the relevant model which is able to describe the economy.

18 See Muth (1961).

19 *Ibid.*, pp. 317-8.

Using supply and demand micro-models, Muth formulated his rational expectation theory. Considering, for example, expected prices in a micro framework one would have:

$$P_t^* = E\ (P_t) \tag{44}$$

where $E\,(P_t)$ is the reduced form prediction of the relevant supply-demand model derived from microeconomic theory.

Ten years later, in the early seventies, authors such as Lucas, Sargent, Barro and McCallum have begun to attempt to introduce rational expectations in macroeconomic models. [20] The Lucas model with rational expectations has already been presented in the previous section. Nowadays, the neutrality propositions derived from rational expectations are reasonably well-wnown: systematic measures of economic policy have no real effect in the economy; real variables are statistically independent from the systematic part of monetary and fiscal policy. Rational expectation models imply that only the non-systematic part of monetary and fiscal policy can have real effects in the economy.

Clearly, these policy — inefectiveness implications were very serious for short-run economic policies which tended to emphasize fine-tuning of the economy. Very rapidly, criticisms of rational expectation models begun to appear in the literature. For instance, one could mention papers by Modigliani, Gordon, Tobin, and Okun. [21]

The most common criticisms have to do with price stickyness or — carrying the question to the first derivative — with inflation rigidity. Price adjustments are supposedly slow and sluggish in the real world — particularly in the downward direction — due to the existence of noncompetitive markets, oligopolistic markets, "customer markets", and "contract markets". Price flexibility would be found only in auction markets — where there are no long-run contracts — such as commodities or financial markets.

In order to take into consideration the downward rigidity of inflation, one could think of an expectations model where, for example, the expected inflation would be equal to the greatest of the following two results: past inflation or inflation predicted by relevant economic theory, that is,

$$\Delta P^* = \max\ \{E\ (\Delta P),\ \Sigma W_i \Delta P_{t-i}\} \tag{45}$$

[20] See Lucas (1972 and 1973), Sargent (1973), Barro (1976) and McCallum (1979*a* and 1979*b*).

[21] See Modigliani (1977), Gordon (1977) and Tobin (1972*a*).

The public would not take into account "rational expectations" when predictions of economic theory suggested inflation rates inferior to past rates of inflation – at least not entirely.

When inflation is accelerating, then rational expectations would become relevant and inflation stickyness could be neglected. But when inflation is decelerating, there can be rigidities in the process that disturb the elegance of the rational expectations idea. In other words, rational expectations would be valid for accelerating inflation, but invalid – at least partially – for decelerating inflation. The past history of inflation would represent a "floor rate" that has to be taken into account in the process of expectation formation.

It seems that inflation rigidity plus rational expectations lead us to a very difficult real world: expansive policies would have practically no positive real output effects, but contractionary policies would continue to have negative real output effects. In fact, the above comments imply that accelerating money (or accelerating inflation) does not have real effects, but decelerating money (or decelerating inflation) does have a serious real side effect – unemployment and lower growth.

Clearly, this situation is almost dramatic for the policy-maker: it would be almost useless to attempt to expand the economy through policy, but it would continue to be very costly to contract the economy in terms of real effects on employment and output.

For example, as shown in the previous section, when one neglects rational expectations, there is a positive relation between Δ^2P and h or between Δ^2M and h implied by conventional inflation models with traditional hypotheses about expectations. On the other hand, the implication of rational expectations is that even these positive relations between second derivatives of the price level or the money supply and real variables might not exist if the accelerations of money and inflation are also anticipated. Thus, even in the short-run, and even for the second derivative, one would have a vertical Phillips Curve, that is, no trade-off between "anticipated" derivatives of the price level or the money supply and unemployment.

But, what would happen if we take into account the rational expectation thesis for accelerating inflation (positive second derivative of the price level or the money supply) but retain the more conventional model for decelerating inflation (negative second derivative of the price level a the money supply) due to inflation rigidity? In this case, we would obtain a Phillips Curve or aggregate supply relation that would be vertical for positive values of $\Delta^2 P$ or $\Delta^2 M$ and nonvertical for negative values — a kinked accelerationist Phillips Curve.[22]

This hypothesis of a kinked Phillips Curve — which combines inflation rigidity and rational expectations — has very significant implications for real world policy, and it will be tested for Brazil in this paper in Section 5, just after a formal presentation of estimates of the models presented in the previous sector.

4 — Estimates of the models for Brazil

In this section, we will present some estimates of the equations of the LP model and the L model for Brazil. The sample period is 1950/79 (30 observations) and the estimation method was OLS.[23] The basic annual data for our regressions are summarized in Table 9, including sources as well as notation.

The main regressions are collected in the following eight tables, whose results will be analyzed in the following paragraphs.

We shall start with the Laidler-Parkin model, amounting to a price equation, an aggregate demand equation, some complementary equations, and reduced forms. After that, we will move to the Lucas model — aggregate supply, aggregate demand and reduced forms.

22 In a less extreme result, the implication would be that for positive values of $\Delta^2 P$ and $\Delta^2 M$, aggregate supply would be "more vertical" and for negative values it would be "more horizontal". The real effects would be significant for negative values of $\Delta^2 P$ and $\Delta^2 M$, but practically negligible for positive values. It ought to be mentioned that this implication is derived simply from a mix of rational expectations and downward inflation rigidity. But Simonsen (1979) has shown recently that legal *indexation mechanisms* based on past inflation could also explain this type of assymetry. His point is that indexation contributes to make price and wage inflation even more rigid in the downward direction. Therefore, one should say that in countries where there are indexed contracts — generally based on past inflation — the problem of inflation rigidity becomes even more dramatic. Tobin (1972*b*) also discusses the existence of a floor rate of wage change, which could represent a barrier to the total absence of trade-offs in the long-run in an economy where relative wages are an important concern.

23 We are aware of the simultaneity problem of our models, which might lead to biases in OLS estimation. However, after some experimentation with 2SLS and even 3SLS estimation, we did not find significant differences in terms of estimated values, and consequently, for the sake of simplicity, we have preferred to report OLS results.

Table 1

Laidler-Parkin Price Equation
(Annual Data, 1950/79)

Explanatory Variables	Dependent Variable: $\Delta^2 P$		Dependent Variable: ΔP	
	I.1	I.2	I.3	I.4
Constant	1.1303	1.0994	3.3269	1.9890
	(0.89)	(0.87)	(1.18)	(0.58)
h^I	0.6330	0.7341	0.5836	0.6952
	(3.85)	(3.91)	(3.34)	(2.94)
h^A		−0.3701		−0.3004
		(−1.10)		(−0.71)
ΔP_{t-1}			0.9128	0.9649
			(9.09)	(7.71)
Statistics				
R^2	0.3461	0.3740	0.7539	0.7586
DW	2.4126	2.3446	2.2950	2.2993
SER	6.9172	6.8920	6.9479	7.0127

Table 2

Laidler-Parking Aggregate Demand Equation
(Annual Data, 1950/79)

Explanatory Variables	Dependent Variable: Δh		Dependent Variable: h	
	II.1	II.2	II.3	II.4
Constant	−1.4721	−1.4505	−1.4632	−1.4226
	(−2.41)	(−2.49)	(−2.33)	(−2.38)
$\Delta M-\Delta P$	0.2092	0.1784	0.2069	0.1705
	(3.27)	(2.83)	(3.02)	(2.51)
$(\Delta M-\Delta P)_{t-1}$	0.1152	0.1419	0.1152	0.1424
	(1.78)	(2.25)	(1.75)	(2.22)
h_{t-1}			0.9925	0.9767
			(14.62)	(14.99)
Δh^A		0.1693		0.1734
		(1.92)		(1.92)
Statistics				
R^2	0.3719	0.4498	0.8936	0.9072
DW	1.6993	1.9251	1.6843	1.8918
SER	2.5028	3.3870	2.5498	2.4280

Table 3

Complementary Equations
(Annual Data, 1950/79)

Explanatory Variables	Dependent Variable: Δh^{I} III.1	Dependent Variable: Δh^{A} III.2
Constant	—0.091	—0.0537
	(—0.20)	(—0.07)
Δh	1.1775	
	(7.43)	
Δh^{A}	—0.3859	
	(—4.22)	
h^{A}_{t-1}		—0.5495
		(—2.45)
Δh^{A}_{t-1}		—0.2516
		(—1.35)
Statistics		
R^2	0.6852	0.4079
DW	2.0997	2.0535
SER	2.4572	4.2097

Table 4

Laidler-Parkin Reduced Forms *
(Annual Data, 1950/79)

Explanatory Variables	Dependent Variable: WP IV.1	Dependent Variable: h IV.2	Dependent Variable: h^{I} IV.3	Dependent Variable: Δh IV.4
Constant	—0.0784	0.4565	3.0864	0.4565
	(—0.02)	(0.33)	(2.00)	(0.33)
ΔM	0.5166	0.1239	0.2277	0.1239
	(2.70)	(1.79)	(2.77)	(1.79)
ΔP_{t-1}	0.4113	—0.1683	—0.4020	—0.1683
	(2.39)	(—2.58)	(—5.46)	(—2.58)
h_{t-1}		0.9501		—0.0499
		(12.14)		(—0.64)
h^{I}_{t-1}	0.3631		0.6987	
	(1.91)		(8.58)	
h^{A}_{t-1}	0.2294	—0.2430	0.3953	—0.2430
	(0.65)	(—1.74)	(2.60)	(—1.74)
Statistics				
R^2	0.8081	0.9083	0.8943	0.4591
DW	2.3042	1.9840	1.8154	1.9840
SER	6.3769	2.4136	2.7352	2.4136

* Also Valid for Lucas type 1 Reduced Forms.

Table 5

Lucas Type 1 Aggregate Supply Equation
(Annual Data, 1950/79)

Explanatory Variables	Dependent Variable: h^I		
	V.1	V.2	V.3
Constant	—0.1832	—0.5369	—0.4336
	(—0.25)	(—0.45)	(—0.42)
$\Delta^2 P$	0.1979	0.5467	0.4918
	(1.96)	(3.85)	(3.91)
h^I_{t-1}	0.7473		
	(6.88)		
h^A			0.7435
			(3.06)
Statistics			
R^2	0.7623	0.3461	0.5145
DW	1.4856	0.9073	1.3506
SER	3.9467	6.4289	5.6411

Table 6

Lucas Type 2 Aggregate Supply Equation
(Annual Data, 1950/79)

Explanatory Variables	Dependent Variable: h^I	
	VI.1	VI.2
Constant	—0.2700	—0.6059
	(—0.22)	(—0.47)
$\Delta^2 M$	0.4105	0.3716
	(2.70)	(2.34)
$\Delta^2 M_{t-1}$		0.3343
		(2.16)
h^A_{t-1}	0.7877	
	(2.79)	
Statistics		
R^2	0.3356	0.2702
DW	0.6014	0.6969
SER	6.5990	6.9162

Table 7

Lucas Aggregate Demand Equation
(Annual Data, 1950/79)

Explanatory Variables	Dependent Variable: ΔP	
	VII.1	VII.2
Constant	—1.2216	—0.7577
	(—0.36)	(—0.19)
ΔM	0.8914	0.8726
	(8.57)	(7.31)
Δh	—1.4175	
	(—3.60)	
Δh^I		—0.8485
		(—2.56)
Δh^A		—0.4230
		(—1.74)
Statistics		
R^2	0.8006	0.7755
DW	2.1247	1.8511
SER	6.2539	6.7628

Table 8

Lucas Type 2 Reduced-Forms *
(Annual Data, 1950/79)

Explanatory Variables	Dependent Variable: ΔP		
	VIII.1	VIII.2	VIII.3
Constant	—4.6249	—2.6807	—3.6260
	(—1.28)	(—0.69)	(—0.90)
ΔM	1.0150	0.9441	0.9727
	(9.11)	(7.78)	(7.71)
$\Delta^2 M$	—0.3331	—0.2942	—0.2895
	(—2.10)	(—1.70)	(—1.50)
h_{t-1}	0.4972		
	(2.92)		
h^I_{t-1}		0.5092	0.3363
		(2.94)	(1.59)
h^A_{t-1}			0.3203
			(0.84)
Δh_{t-1}	—1.0337		
	(—2.54)		
Δh^I_{t-1}		—0.5011	
		(—1.53)	
Statistics			
R^2	0.8133	0.7965	0.7836
DW	1.7634	1.8313	1.6710
SER	6.2889	6.5663	6.7712

* For Lucas type 1 Reduced Forms, see IV (LP Reduced Forms).

Table 9

Brazil, Annual Data, 1947/79

Years	P	ΔP	M	ΔM	y	y^I	y^A	h	h^I	h^A
1947	0.86	—	43.8	—	23.53	16.40	42.47	4.15	—2.03	—1.40
1948	0.92	6.77	43.9	0.23	25.28	18.20	45.41	4.50	—0.06	1.28
1949	0.98	6.86	50.3	13.61	26.95	20.13	47.40	4.06	1.58	1.65
1950	1.00	10.54	61.4	19.94	28.70	22.40	48.17	3.52	4.16	—0.88
1951	1.27	15.28	77.0	22.64	30.40	23.83	48.50	2.43	2.22	—4.21
1952	1.42	11.16	88.3	13.69	33.04	25.02	52.91	3.03	—1.05	0.48
1953	1.63	13.79	104.2	16.56	33.88	27.19	53.01	—0.41	—0.84	—3.36
1954	2.07	23.90	128.4	20.88	37.30	29.65	57.18	2.38	—0.60	0.20
1955	2.41	15.21	152.9	17.46	39.66	32.09	61.60	2.19	1.31	3.61
1956	2.80	18.16	183.8	18.41	31.13	34.94	60.12	—1.52	—0.14	—2.82
1957	3.30	13.27	225.5	20.45	44.44	36.93	65.73	—0.61	—2.72	2.06
1958	3.73	12.25	301.6	29.08	47.80	42.01	67.05	—0.02	4.15	0.04
1959	5.14	32.06	385.4	24.52	50.53	48.01	70.51	—1.43	7.24	1.20
1960	6.64	25.61	540.3	33.78	55.44	52.61	74.08	1.00	8.29	1.97
1961	9.10	31.52	781.0	36.92	61.15	58.21	79.08	3.97	10.30	5.23
1962	13.8	41.64	1 213.5	43.99	64.36	62.76	84.04	2.25	9.67	6.55
1963	24.2	56.17	1 926.8	46.24	65.54	62.88	84.90	—2.77	1.73	3.54
1964	46.1	64.45	3 560.6	61.41	67.27	66.12	86.04	—7.00	—1.37	0.85
1965	72.3	45.00	6 533.3	60.70	69.10	63.00	97.90	—11.10	—14.30	9.75
1966	99.8	32.33	8 842.7	30.27	71.70	69.20	83.60	—14.30	—13.10	—10.10
1967	128.0	24.89	12 098.8	31.35	75.20	71.30	91.30	—16.40	—18.20	—5.27
1968	159.0	21.69	17 086.0	34.52	83.00	80.80	95.40	—12.60	—13.80	—4.90
1969	192.0	18.86	22 598.0	27.96	91.90	90.60	99.00	—9.98	—10.60	—5.21
1970	230.0	18.00	29 054.2	25.13	100.00	100.00	100.00	—8.36	—9.75	—8.23
1971	277.0	18.59	37 904.0	26.50	113.30	114.30	111.40	—2.71	—3.51	—1.45
1972	324.0	15.07	49 653.2	27.00	126.60	129.60	116.0	1.55	0.93	—1.42
1973	373.0	14.08	72 953.6	38.48	144.20	150.10	120.10	7.73	7.49	—1.97
1974	480.0	25.22	100 926.8	32.46	158.30	164.90	130.30	10.20	8.77	2.16
1975	613.0	24.46	134 245.3	28.63	167.30	175.20	134.70	8.92	6.70	1.46
1976	866.0	34.55	188 872.3	34.14	182.30	193.90	140.30	10.70	8.71	1.52
1977	1 236.0	35.58	260 109.5	32.03	190.80	201.40	153.80	8.39	4.38	6.08
1978	1 714.0	32.69	365 198.0	33.91	202.20	217.70	151.20	7.30	4.04	0.96
1979	2 638.0	43.12	557 130.3	42.24	215.10	232.70	156.00	6.70	2.57	0.08

SOURCES: Central Bank (*Bulletin*) and Getulio Vargas Foundation (*Conjuntura Economica*).

NOTATION: P = General Price Level Index (1965-1967 = 100) (annual average).
ΔP = Rate of Inflation measured as 100. log (P/P_{t-1}).
M = Money supply (Cr$ million) (annual average).
ΔM = Rate of Money Growth measured as 100. log (M/M_{t-1}).
y = Real Outup Index (1970 = 100) — Real GDP.
y^I = Industrial Outup Index (1970 = 100).
y^A = Agricultural Output Index (1970 = 100).
h = 100. $\log(y/y^*)$ where log y^* is derived from a trend regression (slope: 0,0684).
h^I = 100.$\log(y^I/y^{I*})$ where log y^{I*} is derived from a trend regression (slope: 0,0813).
h^A = 100.$\log(y^A/y^{A*})$ where log y^{A*} is derived from a trend regression (slope: 0,0402).

NOTES:

1) $\Delta^2 P = \Delta P - \Delta P_{t-1}$

$\Delta h = h - h_{t-1}$

$\Delta h^I = h^I - h^I_{t-1}$

$\Delta h^A = h^A - h^A_{t-1}$

$\Delta^2 M = \Delta M - \Delta M_{t-1}$

$\Delta y = 100. \log (y/y_{t-1})$

$\Delta y^I = 100. \log (y^I/y^I_{t-1})$

$\Delta h^A = 100. \log (y^A/y^I_{t-1})$

$\Delta y = Wh + 6{,}84$

$\Delta y^I = Wh^I + 8{,}13$

$\Delta y^A = Wh^A + 4{,}02$

$\Delta y^* = 100. \log (y^*/y^*_{t-1}) = 6{,}84$

$\Delta y^{I*} = 100. \log (y^{I*}/y^{I*}_{t-1}) = 8{,}13$

$\Delta y^{A*} = 100. \log (y^{A*}/y^{A*}_{t-1}) = 4{,}02$

2) Notice that here in this Table 0 the symbols P and M are used for the price level and the money supply, and *not* the *logs* of the price level and the money supply as in the rest of the text. The same is true for y, y^I, y^A as well as y^*, y^{I*}, and y^{A*}. Notice also that we are working with logarithmic rates of change for ΔP, ΔM and Δy instead of percentage rates of change (but see footnote 1). With logarithmic changes, there are at least two advantages: a) symmetry between negative and positive changes; b) the connection between h and Δy or Δh and Δy is precise, and not merely an approximation.

L-P Price Equation

The basic price equation of the L-P model, as shown in Section 2, can be written as:

$$\Delta^2 P = ah - a(1 - b)\ h_{t-1}$$

In the estimation process, we have considered this formulation, but preliminary results have led us to introduce the following modifications: a) assume $b = 1$, which corresponds to a simple expectation hypothesis where $\Delta P^* = \Delta P_{t-1}$; b) replace h — representing total output — by h^I which refers to industrial output; c) take into consideration a supply shock variable — measured by agricultural output detrended — as a possible inflationary effect in addition to excess demand and price expectations; d) inclusion of a constant term in the equation. The assumption $b = 1$ as well as the use of h^I were simply consequences of empirical results, to the extent that regressions with h^I only were much better then regressions with h^I and h_{t-1} on one hand and much better than regressions with total h on the other hand. The inclusion of a supply shock variable — as well as a constant term — was found to be necessary in order to capture possible autonomous effects on inflation.

The first two regressions presented in Table 1 are quite satisfactory, revealing a strong Phillips Curve effect in Brazil, as well as a correct but non-significant sign for the supply shock. The last two results allow the coefficient of ΔP_{t-1} to differ from one, but the results continue to confirm the accelerationist — type Phillips Curve for Brazil, with good R^2 and satisfactory DW statistics.

L-P Aggregate Demand Equation

The basic relation from Section 2 is

$$\Delta h = c\ (\Delta M - \Delta P)$$

which can also be written as

$$\Delta y = \beta + c\ (\Delta M - \Delta P)$$

that is, deviations between real and potential growth are explained by the rate of change of real balances — a monetarist aggregate demand formulation.

Empirically, we have decided to add an additional lagged effect of $(\Delta M - \Delta P)$ on Δh and have considered a constant term.

Moreover, we have also considered the possibility that the demand side would affect more intensely Δh^I and, in consequence, the second and fourth regressions of Table 2 include Δh^A as an additional variable, since $\Delta h = F\ (\Delta h^I, \Delta h^A)$. [24] Furthermore, in two regressions, we have relaxed the restrictions related to a unitary coefficient for h_{t-1}, by changing the dependent variable and taking h_{t-1} to the right-hand side.

Once more, the results come out quite favorable. For example, the positive effects of real money growth in t and $t-1$ are statistically significant, with a total sum greater than 0.3. This figure should not be regarded simply as the inverse of an income-elasticity of the demand for money around 3, since it can be argued that the L-P aggregate demand model is more than a simple quantity theory formulation. [25] The effect of agricultural growth ($\Delta h^A = \Delta y^A - \Delta y^{*A}$) also contains some significance, contributing to improve R^2 and DW statistics – and this could be an indication that the real money growth effect is more concentrated on the manufaturing sector. Moreover, h_{t-1} appears with a coefficient not significantly different from 1 in the last two regressions, as expected.

Finally, the constant terms are significant, with an interesting interpretation: when real money growth is zero, since $\Delta h = \Delta y - \beta$, the difference between the actual rate of real growth and the potential rate is -1.5. As $\beta \cong 7.0$, zero real money growth would imply an actual growth around 5.5%. Obviously, the "financing" of such real growth would have to be made by a growing income-velocity of money.

Complementary Equations

The use of h, h^I and h^A in our estimation forces us to establish some interrelations among them as well as to attempt to understand more specifically the shock movements of h^A (or Δh^A).

In a previous footnote, it has already been pointed out that $\Delta h = F\ (\Delta h^I, \Delta h^A)$ since $\Delta y = F\ (\Delta y^I, \Delta y^A)$. As a matter of fact, assuming that the services sector also depends on industry and agriculture, one can practically say that Δy is a weighted average of Δy^I and Δy^A. As the LP Price Equation considers h^I

[24] This relationship is obvious if one recalls that, except for constant terms, it is equivalent to $\Delta y = F\ (\Delta y^I, \Delta y^A)$, that is, the rate of growth of total real output is a function of industrial growth and agricultural growth.

[25] For example, one can say that the underlying model behind the *L-P* aggregate demand is an *IS-LM* model with ΔM being used as a proxy for both monetary and fiscal effects. One should also check the numerical example in the final subsection of Section 4 (*Small Summary*).

and the LP aggregate demand equation considers h (or Δh) — and this also occurs in the L1 and L2 models — one needs a simple equation to link Y^r and h or Δh^I and Δh. Such equation is the first regression of Table 3. If it were rewritten with Δh on the left-hand side, the regression would simply imply that

$$\Delta h = 0.85 \, \Delta h_I + 0.33 \, \Delta h_A.$$

The other complementary equation in Table 3 serves only to confirm the cobweb movements of agriculture. Agricultural growth goes beyond the average ($\Delta h^A > 0$) whenever we start from a bad crop in the previous year (negative h^A_{t-1}).

The R^2 and DW statistics for these two equations are satisfactory.

L-P Reduced Forms

We turn now to L-P reduced forms. Lt us recall that, as pointed out in Section 2, such reduced forms are also valid for the L1 model. Basically, the theoretical variables to be considered are ΔM, ΔP_{t-1}, and h_{t-1}. However, due to the modifications introduced in the models, we have also considered constant terms and the lagged values of h^I_{t-1} and h^A_{t-1}.

Table 4 presents some selected reduced forms, which should be considered valid for both the LP model and the L1 model.

First of all, one must emphasize that all estimated coefficients have the expected signs and the level of significance seems to be quite satisfatory. The reduced forms explain 81% of inflation variability, 89-91% of output variability (already detrended), as well as 46% of the variability in the actual rate of real growth (notice that $\Delta h = \Delta y + \beta$).[26] There is no relevant autocorrelation problem.

The short-run effects or short-run multipliers of ΔM are 0.52 on prices and 0.12 on real output and 0.23 on industrial output. It can be shown that the long-run multipliers are close to one for prices, and close to zero for real and industrial output.[27] These numbers reflect to a great extent the implications of (35) and (36) in Section 2, even through they to not lead precisely to zero or unitary long-run multipliers. This imprecision must

[26] Notice that the second and fourth reduced forms are basically the same, except for a change in the dependent variable which affects R^2 and the coefficient of h_{t-1}, but does not affect other coefficients and other statistics.

[27] These long-run multipliers are obtained with $\Delta P = \Delta P_{t-1}$, $h = h_{t-1}$, and $h^I = h_{t-1}$.

be attributed to the non-imposition of restrictions in the estimated reduced forms as, for example, the inclusion of $(\Delta M - \Delta P_{t-1})$ in the last three reduced forms. With these restrictions, multipliers for real variables of ΔM are zero – only the second derivative or $\Delta^2 M$ has real effects.

Lucas Type 1 Aggregate Supply

Lucas basic formulation is:

$$h = \gamma\ (\Delta P - \Delta P_{t-1}) + \lambda h_{t-1} = \gamma\ \Delta^2 P + \lambda h_{t-1}$$

With h^I instead of h, Table 5 presents such an aggregate supply relation. The short-run coefficient of the Phillips Curve between h and $\Delta^2 P$ is 0.20 but the long-run one is $0.20/1 - 0.75) = 0.80$.

The presence of a lagged h^I on the right-hand side is the main aspect that differentiates such a model from the L-P Price Equation. When one assumes $\lambda = 0$, we have the relation:

$$h^I = -0.5369 + 0.5467\ \Delta^2 P$$

which is equivalent to the L-P Price Equation:

$$\Delta^2 P = 0.1303 + 0.6330\ h^I$$

The simultaneity question is obvious in this case, and one finds no compelling reason to emphasize either one or the other variable on the left-hand side. The point is that we have one equation (aggregate supply or Phillips Curve) and two basic macroeconomic variables — inflation acceleration and industrial output detrended as a proxy for unemployment. One needs aggregate demand to close the model.

Finally, the introduction of a supply shock variable is particularly significant here in the L1 aggregate supply model. In fact, considering econometric problems of a lagged dependent variable and the low DW of the simple relation between h^I and $\Delta^2 P$, the third regression of Table 5 seems to be the most appropriate, characterizing: 1) the positive price effect on aggregate supply; 2) agricultural supply shocks affecting positively the industrial movements, as one should expect.

Lucas Type 2 Aggregate Supply

Consider the basic equation, mixing aggregate supply and rational expectations:

$$h = \gamma\ (\Delta M - \Delta M_{t-1}) + \lambda h_{t-1} = \gamma\ \Delta^2 M + \lambda h_{t-1}$$

Empirically, the L2 Aggregate Supply had to be slightly changed after some experimentation with Brazilian data. A constant term was added. An additional lagged term for $\Delta^2 M$ was introduced. The variable h was replaced by h^I. A supply shock variable was also included.[28] The lagged dependent variable was suppressed.

In spite of lower R^2 and DW as compared to the L1 formulation, the important aspect to be emphasized is the adequate level of significance for the coefficients of monetary acceleration. The estimation tends to confirm the basic positive relation between $\Delta^2 M$ (and $\Delta^2 M_{t-1}$) and h^I, with coefficients which sum up to 0.7 and t-scores greater than 2. It is true that R^2 and DW are low, but the inclusion of a lagged dependent variable was found to be impractical, in spite of obvious but meaningless improvements in the statistics. In fact, with the lagged dependent variable, the other explanatory variables lose entirely their significance.

Anyway, the two regressions of Table 6 exemplify reasonably well the implications of the L2 model as far as aggregate supply is concerned. We will return to them when we discuss the inflation rigidity hypothesis in Section 5.

Lucas Aggregate Demand

When we take into consideration ΔM instead of ΔY for usefull aggregate demand analysis — as we did in Section 2 — then we obtain this simple aggregate demand formulation for Lucas model:

$$\Delta P = \Delta M - \Delta h - \beta$$

or

$$\Delta P = \Delta M - \Delta y$$

Comparing with the L-P model the main difference is the implicit unitary coefficient for Δy or Δh, besides the exchange of ΔP from the right-hand to the left-hand side and vice-versa for Δy.[29]

28 Since the L2 Aggregate Supply equation corresponds already to a reduced form for h (or h^I), with only predetermined variables on the right-hand side, we have decided to consider h^A_{t-1} rather than h^A as the supply shock variable. Naturally, according to the complementary equations in Table 3, h^A_{t-1} is the main determinant of movements in h^A.

29 Another minor difference is the presence of β in the L equation, but such presence could also have been introduced in the LP model since the very beginning — see also Laidler (1975).

For estimation purposes, we have relaxed both the unitary coefficients for ΔM and Δh (or Δy). The results – quite satisfatory – are in Table 7. The most simple one is:

$$\Delta P = -1.2216 + 0.8914\ \Delta M - 1.4175\ \Delta h$$

with very significant t-scores for ΔM and Δh, as well as good statistics.

An in the case of the LP model, one is tempted to think in terms of an inverted money demand in the quantity theory tradition, with an income-elasticity of money demand greater than 1. It is not incorrect to interpret our formulation of Lucas aggregate demand in this fashion, but it must be emphasized that the underlying model could go beyond the quantity theory.[30]

Finally, the other regression in Table 7 includes separately Δh^I and Δh^A instead of Δh. It can be observed that both signs are correct, and the coefficient and t-score for industrial growth is greater than the ones for agricultural growth. Moreover, the statistics for aggregate demand with Δh^I and Δh^A are slightly worse than the ones for Δh only.

Lucas Type 2 Reduced Forms[31]

Finally, we comment on Lucas Type 2 Reduced Form for ΔP. It must be recalled that, as far as the real variable h^I (or Δh^I) is concerned, the L2 aggregate supply formulation *is* already a reduced from.

Formally, the analysis of Section 2 indicated that ΔM, ΔM_{t-1} and h_{t-1}, should be in the reduced form for ΔP, that is,

$$\Delta P = \Delta M - \gamma\ (\Delta M - \Delta M_{t-1}) + (1 - \lambda)\ h_{t-1} - \beta$$

An interesting result is related to the short-run negative effect of $\Delta^2 M$ on ΔP, as well as the unitary positive effect of ΔM.

Empirically, we have relaxed the unitary constraint for ΔM and have introduced additional explanatory variables in the reduced forms – basically lagged values of h_I and h_A as well as lagged values of Δh and Δh_I. These additional variables are related to the modifications introduced in the empirical estimation.

There are four alternative reduced forms in Table 8.

30 See footnote 25.

31 Recall that Lucas Type 1 Reduced-Forms are not different from LP Reduced Forms, and in consequence they have already been analyzed.

The reduced form results for ΔP are quite favorable in the L2 model. Signs are generally correct, and in particular the coefficient for ΔM is very close to 1. Levels of significance are satisfactory. R^2 values vary between 0.78 and 0.81 for inflation, with DW statistics between 1.67 and 1.83.

In fact, it seems to be difficult to choose between the reduced form for ΔP in Table 4 – with an R^2 of 0.81 and the presence of a lagged dependent variable – and these three reduced forms in Table 8 – with R^2 between 0.78 and 0.81.

Anyway, all these reduced form favorable results confirm that both the L-P model and the L model perform quite well for the last 30 years as far as Brazil is concerned.

Small Summary

If one were to summarize this estimation, one would make use of the simple model formed by (39) and (40).

Taking the L1 model as an example, we would have, as an *approximation* for the Brazilian case:

$$h^I = 0.55\ \Delta^2 P$$

$$\Delta P = \Delta M - 1.42\ \Delta h$$

In Diagrams 1 and 2, these two relations are presented. It can be seen in those diagrams that they reflect very well the simultaneous movements of ΔP and h.

Let us assume for simplicity here that $\Delta h = \Delta h^I$ or $h = h^I$.

The implicit reduced forms are then:

$$h^I = 0{,}309\ \Delta M + 0{,}439\ h_{t-1} - 0{,}309\ P_{t-1}$$

$$\Delta P = 0{,}561\ \Delta M + 0{,}797\ h_{t-1} + 0{,}439\ \Delta P_{t-1}$$

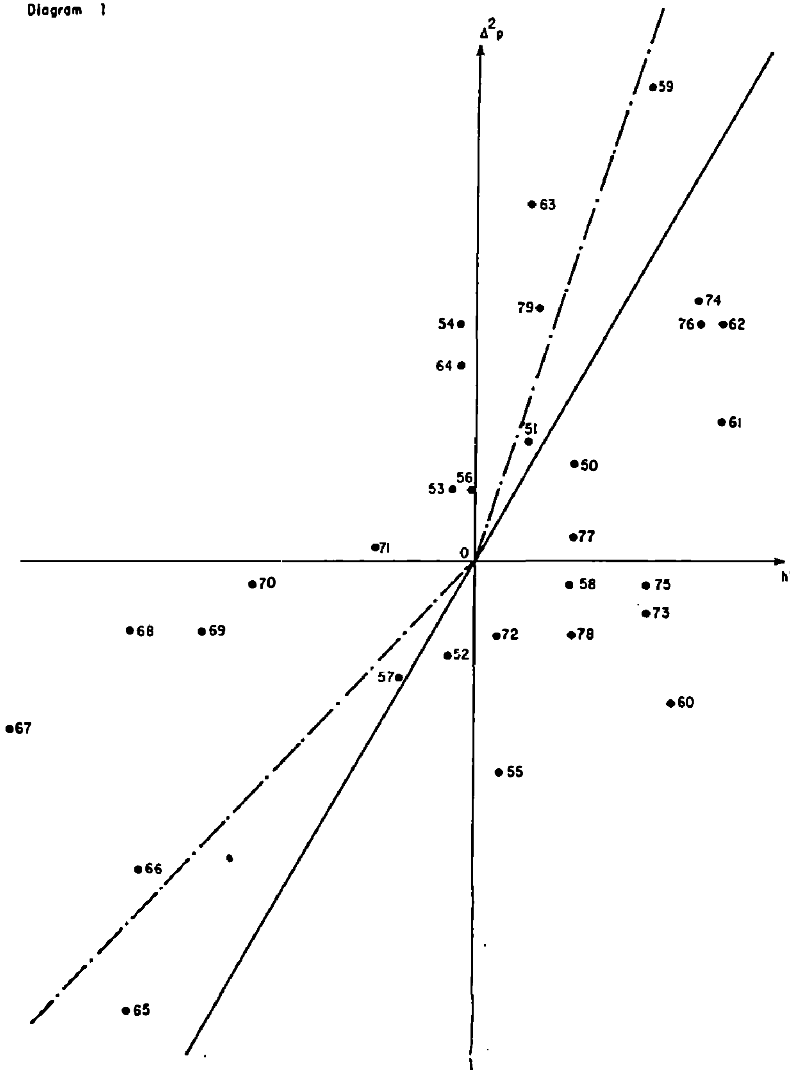
The implicit *final* forms are:

$$h^I = 0{,}878\ h^I_{t-1} - 0{,}439\ h^I_{t-2} + 0{,}309\ \Delta M - 0{,}309\ \Delta M_{t-1}$$

$$\Delta P = 0{,}878\ \Delta P_{t-1} - 0{,}439\ \Delta P_{t-2} + 0{,}561\ \Delta M$$

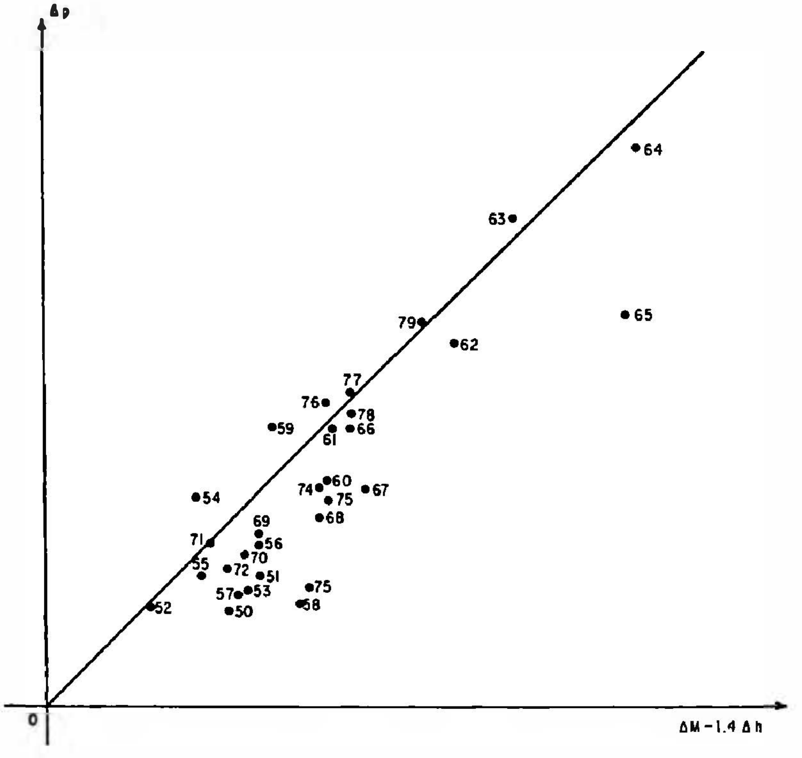
The following table presents an exercise describing the movements of these variables as well as of $\Delta y^I = h^I - h^I_{t-1} + \Delta y^{I*}$

Diagram 1



$\Delta^2 p$
h^1
0
59
63
74
79
54
76
62
64
61
51
50
56
53
77
71
70
58
75
73
68
69
72
78
52
57
60
67
55
66
65

Diagram 2



starting from an "inflationary equilibrium" and assuming a gradualist policy of monetary deceleration in Brazil:

	Money Growth ΔM	Inflation Rate ΔP	Ind. Output (Detrended) h^I	Ind. Growth (Actual) Δy^I	Potential Ind. Growth Δy^{I*}
Year —1	50	50	0	8	8
Year 0	50	50	0	8	8
Year 1	35	42	—5	3	8
Year 2	20	26	—9	4	8
Year 3	10	10	—9	8	8
Year 4	10	3	—4	13	8
Year 5	10	4	1	13	8
Year 6	10	8	2	9	8

This illustration reveals the dillemas and cycles of modern economies, caused by a money growth decceleration policy. We have initially "recession" (a growth recession) and later a boom period (besides cyclical movements in the rate of inflation), that is, all possible combinations of inflation and growth, with oscillatory movements tending to converge to 10, 10, 0, 8 and 8, respectively. Moreover, notice the absence of simple short-run correlations between ΔM and ΔP or ΔM and h in the example, in spite of a monetarist-type aggregate demand formulation. The Phillips Curve movement occurs in years 1 and 2, but is reversed in the following years. In the long run, there is no trade-off between inflation and recession.

In the next section, it will be shown that problems caused by a mix of rational expectations and inflation rigidity might disturb the Phillips Curve relation between h^I and $\Delta^2 P$ in such a way that those dillemas and cycles of modern economies can be even more dramatic and difficult to be cured by economic policy.

5 — A test of the inflation rigidity hypothesis

In the previous section, we have estimated some accelerationist Phillips Curve-type equations under different names. In the LP model, one basic equation was

$$\Delta^2 P = 1.1303 + 0.6330 \ h^I$$

or, including the supply shock variable,

$$\Delta^2 P = 1.0994 + 0.7341 \ h^I - 0.3701 \ h^A$$

In the LI model, we had an inverted version of the LP Phillips Curve, or

$$h^I = -0.5369 + 0.5467 \ \Delta^2 P$$

or, including the supply shock variable,

$$h^I = -0.4336 + 0.4918 \ \Delta^2 P + 0.7435 \ h^A$$

or, including the lagged dependent variable,

$$h^I = -0.1832 + 0.1979 \ \Delta^2 P + 0.7473 \ h^I_{t-1}$$

Moreover, in the L2 model, a policy-oriented Phillips Curve-type equation, derived from rational expectations and aggregate supply theory, was

$$h^I = -0.6059 + 0.3716\ \Delta^2 M + 0.3343\ \Delta^2 M_{t-1}$$

or, with a supply shock,

$$h^I = -0.2700 + 0.4105\ \Delta^2 M + 0.7877\ h^I_{t-1}$$

All these results naturally imply the basic short-run Phillips Curve accelerationist trade-off between inflation acceleration or monetary acceleration and real output, as well the "no trade-off in the long run" implication, because in a steady state the second derivatives of P or M will be zero by definition.

It would be interesting, however, to go deeper into these relations in order to test the hypothesis introduced in Section 3 and related to a mix of rational expectations and inflation rigidity. In particular, the equations derived from Lucas model — both in the L1 and L2 forms — are quite adequate for such a test.

As a matter of fact, we are interested in testing the hypothesis that the real side effects of inflation deceleration or money growth deceleration ($\Delta^2 P < 0$ or $\Delta^2 M < 0$) are much greater than the real effects of accelerating inflation or accelerating money ($\Delta^2 P > 0$ or $\Delta^2 M > 0$). In fact, the latter effects might even be entirely absent due to rational expectations behavior.

With that purpose, we have run a few regressions dissagregating the right-hand side variables $\Delta^2 P$ or $\Delta^2 M$ between positive and negative values. that is, $\Delta^2 P^+$ or $\Delta^2 P^-$ and $\Delta^2 M^+$ or $\Delta^2 M^-$. The method was OLS and the period was the same of the previous regressions: 1950/79 (annual data). The results are presented in Table 10. For example, $\Delta^2 P^+$ is equal to $\Delta^2 P$ when we have a positive value and zero otherwise; $\Delta^2 M^-$ is equal to $\Delta^2 M$ when we have a negative value and zero otherwise; and so on.

The results tend to confirm our hypothesis. One finds no significant connection between accelerating inflation or monetary acceleration and the real variable expressed by h^I or industrial output detrended. Rational expectations seem to be acting on the positive axis of the Phillips Curve leading to an almost total absence of trade-offs between inflation and real output, even in the short-run.

On the other hand, when we move to the negative axis, the situation changes dramatically: the trade-off is found, with very

Table 10

Inflation Rigidity and Rational Expectations
Dependent Variable: h^I

Explanatory Variable	X.1	X.2	X.3	X.4	X.5
Constant	2.1586	1.1043	2.9014	0.8104	0.2055
	(2.09)	(0.59)	(1.87)	(0.42)	(0.08)
Δ^2P+	−0.1337	0.3228	0.0254		
	(−0.92)	(1.32)	(0.12)		
Δ^2P-	0.5977	0.8386	1.0644		
	(3.64)	(2.84)	(4.42)		
h^I_{t-1}	0.7773				
	(8.03)				
h^A			0.9504		
			(4.10)		
Δ^2M+				0.1795	0.3807
				(0.51)	(0.99)
Δ^2M-				0.5508	0.3405
				(2.25)	(1.39)
$\Delta^2M^+_{t-1}$					0.0947
					(0.26)
$\Delta^2M^-_{t-1}$					0.4391
					(2.00)
h^A_{t-1}				0.8856	
				(2.82)	
Statistics					
R^2	0.8206	0.3754	0.6204	0.3492	0.2888
DW	1.6767	0.8531	1.4851	0.7016	0.6162
SER	3.4939	6.3984	5.0828	6.6555	7.0954

significant coefficients. It seems that downward inflation rigidity contributes to maintain a short-run trade-off between decelerating inflation or decreasing money growth and real output.

These comments are valid for the 5 cases presented in Table 10. Graphically, one could imagine a kinked Phillips Curve (look at Diagram 1 for an example), which would be almost vertical for positive values of Δ^2P or Δ^2M, but would have a very significant slope (positive in our case [32] for negative values of Δ^2P or Δ^2M. It is true that the regressions with Δ^2P^- tend to be more significant

[32] Notice that $h^I \equiv y^I - y^{I*} \equiv \Delta y^I - \Delta y^{I*} + h^I_{t-1}$. Evidently, h^I corresponds to an inverted indicator of unemployment. If the graph we are describing had $-h^I$ in one of the axis instead of h^I, then the lower part of the Phillips Curve would have a negative slope — just like in the inflation-unemployment graph.

than the ones with Δ^2M^- and Δ^2M_{t-1}, but the point is that in both cases the differences between acceleration movements and deceleration movements are very clear.[33]

In fact, and this is interesting to emphasize, the 0.5-0.6 slopes of the conventional Phillips Curves for Δ^2P are "divided" between a non-significant coefficient for Δ^2P^+ and a 0.8-1.1 coefficient for Δ^2P^-. In words, a 5% inflation deceleration provokes a 5% reduction of real excess demand, that is, a 5% deceleration of the rate of industrial growth in the short-run. But a 5% inflation acceleration has practically no real effects. As far as money growth variables are concerned, the results have lower significance than the ones for inflation, but the trade-off coefficients for negative values Δ^2M^- and Δ^2M_{t-1} are situated between 0.55 and 0.80: a 5% deceleration in money growth can provoke up to a 4% real deceleration in the economy in the short-run.

Therefore, rational expectations and downward inflation rigidity seem to lead us to a very complicated real world as far as short-run policy is concerned. There are no real gains from monetary and inflation accelerations, but there are real losses — although temporary — from monetary and inflation deceleration.

The relevance of our findings ought to be emphasized. An antiinflationary policy — say a 5% to 15% contraction of money growth — would have very severe real effects, by reducing industrial output (detrended) and its rate of growth (see the numerical example of the Section 4 — final subsection). On the other hand, assymetrically, an acceleration of the expansion of the money supply would just lead to more inflation.

These results have a direct implication on the effects of stop and go economic policies. Due to the assymetry of effects, a policy of equal accelerations and decelerations in the money supply would not be neutral with respect to inflation and unemployment. For example, starting from a 20% money growth, a variable policy with 30%, 20%, 30% and 20% in the following years would lead in the short-run to higher inflation and higher unemployment than before. Higher inflation will be caused not only by the new average (25%) but also by the assymetric effects of accelerations and decelerations, while higher unemployment would be caused by the *negative* movements in the third and fifth year, since the *positive* movements would have no real effects.

33 Clearly, this result could be regarded simply as an indication of non-linearity in the relationship between Δ^2P and h^t. But it must be emphasized that this non-linearity is not of the same type as the one often obtained for short-run Phillips relations, since it continues to be valid in the long run.

Obviously, two implications of the kinked Phillips Curve are: a) if one starts from a low and satisfactory level of inflation, the best policy is one of a constant rate of growth in the money supply; b) if one starts from a high and unfavorable level of inflation, then the Government should attempt to undertake a consistently gradual program of money growth deceleration — without stop and go movements — in order to bring inflation down. But the cost of stopping inflation will continue to occur, that is, the unfavorable side effect in terms of less employment and less output. On the other hand, if there is no consistent program, stop and go policies — combining expansions and contractions — will leave us with both inflation and unemployment, because of the assymetries implicit in this mix of rational expectations and inflation rigidity behavior.

6 — Summary and conclusions

This paper has discussed inflation-real output macromodels and presented estimates of such models for Brazil. Moreover, problems related to rational expectation formulations and inflationary rigidity have been taken into consideration.

In Section 2, there was a formal presentation of the LaidlerParkin model and the Lucas model, leading to formulations that permit a simultaneous determination of variables such as ΔP, $\Delta^2 P$, h, Δh, Δy, that is, inflation and its acceleration, industrial output gap and its rate of change, real growth, etc. Besides the simultaneity aspect, there was some emphasis on dynamic properties such as long and short-run multipliers as well as cyclical movements of the variables, led by instability in economic policy.

Rational expectations ideas were rapidly reviewed in Section 3, going from the micromodels of Muth to the macromodels of Lucas and others such as Barro, Sargent, and McCallum. The policy-inefectiveness implication of rational expectations was mentioned, as well as the main criticisms to this approach. In particular, we emphasized the possible existence of a certain degree of rigidity in inflation rates, specially in the downward direction.

Estimates of the models for Brazil are in Section 4. In general, the equations of the Laidler-Parkin model and the Lucas model perform very well, with correct and significant signs, as well as reasonable statistics (R^2, DW), given the nature of certain dependent variables (rates of change and changes in the rate of change). A simplified summary of the results is presented, with some diagrams and a simulation exercise.

Finally, in Section 5, we consider the possibility of a kinked-type accelerationist Phillips Curve in Brazil where inflation acceleration would have much less (positive) real effects than inflation deceleration in terms of (negative) real effects. The simple tests suggest that such a possibility cannot be rejected. Obviously, this result has important implications for economic policy to the extent that a mix of rational expectations behavior and inflation rigidity can make the inflation-recession dillema even more serious.

The main conclusions of this paper are as follows. There is no doubt that there exists in the short-run a dillema between recession or real excess supply and the acceleration of inflation. Moreover, it is clear that money has a relevant effect on both variables in the short-run. However, there is a serious probability that this dillema is much more meaningful in policy contractions than in policy expansions, since in the latter case real effects of policy could be minor. All these results imply that one should avoid stop and go policies and at the same time should aim at a low and constant rate of monetary growth, in order to halt inflation. It seems that models such as the ones presented in this paper are valid both for countries such as the United States or Brazil — the macroeconomic mechanisms that cause the trade-off are present in both countries. Further research would be necessary to take into account the openess of modern economies in these models, [34] but preliminary exercises suggest no significant differences from the present results.

Clearly, simple macromodels such as the ones presented in this paper should not be used strictly for projection exercises. Of course, the very simplicity of the models indicates that its usefulness is much more related to its contribution to a general understanding of the inflation-recession dillema than to its capacity to make precise forecasts about inflation and real growth in the following year. Therefore, the general interest of the type of analysis undertaken here should not be closely linked to the behavior of inflation and growth in Brazil in 1979 or 1980, as compared to the "predictions" of our theory. But it seems to be correct to say that the results obtained in this paper provide some hints as to the main reasons for the substantial acceleration of inflation in Brazil since 1976. Besides some obvious supply shocks, this acceleration is certainly associated to excess demand and higher

[34] In particular, money supply could be made endogenous and import and export prices would have to be considered in the price equation, as well as the exchange rate.

rates of monetary expansion since 1973. The 1979/80 inflationary explosion represented the *climax* of these supply shocks and aggregate demand pressures.

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Equations for import demand in Brazil: 1960/80 *

Maria de Fátima S. Pombal Dib **

1 — Introduction

The purpose of this article is to obtain estimates for import demand in Brazil in the period from 1960 to 1980.

In view of the present-day aim to obtain a considerable balance of trade surplus and the vital need for measures designed to restore control of the balance of payments, import control policies are a part of any strategy for solving the problem of Brazil's foreign accounts. It is, therefore, desirable to improve our knowledge of the behaviour of import demand.

Given the great variation of the forms of control that have been attempted in recent Brazilian history, [1] it is difficult to synthesize the policy measures into a small number of variables open to statistical observation, and which can be the subject of an econometric treatment. Despite these difficulties, the attempt to give an econometric treatment to import demand is justified by the need for the availability of quantitative assessments concerning the impact of changes, for example, in the exchange policy, in the level of economic activity, in tariff policy, or in variables

Editor's note: Translation not revised by the author.

* This article presents some of the results of the master's dissertation entitled "Brazilian Imports: Control Policies and Demand Determinants", presented to the Economics Department of the Pontifícia Universidade Católica do Rio de Janeiro.

The author thanks Dionísio Dias Carneiro for this supervision while accepting total responsibility for any errors that the thesis may still contain. Another version of this article was published in *Revista Brasileira de Economia*, vol. 35, n.º 4, October/December 1981.

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1 See, for example, Dib (1983, Chapter II).

outside internal control, such as the international prices of Brazilian imports.

With a view to obtain accurate estimates for the price and income elasticities of import demand, this article is structured in the following way: the first section gives a brief review of the literature dealing with the import demand models that are used; the second section analyses the data utilized, and the third section presents the results obtained in the econometric estimates for Brazil between 1960 and 1980, and compares our results with those obtained by Lemgruber (1976), Weisskoff (1979) and Khan (1974); and finally, the last section contains various comments on the need to make import control compatible with the definition of viable growth strategies for the next few years. The results that are presented confirm and extend those obtained by the author in a preliminary study. [2]

2 — Import demand models

The methodology used in the research sought to incorporate most of the information available in the literature on the subject. The theory has identified the relative income and price variables as the determinants of import demand. However, as will be seen, certain additional factors can also be related, such as trend variables and changes in economic policy management.

2.1 — The simplified model

According to Lemgruber (1976), Pastore, Barros e Kadota (1976) and Thirlwall (1977), the specification of the aggregate import demand equation relates the quantity of imports demanded to domestic economic activity, measured by the real income variable, and to the price of the imported product in relation to that of the domestic substitute. [3] The estimated equation for the annual data has the following form:

$$\log M^d = \alpha_0 + \alpha_1 \log Y + \alpha_2 \log \frac{r_m (1 + t) \lambda}{p} + \mu_t \qquad (1)$$

where:

M^d = quantity of import demanded

[2] Dib (1981).

[3] It is assumed that the demand function is zero degree homogeneous in the prices and income.

Y = real income

P_m = price of imports in foreign currency

P = domestic price index

λ = exchange rate (Cr$/US$, annual average)

t = tariff protection rate

and

μ_t = is the independently and normally distributed sample error.

Working on the hypothesis that Brazil is a "small" country in relation to the supplier markets, we will assume that the supply of imports is infinitely elastic at the price of imports in foreign currency, and we will also suppose that the domestic prices and real domestic income are exogenous,[4] the equation (1) being estimated by the ordinary least squares method with annual data. Besides this, it is initially accepted that the adjustment lag is not for more than a year, that is to say that the importers are aleays over their demand function ($M^d = M$).

As the equation is specified in logarithmic terms, α_1 and α_2 are respectively, the price and income elasticities of the imports. The theory predicts that the α_2 sign is negative, while the income elasticity sign, α_1 is positive.

Similarly to Khan (1974), Lemgruber (1976) and Weisskoff (1979), we started from the simplest formulation of the aggregate import demand and, after this, we relaxed certain hypotheses so as to qualify the initial model. In this way we will add to the initial hypotheses, the possibility of adjustment lag, of dummy variable influences (as a result of atypical movements that may have characterized any specifi year, and we will introduce trend variables designed to separate the cyclic and secular effects of economic activity.

2.2 — The adjustment lag model

In a study designed to estimate import functions for fifteen underdeveloped countries, including Brazil, Khan (1974) attempted

[4] This is a simplifying hypothesis. Strictly speaking, given the dependence of investment and the level of activity in general on non-substitutable imports, imports (or part of them) should be a part, in a more general model, of the income determinants.

to incorporate the role of quantitative restrictions on trade. This is necessary to the extent that the quantitative controls have appeared as a characteristic common to underdeveloped countries. Although the exact role of these quantitative restrictions cannot be quantified, Khan shows that, by assuming certain hypotheses, it is possible to obtain approximations and tests that can measure their importance.

In the context of the underdeveloped countries, then, an additional source of bias in the estimation of the equation (1) is the omission or lack of attention given to the quantitative restrictions in the flow of imports, leading to inconsistend and tendentious estimates.

In this way, in order to avoid these sources of bias, Khan (1974) adds to the initial hypotheses the possibility of adjustment lags, specifying a disequilibrium formulation, in which additions are related to the difference between the demanded imports in the period t and the effective imports in the period $(t - 1)$.

$$\Delta \log M_t = \gamma (\log M_t^d - \log M_{t-1}) \qquad (2)$$

where,

$$\Delta \log M^t = \log M_t - \log M_{t-1}$$

and γ is the coefficient of adjustment, $0 \leqq \gamma \geqq 1$.

The equation (1), presented in this way, is a relation of equilibrium implying instantaneous adjustment, on the part of the importers, to the changes in the relative prices of the imports and in the real income. Given that there is an adjustment cost of the effective imports at their desired level, and given that some imports are undertaken by way of contracts with a duration of more than one year, being unable to respond immediately to changes in demand, it is easy to notice that there exists the possibility of an adjustment lag.

This process of partial adjustment introduces a structure of distributed adjustment lags (i. e., where the prices decline geometrically) in the determination of imports.

In this way, substituting the equation (1) into (2) and resolving to M_t, we will have:

$$\log M_t = y\,\alpha_0 + y\,\alpha_1 \log Y + y\,\alpha_2 \log \frac{P_m (1 + t)\,\lambda}{P} +$$

$$+ (1 - y) \log M_{t-1} + y\,\mu_t \qquad (3)$$

where $y\alpha_1$ and $y\alpha_2$ are, respectively, the short-term price and income elasticities.

2.3 — Cyclical and secular effects on import demand

The unsophisticated specification of the aggregate import demand equation discussed previously has been criticized as it does not allow for a distinction between cyclical factors and secular factors.

The effects of the cyclical factors may be substantially different from the effects arising from secular actors. The introduction of the current real income as an explanatory variable may include only the cyclical influences on imports. Thus, to ignore the role of the secular factors may lead to a tendentious estimate of the income elasticity, especially concerning the long-term uses of the estimated equations.

Previous attempts to formulate an import demand equation in which it would be possible to isolate the two effects, were made by Branson (1968) for the United States, and by Marston (1971) for Great Britain. The method by which Khan and Ross (1975) took these cyclical and tendency factors into account will be explained below.

Specifying, firstly, the import demand equation similar to equation (1), but incorporating the quantity of imports and the income in terms of tendency, we have:

$$\log M^{*d} = \alpha_0 + \alpha_1 \log Y^{*} + \alpha_2 \log \frac{P_m (1 + t) \lambda}{P} + V_t \quad (4)$$

where the asterisk indicates the value of the tendency, that is to say, the potential value of the variable and V_t is the sample error.

Additionally, a new equation relates the effective import demand deviation in relation to its potential level to the deviation in effective real income in relation to the potential income. That is to say:

$$\log M^{d} - \log M^{*d} = \Phi (\log Y - \log Y^{*}) + W_t \quad (5)$$

Substituting (4) into (5) and resolving to M^d we obtain the final import demand equation:

$$\log M^{d} = \alpha_0 + \alpha_2 \log \frac{P_m (1 + t) \lambda}{P} + \Phi \log Y +$$

$$+ (\alpha_1 - \Phi) \log Y^{*} + \mu_t \quad (6)$$

where $\mu_t = (V_t + W_t)$

The coefficient sign of the potential income, $(\alpha_1 - \Phi)$ will therefore depend on the relative magnitude of the two parameters and may even be negative. [5]

Many studies have not included the potential income as an explanatory variable in the import demand equation and this explains why nearly all of them fail to find evidence of negative elasticity, given that their results only concerned the estimate of the parameter Φ.

While the bulk of econometric work in this fields concluded that both the level of economic activity and the trend variable are important, there is disagreement on the importance of cyclical variables. Marston (1971), Gregory (1979) and Hughes and Thirlwall (1977) found significant positive results for the cyclical variables, while Barker (1970 and 1977) concluded that cyclical influences, such as demand pressures, the utilization of capacity or stocks, are insignificant, unstable and generally present a sign contrary to that expected.

Effects of variations at the levels of economic activity, tendency variables and cyclical variables (a measure of the level of utilization of capacity) should then be included in the import demand functions. Although the need to consider these explanatory variables is recognized, there is a problem of identifying these variables, given that any two out of three will imply the third. Therefore, the introduction of the three variables make the model overdetermined and the decision between them cannot be solved empirically.

Barker (1979) observes that the parameters relative, respectively, to the income, cyclical and tendency variables cannot be distinguished independently. With the aim of illustrating this problem of interpretation of the parameters, the author estimated a series of import demand functions for the United Kingdom, at an aggregate and disaggregate level, removing, individually, the variables of economic activity, the cyclical variables and the tendency variables. On the whole, the estimates conform to the results that could be expected a priori.

When the trend variable is omitted, the income elasticity increases and the elasticity of cyclical demand is reduced. Similarly,

[5] An alternative way to present the equation (6) would be:

$$\log M^d = \alpha_0 \alpha_1 \log \frac{P_m (1 + t) \lambda}{P} + \Phi (\log Y - \log Y^*) + \alpha_1 \log Y^*$$

This is the formulation used by Artus (1973) and Goldstein, Khan and Officer (1979) although it is obvious that the results of one can be derived from the other.

the omission of the cyclical demand variable would result in an increase of the income elasticity and a reduction of the tendency elasticity. Finally, when the income variable is omitted, both the cyclical demand elasticity and the tendency elasticity grow. The only exception to the algebraic model found by Barker (1979) is an increase in the tendency variable when the cyclical variables is withdraw, which is perhaps associated to a combination of negative cyclical demand elasticities with goods for the demand is declining through time.

In view of these empirical results, we can conclude that the omission or restriction of certain variables in estimating the import demand functions may lead to badly biased results.

According to Barker's models (1979), the distinction between the effects of variations in the levels of economic activity, tendency variables and cyclical variables (a measure of the level of utilization of capacity) can be explained by means of an algebraic derivation based on the simplest formulation of the import demand equation. [6]

The first specification incorporates into the simplest form the tendential variable T:

$$\log M^d = \alpha_0 + \alpha_1 \log Y + \overset{\prime}{1}\, T + \\ + \alpha_2 \log \frac{P_m (1 + t)\lambda}{P} + \mu_t \tag{7}$$

In the second, instead of the real income a capacity variable is introduced, measured by the ration between real product and potential product (Y/Y^*), and the tendency variable (T):

$$\log M^d = \alpha_0 + \alpha_1 \log \frac{Y}{Y^*} + \alpha \overset{\prime}{1}\, T + \\ + \alpha_2 \log \frac{P_m (1 + t)\lambda}{P} + \mu_t \tag{8}$$

Whilst in the third, an attempt is made to include the influence of both effects by utilizing the real income (Y) and the level of utilization of capacity (Y/Y^*).

$$\log M^d = \alpha_0 + \alpha_1 \log Y + \alpha \overset{\prime}{1} \log \frac{Y}{Y^*} + \\ + \alpha_2 \log \frac{P_m (1 + t)\lambda}{P} + \mu_t \tag{9}$$

[6] See, in this respect, Barker (1979, pp. 63-4).

Finally, the last specification incorporates the potential income (Y) and the capacity variable (Y/Y^*):

$$\log M^d = \alpha_0 + \alpha_1 \log (Y/Y^*) + \alpha' 1 \log Y^* +$$
$$+ \alpha_2 \log \frac{P_m (1 + t) \lambda}{P} + \mu_t$$

It should be stressed that none of the specifications (10 or 6a) mentioned present any problems related to the identification of variables as suggested by Barker (1979), because none of them is of the form:

$$M^d = \alpha_0 + \alpha_1 Y + \alpha_2 c + \alpha_3 t$$

where:

$$c = \frac{Y}{Y^*} \; e \; Y_t^* = Y_0 (1 + \bar{r})^t$$

M^d = quantum total of imports;
Y = real product;
c = utilization of capacity variable;
t = tendency variable;
Y^* = potential product;
Y_0 = real product in base year (period 0);
Y_t^* = potential product in period t; and
$\bar{r}$ = average growth rate of the aggregate potential product.

Thus, we consider as alternative specifications Khan's model of partial adjustment (1974); and the Khan and Ross model (1975) and Barker's model (1979) in which, besides the real income variable, other variables are also included so as to attempt to separate the cyclical effects from the tendential effects.

Given that in estimating the import demand equations there is a particular interest in obtaining price and income elasticities, the logarithmic specification in the most convenient, as these elasticities can be obtained directly from the estimated equations.

So, as the equations estimated here are specified in logarithmic terms, α_i, $i = 0, \ldots 2$, they are, respectively, the elasticities that are sought.

It should be expected, a priori, that the signs of the income elasticities will be positive and that the price elasticity sign will be negative.

While there appears to be no doubt as to the price and income elasticity signs, the interpretation of the "elasticities" relative to the level of the utilization of capacity, of potential income and of the tendential variable is less obvious. [7]

The inclusion of these variables is an attempt to include costs, other than those directly measured by the prices, related to the purchase of particular goods. The import-domestic production balance depends on the cost and price of obtaining these goods, especially the cost of having to queue up when there is excessive demand not absorbed by price increases or, alternatively, if the goods are produced to order. Normally an increase in the utilization of internal installed productive capacity corresponds to an increase in the waiting time for domestic consumers, who will try to avoid this cost increase by seeking external suppliers. [8] In addition, there are technological factors which are not being considered, at least in their short-term effects, by the activity variable. The exhaustion of installed productive capacity, in periods of intense growth in industrial production, induces a rise in the level of investment and, consequently, in the gross fixed capital formation. If the economy in question does not yet possess a sufficiently dynamic production goods sector compatible with the size of its productive base, the effect of an increase in the demand for these goods, especially capital goods, tend to overflow abroad through an increase in the imports of such products. [9] The expected sign of the elasticity relative to the utilization of capacity is, therefore, positive.

3 — The data

Given the peculiarities associated with the laws that control the importation of wheat, the leading Brazilian import, and of oil, the main intermediate product, we decided to exclude them form the econometric analysis of the determinants of the quantum imports.

Bearing in ming the great variety of economic policy measures designed to solve the balance of trade deficits in the post-war

7 See the development of the Khan and Ross model (1975) in previous paragraphs.

8 See Abreu and Horta (1982).

9 The years 1972 and 1973 are good examples of such an argument.

period, particularly as regards exchange policy, the period 1960/80 was chosen because it presented greater stability in terms of foreign exchange regulations, with the exchange rate reflecting more adequately of the scarcity of foreign exchange.

We sought to estimate an *ad valorem* tariff index that would reflect in the most suitable way possible the effective cost of imports. Thus, the tariff on imports was calculated by way of the taxes effectively received during the period, obtaining the effective aliquot by the quotient between the import tax received and the FOB value of the imports, not including wheat and oil.[10] It should be observed that the series of tariffs becomes somewhat difficult to construct given the diversity of policies introduced in the post-war period (1947/80). We believe, however, that this series can be used as a good indication of variations in the nominal protection granted by the country to productive activities.

The indices of price and import quantum (oil and wheat excepted) P_m and M^d, were constructed on the basis of information provided by the Getulio Vargas Foundation National Accounts Center.[11]

The domestic price index (P) was the wholesale Price Index, internal availability, obtained in *Conjuntura Econômica*. The object of this choice was to obtain an index that reflected the behaviour of the prices of marketable goods and which, therefore, was the most appropriate as a deflator of the cost of imports.

The other data were the real product index for the real income variable, and the price of the dollar in cruzeiros (annual average) for the exchange rate variable (λ) published monthly by *Conjuntura Econômica*.

The measure of the level of the utilization of capacity corresponds to the ratio between the Real Product and the Potential Product (Y/Y^*). The data used for the potential product were those estimated by Resende and Lopes,[12] the implicit growth rate of which was 7.21% a year.

Finally, we included the dummy variable in 1974 ($D = 1$ for 1974, $D = 0$ for the other years) to reflect the speculative demand

10 That is to say: $\dfrac{RT}{(M - MP - MT)^{\lambda}}$

where: RT = total import tax revenue in cruzeiros
M = total value of imports (FOB)
MP = value of petroleum imports (FOB)
MT = value of wheat imports (FOB)
λ = average exchange rate Cr$/US$

11 Not published by *Conjuntura Econômica*.

provoked by generalized expectations of scarcity and international inflationary acceleration, which are not reflected in the conventional relative price and income variables.

4 — The empirical results

The specifications presented were estimated in two versions as far as the relative price variable was concerned. In the first case, we removed from the price variable the import tariff index while, in the second case, the effective real cost of imports incorporates the international price of imports (P_m), the real exchange rate (P) and the tariff index $(1 + t)$. This procedure was designed to test the importance of the effects provoked by the imposing of tariffs on exports.

The estimated regressions relative to the two cases are reproduced, respectively, in Tables 1 and 2 below.

The introduction of the tariff index in the relative price variable allows a slight improvement in the statistical quality of the test and a small increase in the value of price elasticities in all the estimated specifications. Although this result shows the sensitivity of the import quantum demand in relation to the variations of the tariff aliquots, tending to suggest that it is reasonable to include this variable in the exercise, the small level of difference between the two cases is due to the small variation shown by the average tariff calculated for the period.

Generaly speaking, all the results obtained are very satisfactory from the econometric standpoint; all the coefficients show expected signs and reasonable accuracy in the estimates.

The results of the tables seem to confirm the evidence of a high income and price elasticity of import demand in Brazil, even if just the simplest specification is considered (equation (1)).

The inclusion of other variables as a measure of the level of economic activity, so as to differentiate the cyclical and tendencial effects in import demand, however, implies obtaining extremely accurate coefficient estimates and values for price and income elasticities more in accordance with what would be expected a priori.

When we disaggregate the activity variable into real income (Y) and potential income (Y^*), the real income elasticity rises considerably, while the potential income elasticity appears negative

12 See Resende and Lopes (1981).

Table 1

Equations of Total Import Demand – Brazil (Case I) (Excluding Petroleum and Wheat)

Variables Specifications	(c)	(Y)	(Y^*)	($\frac{Y}{Y^*}$)	(T)	(Pr)	(M_{t-1})	DUM–74	R^2	F	DW	RHO	Standard Error of Estimate	Period
1. Equation (1)	1,02382 (0,570302)	1,50895 (8,11513)	—	—	—	—0,750944 (—2,13151)	—	—	0,9689	264,894	1,7142	0,67716	0,127498	1960/80
2. Equation (3)	3,88530 (2,72239)	0,588677 (2,43988)	—	—	—	—0,993936 (—3,45908)	0,580542 (3,57638)	—	0,9729	179,395	—	—	0,123873	1961/80
3. Equation (6)	5,22886 (3,04040)	3,29327 (5,96897)	—2,23223 (—3,50002)	—	—	—1,12833 (—3,74899)	—	—	0,9760	217,068	1,8032	0,33696	0,115421	1960/80
4. Equation (7)	—3,74929 (—1,72407)	3,29111 (—3,74484)	—	—	—0,154750 (—3,49102)	—1,12850 (—3,74484)	—	—	0,9760	216,629	1,9050	0,33725	0,115535	1960/80
5. Equation (8)	9,49711 (6,48739)	—	—	3,29231 (5,96775)	0,0736592 (8,46144)	—1,12764 (—3,74783)	—	—	0,9760	217,275	1,8025	0,33727	0,115367	1960/80
6. Equation (9)	5,22888 (3,04042)	1,06104 (8,45931)	—	2,23224 (3,50005)	—	—1,12833 (—3,74901)	—	—	0,9760	217,069	1,8032	0,33696	0,115421	1960/80
7. Equation (10)	5,22893 (3,04045)	—	1,06104 (8,45938)	3,29329 (5,96907)	—	—1,12834 (—3,74904)	—	—	0,9760	217,069	1,8033	0,33695	0,115421	1960/80
8. Equation (1*)	2,13412 (1,46782)	1,48877 (10,2784)	—	—	—	—0,969959 (—3,38905)	—	0,288036 (3,38905)	0,9817	285,696	1,8385	0,67694	0,100898	1960/80
9. Equation (9*)	5,40551 (3,98200)	1,07411 (10,8063)	—	1,96328 (3,84375)	—	—1,18851 (—4,99188)	—	0,293798 (3,29150)	0,9861	265,586	2,1366	0,34274	0,098315	1960/80
10. Equation (9**)	5,62905 (3,28754)	0,797011 (2,99751)	—	1,48041 (1,59772)	—	—1,22782 (—4,01600)	0,274927 (1,10314)	—	0,9771	149,047	—	—	0,118040	1961/80
11. Equation (6**)	5,62902 (3,28755)	2,26740 (2,09194)	—1,48039 (—1,59771)	—	—	—1,22781 (—4,01602)	0,274932 (1,10317)	—	0,9771	149,048	—	—	0,118040	1961/80

OBS.: The Durbin statistic for the coefficient of import lag (M_{t-1}) in equation (3) is —0.3957 and is not defined for equations 9** and 6**

c = Constant

Y = Real Product

Y* = Potential Product

(Y/Y*) = Utilization of capacity variable (measured by product gap)

T = Tendency variable

Pr = Relative prices (measured by $\frac{Pm \cdot \lambda}{P}$)

M_{t-1} = Import lag (t—1)

DUM—74 = Dummy variable 1974

* = Includes the dummy variable (1974) in the original specif.

** = Includes the dependent lag variable in the original specification

The figures in brackets refer to the statistic *t*.

Table 2

Equations of Total Import Demand — Brazil (Case II) (Excluding Petroleum and Wheat)

Variables Specifications	(c)	(Y)	(Y*)	$\left(\frac{Y}{Y^*}\right)$	(T)	(Pr)	(M_{t-1})	DUM—74	R^2	F	DW	RHO	Standard Error of Estimate	Period
1. Equation (1)	1,65544 (0,922075)	1,48537 (8,59149)	—	—	—	—0,859235 (—2,46518)	—	—	0,9709	283,438	1,8074	0,6692	0,123382	1960/80
2. Equation (3)	4,31924 (3,39228)	0,538845 (2,53807)	—	—	—	—1,05358 (—4,17940)	0,599360 (4,17367)	—	0,9755	198,938	—	—	0,117883	1961/80
3. Equation (6)	5,68561 (3,54557)	3,25277 (6,48828)	—2,21122 (—3,81254)	—	—	—1,20516 (—4,30444)	—	—	0,9709	243,435	1,9501	0,31308	0,109133	1960/80
4. Equation (7)	—3,20843 (—1,57587)	3,25106 (6,47591)	—	—	—0,153328 (—3,80377)	—1,20559 (—4,30047)	—	—	0,9785	242,931	1,9515	0,31331	0,109244	1960/80
5. Equation (8)	9,87583 (7,25029)	—	—	3,25205 (6,40808)	0,0723022 (9,05004)	—1,20454 (—4,30358)	—	—	0,9786	243,666	1,9197	0,31330	0,109082	1960/80
6. Equation (9)	5,68559 (3,54557)	1,04155 (8,04689)	—	2,21121 (3,81254)	—	—1,20515 (—4,30443)	—	—	0,9786	243,436	1,9501	0,31308	0,109133	1960/80
7. Equation (10)	5,68560 (3,54556)	—	1,04155 (8,04697)	3,25276 (6,48826)	—	—1,20515 (—4,30441)	—	—	0,9786	243,433	1,9502	0,31308	0,109133	1960/80
8. Equation (1*)	2,45807 (1,69082)	1,47171 (10,4190)	—	—	—	—1,02020 (—3,62494)	—	0,274062 (3,31111)	0,9827	303,409	1,8800	0,67875	0,0979614	1960/80
9. Equation (9*)	5,49718 (4,20469)	1,06934 (10,9972)	—	1,88861 (3,83209)	—	—1,20363 (—5,27742)	—	0,270644 (3,18458)	0,9872	289,046	2,0015	0,35332	0,0871166	1960/80
10. Equation (9**)	5,81348 (3,71749)	0,78820 (3,20306)	—	1,12399 (1,66697)	—	—1,25837 (—4,52977)	0,264866 (1,12652)	—	0,9795	167,405	—	—	0,111521	1961/80
11. Equation (6**)	5,81349 (3,71749)	2,21218 (2,19837)	—1,42398 (1,66696)	—	—	—1,25838 (—4,52975)	0,264869 (1,12654)	—	0,9795	167,404	—	—	0,111521	1961/80

OBS.: The Durbin statistic for the coefficient of import lag (M_{t-1}) in the equation (3) is —0,4773 and is not defined for equations 9** and 6**

c = Constant
Y = Real Product
Y* = Potential Product
(Y/Y*) = Capacity variable (measured by product gap)
T = Tendency variable
Pr = Relative prices $\left(\text{measured by } \frac{Pm\ (1 + t)\lambda}{P}\right)$

M_{t-1} = Import Lag
DUM—74 = Dummy variable 1974
* = includes dummy variable (1974) in the original specification
** = includes the dependent lag variable (M_{t-1}) in the original especifications
The figures in brackets refer to the statistic t.

which means say that import quantum demanded varies inversely with the trend incometrend (see equation (6)).

In the same way, the negative coefficient of the tendency variable (*T*) can be interpreted as reflecting, according to Weiss-koff, a growing decline in Brazilian dependence on aggregate imports during the period in question (equation (7)).

It is worth mentioning, however, that when the utilization of productive capacity variable is introduced as a measure of cyclical influence in import demand, the trend elasticity becomes positive (equation (8)). Bearing in ming these contracdictory results in relation to the trend variable, its interpretation starts to demand a little more attention and will be the subject of more detailed qualification in the following paragraphs.

The high elasticity in relation to the utilization of capacity variable, suggests the importance of extra-price factors in the competition between imported goods and goods produced internally.

The reduction in estimated income elasticity can be seen as a result of the separation of cyclical effects from secular effects. On the other hand, the increase in the estimated value of price elasticity in relation to the simpler specification — which occurs in all the other cases — reflects the separation of price and extra-price effects of a cyclical nature relevant to the competition between domestic production and importation.

As one approaches the full utilization of productive capacity, the surplus demand for domestic production tends to switch abroad in the form of a rise in import demand. Besides thise, the greater rigidity that characterizes the present list of Brazilian imports, above all in relation to intermediate goods imports (particularly oil) and capital goods (particularly those made to order) — which represent almost the total volume of imports [13] — reflects the greater influence of extra price effects *vis-à-vis* price effects. In this way, changes of the imported quantum seem to respond more to oscillations in the level of economic activity than to modifications in the relative price structure.

One can conclude that the introduction of the potential income variables, and the tendency and utilization of capacity variable, allows us, in the light of the Khan and Ross models (1975) and the Baker model (1979), to distinguish between cyclical and secular factors relative to the movement in the

[13] For an analysis of the evolution of post-war imports, see, for example, Dib (1983, Chapter III, Section 3.2).

variables for measuring economic activity and between price and extra-price effects in import demand.[14]

The results obtained with the inclusion of a dummy variable for the year 1974, in an attempt to explain the atypical accumulation of speculative stocks, confirms the peculiar character of that year, though not altering in any substantial way the estimated elasticity values (equations (1*) and (9*)). The reduction of income elasticity in those cases can be explained by the greater influence of extra-price and income effects, which characterize atypical years in terms of import performance, as well as the fear of generalized international inflation and a new period of drastic import restrictions.

On the other hand, if one considers the estimates with distributed lags, that is to say, in which the hypothesis of partial adjustments is introduced, the coefficient of import lags (M_{t-1}) appears significant at a level of at least 95% in equation (3), which leads us to reject the hypothesis that Brazilian imports adjust to the information relevant to the demand function in a period of one year or less. However, in the specification of equation (9**) the income elasticity and the elasticity of lagged quantum variable become non-significant (5% significance level), which raises doubts about the effective adjustment period of the effective imports to their desired level. Perhaps the inclusion of the capacity variable picks up these short term effects, it not being reasonable to suppose, at least for total imports, that this period is greater than one year. Recent work by Abreu and Horta (1982), tries to shed some additional light on this question, through estimates for disaggregate equations by category of use, suggesting that the lags do not seem to be relevant, except in the case of capital goods.

If we accept the hypothesis of an adjustment greater than a year, it is important to distinguish between short-term elasticities, given by the disequilibrium equations, and long-term elasticities, coming from the equilibrium equations. The income and price elasticities, will be calculated from the short-term elasticities, according to $\gamma\, \alpha_i/1 - (1 - \gamma)$ where γ is the adjustment coefficient and $\gamma\, \alpha_i$ are the short-term income and price elasticities, given by equation (3). If $(1 - \gamma) \neq 0$ then the adjustment of imports to their desired level is not instantaneous, that is to say, there exists an adjustment lag. We can conclude immediately, through

14 It is also important to point out the similarity and independence shown by the elasticities in relation to the real income, potential income and utilization of capacity variables, in equations (6), (9) and (10), given that any two out of the three will imply the third.

certain arithmetic exercises, that high price and income elasticities persist both in the long-term as well as the short-term (see Table 3).

It is useful to compare the empirical results of this research project with those obtained by Lemgruber (1976), Weisskoff (1979) and Khan (1974).

The equation estimated by Lemgruber (1976) given the small number of degrees of freedom, with only a few annual observations, was based on the simplest specification, similar to equation (1). The data utilized were the price and import quantum indices published monthly by *Conjuntura Econômica*, the industrial product index as a variable of economic activity, the general wholesale price index (IPA) as (P) and the annual average rate of exchange (λ). One can note that Lemgruber did not inclued in his estimates an import tariff index. On the other hand, it was necessary to introduce a *dummy* variable for the year 1974, for the same reasons as explained above. One should emphasize, furthermore, that the data related to the price and import quantum indices do not exclude the items petroleum and wheat.

The general model of import demand adopted by Weisskoff (1979), distinguishes besides the income and price elasticities of import demand, a tendency coefficient and a dummy variable designed to show shifts in the function, due to changes in the direction of economic policy.

All the regressions in Weisskoff were also estimated by ordinary squares, with annual data, covering the 1953/70 period.

Khan (1974), besides using the simplest specification in his regressions, also allows for the proces of partial adjustment.

Table 3

Income and Price Elasticity for Short and Long-Term Imports [1]

		Income	Price
Short-Term	Case I	0.5887	—0.9939
	Case II	0.5388	—1.0536
Long-Term	Case I	1.0140	—1.7121
	Case II	0.8990	—1.7578

1 The long-term elasticities were calculated from the short-term elasticities according to the formula:

$$\frac{\gamma\alpha_i}{1-(1-\gamma)}$$

where: $\gamma\alpha_i$ = short-term elasticities
γ = adjustment coefficient

Contrary to Lemgruber (1976) and Weisskoff (1979) Khan's regressions were estimated subject to an auto-regressive process of the first order, through the two stage least squared method. [15]

The period under study was 1951/69, with annual data. All the variables are at 1958 dollar value.

The results of Lemgruber (1976), Weisskoff (1979) and Khan (1974) are shown in Table 4.

With the exception of the estimate of the equilibrium in the Khan model, all the regressions show a high income elasticity of import demand. However, the results insofar as price elasticity is concerned do not seem to respond in the same way. Contrary to Khan's results, the coefficients estimated by Lemgruber and Weisskoff, indicate that although showing a correct sign, Brazilian imports show a low elasticity in relation to price. This difference between the results seems to result from the fact that Lemgruber, Weisskoff and Khan used price and import quantum data, which included petroleum and wheat. Furthermore, they did not incorporate into their relative-price variables, an import tariff index. This omission will also have introduced a certain bias in the estimates.

Contrary to our observations, Khan affirms that the degree of auto-correlation of the residues, reflected by the non-significant auto-correlation coefficient, both in the equilibrium estimates as well as the disequilibrium estimates, would serve as an indication that the quantitative restrictions have had a non-significant role in Brazil's flow of trade. Besides this, Khan finds evidence of a non-significant import lag coefficient and concludes that the adjustment of importation is instantaneous.

On the other hand, Weisskoff's results concerning trend elasticity confirm our own, to the extent that both show evidence of a strong influence of tendencial factors, which Weisskoff interprets as reflecting successful import substitution. The other results cannot be compared, since Lemgruber (1976), Weisskoff (1979) and Khan (1974) restricted their exercises to the simplest specification of the aggregate import demand equation.

5 — Final considerations

The different economic policy measures designed to reduce external imbalances, particularly exchange policy and tariff controls, introduced in the post-war years with varying degrees of intensity,

15 The technique used was that of Sargan (1964).

Table 4

Equation for Import Demand in Brazil: Lemgruber — Weisskoff — Khan [1]

	Constant	Activity Variable GNP	Activity Variable GCF	Price	Time	Dummy	M_{t-1}	RHO	R^2	R^{-2}	DW
Lemgruber	—2,0257	1,4930[a] (27,15)	—	—0,4949[c] (2,02)	—	(dum.74) 0,3294[a] (3.84)	—	—	0,0049	0 9935	1,84
Weisskoff	—5,994	2,333[a] (3,034)	—	—0,374 (2,142)	—0,131[b] (2,840)	(dum.68) 0,260[b] (2,796)	—	—	—	0,823	1,89
	—2,744[b] (2,778)	—	1,759[a] (1,488)	0,262[b] (2,921)	—0,080[a] (5,876)	—	—	—	—	0,896	1,77
Khan	2,533 (1,51)	0,107[b] (2,78)	—	1,688[a] (5,50)	—	—	—	0,378[c] (1,49)	—	—	—
	—0,402 (0,20)	0,153[a] (3,51)	—	—1,315[a] (3,97)	—	—	1,153 (1,22)	0,342[c] (1.97)	—	—	—

1 Aggregate annual data were utilized in the regressions
GNP — Gross National Product GCF — Gross Capital Formation

a) significant at a 99% level;

b) significant at a 95% level;

c) indicates that the hypothesis of auto-correlation can be rejected.

represent successive attempts to equate the real needs for importation and saving, peculiar to the stage of growth and differentiation of the Brazilian economy, to its capacity to generate currency reserves.[16] From this point of view the reserve position has emerged as the controlling factor in guiding domestic economic policy in the import-export field, which becomes translated into a policy of granting privileges to the exporters or of toughening import controls.

Parallel to this the situation of the international financial markets has indicated what proportion of the gap in the real resources can be financed by the inflow of foreign capital. In this way, the absorption of foreign capital allowed the investment levels to remain above that which internal saving alone would provide, and permitted the internal product to continue its path of expansion even when the stimulus coming from the international economy had become inverted in the middle of the last decade. Such an option would not create problems, as long as the process of indebtedness stayed in line with the expansion in trade and the supply of international financial resources, with the continued expansion of exports and real prospects for maintaining the growth in internal income.

This policy of growth *cum* indebtedness which requires the continual expansion of export revenues is now facing considerable difficulties due to the cooling off of world trade and the disorganization of the financial markets caused by events over the last two years.[17] Once again in Brazilian economic history, the way out of the impasse brought by the shortage of foreign reserves cannot be seen only in terms of promoting exports, the expansion of which is conditional on a change in the overall international situation and on the recovery of foreign markets. On the other hand, the inflow of loan capital is dependent on the return to normality on the international financial markets. Once again, policies which deal directly with restricting imports have become the order of the day in economic policy debates

On the other hand, there is a trend in imports towards an ever growing concentration on capital goods and intermediary goods, with a marked decline in imports of consumer goods, both durable and non-durable. Such modifications have resulted from structural changes which have taken place in the national productive system over the last few decades arising from different policies of industrial expansion. In this way, imports have evolved to the

[16] See Dib (1983, Chapter II).

[17] See, for example, Arida (1982, Sections IV and VII) and Malan (1982).

point where they have become relatively incompressible and any additional reductions will tend to cause a considerable impact.

This being the case, the need for more accurate knowledge of the behaviour of imports led us to concentrate our efforts on estimating equations for aggregate import demand in Brazil, which will permit an evolution, in a systematic way, of the principal aggregates which have an influence on it.

The results obtained for the various specifications for the total imported *quantum* (excluding petroleum and wheat) included in the previous section, show that the estimated equations depict with a reasonable degree of accuracy the behaviour of import demand in the last twenty years (1960/80).

The decision to exclude petroleum and wheat was instrumental for obtaining such good econometric results. All the coefficients show expected signs and reasonable accuracy in the estimates, confirming the existence of high income and price elasticity in import demand in Brazil. Besides this, the inclusion of other variables, as a way of differentiating cyclical and tendencial effects on import demand, especially the measure of utilization of capacity, reflects the greater influence of extra-price effects, such as variations in the level of economic activity, in relation to price effects.

It is worth mentioning, however, the precariousness of the utilization of capacity measure, based on the ratio between real product and potential product. The adoption of this procedure to estimate the utilization of capacity presupposes, *caeteris paribus,* that the entrepreneurs and the government create productive capacity at constant rates and independent of the economy's performance.[18] In reality, with the continuation of the recession, it is natural that this measure becomes more precarious, as one cannot know with certainty, at this stage, the effects of the recession on the expansion of productive capacity, through a fall in total investment. It is expected, therefore, that the equation will tend to overestimate the effects of the recession on the decline of the imported *quantum,* and underestimate the consequences of a possible recovery of the economy.

The ideal solution would be to possess a direct measure of the utilization of installed capacity, or to alter the method of calculating the potential product which would incorporate the effects of a prolonged fall in the growth rate of the GNP on the investment decisions of the economic agents. Such considerations, however, go beyond the scope of the present study.

18 See Abreu and Horta (1982).

Thus, although certain reservations can still be made, particularly as to the method of calculating the utilization of capacity variable and its effects on the variation of the imported *quantum*, particularly in prolonged periods of low economic growth rates, it is possible to estimate import demand within the econometric "canons", as long as the key income and price variables are adjusted to take into consideration: a) the distinction between the tendency effects and the cyclical effects in the variation of the level of activity; b) that the relative prices incorporate the tariff policy; and c) that exchange rates are reasonably stable.

It is obvious, from the above, that the reality of the eighties is radically different from that of the two previous decades. The changes that are taking place on the international scene tend to make it quite different from that which prevailed until the mid-seventies. The Brazilian economy should, therefore, adjust as quickly as possible to this new reality, characterized particularly by a slower growth in the inflow of foreign reserves, as a consequence of the negative impact on world trade caused by a succession of exogenous shocks which have exaggerated in a disproportionate way the degree of uncertainty in the international financial and trading markets.

The movement towards greater import control seems then to be inevitable in order to achieve a surplus on current account, especially on the trade balance. Thus, the need to keep imports within the limits imposed by the availability of foreign reserves together with a greater influence of extra-price effects *vis-à-vis* price effects in the determinants of aggregate demand, clearly demonstrates the urgency of harmonizing the current policy of controls with a new industrial policy that defines viable strategies of economic growth compatible with external restrictions.

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Brazil's external balance: an evaluation of the monetary approach *

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This paper discusses the extent to which the "monetary approach" can serve as an explanation of the Brazilian external balance. Recent work by Connolly and Dantas (1979) claims that "The simple models of exchange market pressure tested here perform fairly well in the 1955/75 period and very well during 1962/75, in explaining movements of reserves and the exchange rate". We show that their analysis is flawed in a number of respects but that, when all corrections are applied, their conclusions substantially stand.

In a first part we briefly review the monetary approach. The second section presents empirical evidence and in the final section we comment on the Connolly-Dantas paper.

1 — The monetary approach

The monetary approach to the balance of payments grew out of policy oriented modelling at the IMF (1978) with some ancestry in Dutch theory [Prais (1961)]. The approach received its main emphasis toward the end of the 1960s, particularly associated with Johnson (1972), Mundell (1971) and their students [see Frenkel and Johnson, eds. (1976)]. The particular model to be discussed here is due to Girton and Roper (1977) who estimated it for the case of Canada.

* Originally published in *Pesquisa e Planejamento Econômico*, Rio de Janeiro, *10* (2): 481-502, August 1980.

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The monetary approach recognizes the relationship between the balance sheet of the consolidated banking system, the external balance and monetary aggregates. Denoting net foreign assets, the money stock and domestic credit by *NFA, M* and *DC* respectively, the balance sheet identify of the banking system states:

$$NFA + DC \equiv M \tag{1}$$

where M denotes *all* liabilities of the consolidated banking system. The monetary approach uses this identity, combined with assumptions about the monetary sector and the exchange rate regime, to establish a link between money demand changes, changes in domestic credit and changes in net foreign assets.

Denoting by Δ a change we obtain from (1):

$$\Delta\ NFA \equiv \Delta\ M - \Delta\ DC \qquad (1)'$$

In (1)' we still have an identity. The next step is to convert it into a theory by imposing the assumption that money supply always equals money demand. Denoting the change in money demand by $\Delta\ M^d$ and imposing the assumption $\Delta\ M = \Delta\ M^d$ converts (1)' from an identity into a theory, namely the monetary approach:

$$\Delta\ NFA = \Delta\ M^d - \Delta\ DC \qquad (1)''$$

The monetary approach asserts that under fixed exchange rates changes in money demand increase net foreign assets of the consolidated banking system while credit expansion leads to a precisely offseting loss in net foreign assets. With exchange rates flexible or managed, as we see below, the approach amounts to a statement about changes in net foreign assets and/or appreciation. A rise in domestic credit, for example, leads to an offsetting decline in foreign assets or to exchange depreciation, higher prices and thus higher money demand. These details are spelled out below.

The attraction of the monetary approach is that it offers a very aggregative, simple framework for the analysis of net foreign assets. In terms of complexity it thus compares favorably with the alternative approach that would specify separate equations for exports, imports and capital flows, and in this way build up a model of the external balance and changes in net foreign assets. We now develop the monetary approach and start by expressing (1) in percentage changes:

$$r \equiv \Delta M/M - d \tag{2}$$

where $r \equiv \Delta NFA/M$ and $d \equiv \Delta DC/M$ are used for national convenience.

The next step is to impose monetary equilibrium with *nominal* money demand determined by the price level, real income and the nominal interest rate:

$$M = PY^{\phi} e^{-hi} \tag{3}$$

where P, Y and i denote the level of prices, real income and the nominal interest rate. The functional form is that assumed by Cagan (1956).

Differencing (3) yields:

$$\Delta M/M = p + \phi y - h\Delta i \tag{3)'$$

where lower case letters represent percentage changes; thus $p \equiv \Delta P/P$. Substituting (3)' — the demand determined changes in nominal money — into equation (2) yields:

$$r = p + \phi y - h\Delta i - d \tag{2)'$$

The model is closed by the assumption of purchasing power parity:

$$P = EP^{*} \tag{4}$$

where E is the cruzeiro price of foreign exchange and P^{*} the given foreign price level. Differencing (4) leads to an equation for the domestic rate of inflation:

$$p = p^{*} + e \tag{4)'$$

After substitution in (2)' we obtain the final form of the equation describing changes in reserves and exchange rates:

$$r - e = p^{*} + \phi y - h\Delta i - d \tag{5}$$

The equation implies the following predictions:

i) An increase in external inflation leads to reserve increases or appreciation, one-for-one.

ii) Growth in real income leads to reserve accumulation or appreciation.

iii) An increase in interest rates leads to a reserve outflow or depreciation.

iv) An increase in domestic credit creation leads to an equal reserve decumulation or depreciation.

The theory thus implies sign restrictions on two of the right-handside variables — real income and interest rates — and the tighter restrictions of plus and minus unity on foreign inflation and domestic credit creation respectively.

Empirical work would estimate the equation (5)':

$$r - e = a_0 + a_1 d + a_2 p^* + a_3 y + a_4 \Delta i \qquad (5)$$

and test the restrictions $a_0 = 0$, $-a_1 = a_2 = 1$, $a_3 > 0$, $a_4 < 0$.

Equation (5)' is of course not a test of the monetary approach but rather a *joint* test of four hypothesis: (i) the monetary approach, (ii) continuous money market equilibrium, (iii) the functional form and determinants of money demand and (iv) continuous purchasing power parity. There are further hypotheses introduced, once the choice is made on what are the data counterparts of p^*, y and i and how to estimate the equation.

Equation (5)' is what Girton and Roper (1977) call an "exchange market pressure" formulation of the monetary approach and is the model that Connolly and Dantas (1979) applied, in a modified form, to the Brazilian data. We now proceed to an estimate of (5)'.

2 — Estimates for Brazil 1958/78

The exchange pressure model of the monetary approach was estimated with annual data for Brazil for the period 1958/78. The data are described in detail in the appendix we note here simply that foreign inflation was represented by the US wholesale price index, the alternative cost of holding money, because there is no sufficiently long interest series, was proxied by the Brazilian inflation rate of the general price level. Finally domestic credit creation measures the growth of domestic credit less nonmonetary liabilities of the consolidated banking system.

In Table 1 we report the empirical evidence. First we observe that the model explains a large fraction of the variation in exchange market pressure, but that there remains considerable serial correlation even after first-order correction. The theory is supported in that all coefficients have the expected signs: specifically, a higher rate of credit creation is reflected in reserve losses or depreciation; higher foreign inflation leads to reserve gains or appreciation. A higher growth rate of home income implies increased reserve gains or appreciation while increased, home inflation, by raising the opportunity cost of holding money, gives rise to reserve losses or depreciation.

Not only are the signs in accordance with the theoretical specification, but the parameter estimates also accord substantially with the theory. Thus domestic credit creation shows a coefficient of approximately minus unity; the foreign rate of inflation has a coefficient that is significanly different from unity.

The two equations reported in Table 1 differ in the formulation of the income variable. In the first equation we use the current growth rate of income in the other a three year moving average. The specification of the income variable substantially affects the estimates and the precision of estimates of foreign inflation, real growth and the alternative cost of holding money. The formulation that uses a three year moving average implies a less precise estimate for foreign inflation and a more precise estimate of real income growth and changes in the alternative cost of holding money. Along with a more precise estimate for the exchange market effect of real income growth we obtain, though, an unacceptably high estimate of the income elasticity as being 3.13.

In the following chart we show a graph of actual and predicted values corresponding to equation (1) in Table 1. It is apparent that the equation tracks well the main movements in the exchange market pressure variable. In particular the 1964 episode is well-accounted for by the model.

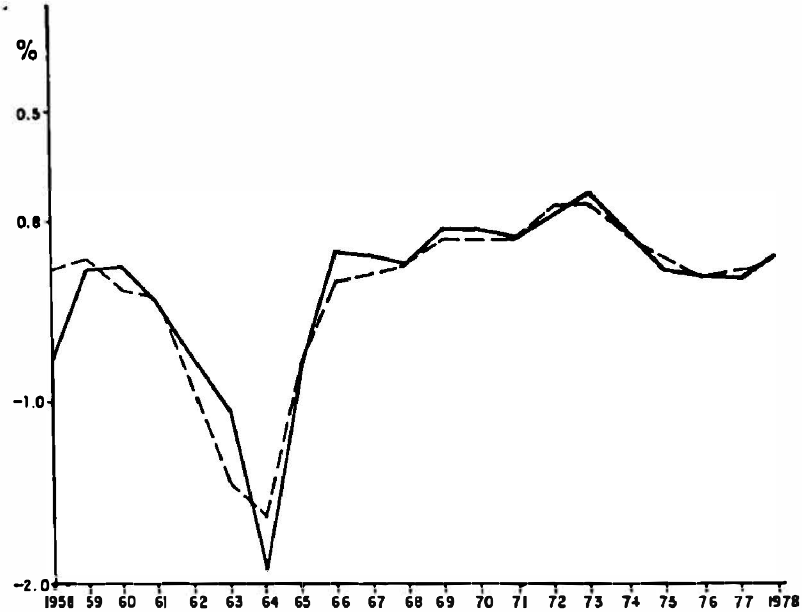
Our conclusion on the empirical evidence is that the monetary approach model describes well the behavior of the exchange

Table 1

$$r - e \quad a_0 + a_1 d + a_2 p^* + a_3 y + a_1 \Delta p$$

1958/78	a_0	a_1	a_2	a_3	a_4	R^2	DW	SER	Rho
1	—.03 (—.17)	—.99 (—7.51)	1.56 (2.27)	1.01 (0.72)	—.46 (—1.74)	.88	1.58	.18	—.41
2	—.18 (—1.20)	—.93 (—9.30)	.77 (1.03)	3.13 (1.90)	—.77 (—2.66)	.91	1.64	.10	—0.47

NOTE: The income growth variable in equation 1 is the current growth rate of real GDP; in equation 2 it is a three- year moving average. All equations estimated with a correction for first-order serial correlation. t-statistics in parenthesis.



pressure variable, although the magnitude and precision of some of the coefficient estimates — income, foreign inflation and the alternative cost of holding money — are not entirely satisfactory.

3 — A comparison with the Connolly-Dantas estimates

In Table 2 we report estimates by Connolly-Dantas (1979). It is apparent that their results, especially for the period 1962/75 are quite splendid. Specifically the coefficients are all significant and they all have magnitudes entirely compatible with the theoretical specification. Moreover, the equation for 1962/75 shows no evidence of serial correlation, after a minor first-order correction. What then accounts for the difference between these results and those we report in Table 1? The differences arise in part from the equation specification, in part from data differences:

i) Connolly-Dantas estimate their equations without a constant. This procedure will tend to raise estimated statistics on the remaining variables.

ii) The equations are estimated without a variable measuring the opportunity cost of holding money. Connolly-Dantas note that they omit the alternative cost of holding money "for simplicity" although it is quite apparent that this amounts to a serious misspecification of money demand, surely incompatible with the monetary approach.

iii) There are data problems, to judge from the detailed description provided in the appendix to Connolly-Dantas. The first concerns the real income series. Connolly-Dantas seem to be using the IMF series which has a break in 1964. Failure to note that break leads to a calculated real income growth rate of 23.3 percent for 1964. It appears that this incorrect number is being used.

iv) The perhaps more serious data problem concerns the series for the growth of domestic credit of the consolidated banking system. This series is constructed, according to their appendix, by taking the difference between the growth rate of M_1 and the growth of net foreign assets, both expressed as a fraction of M_1. That procedure would be correct *only* if the consolidated banking system had only M_1 as a liability. The presence of time deposits, and more importantly of nonmonetary liabilities, implies that domestic credit cannot be calculated in that manner. In our work reported above we used the definition of *net* domestic credit by adjusting the banking system's credit outstanding for all nonmonetary liabilities. [1]

Table 2

The Connolly-Dantas Equation

$$r - e = a_0 + a_1 d + a_2 p^* + a_3 y$$

	a_0	a_1	a_2	a_3	R^2	DW	SER	Rho
1.	195575	—1.01	1.29	1.27	.68	2.22	.81	—.11
		(— 7.42)	(1.27)	(1.26)				
2.	196275	—1.01	1.21	1.46	.91	2.00	.13	—.12
		(—13.09)	(2.04)	(2.45)				

SOURCE: Connolly-Dantas (1979, Table 1).

[1] A further difficulty we encountered was our inability, even reconstructing the Connolly-Dantas data according to their description to reproduce their equations. Our estimates using their data are:

$$r - e = a_0 + a_1 d + a_2 p^* + a_3 y$$

1. 1962/75	—.14	— .95	.12	2.95	.90	2.07	.13	—.10
	(—.87)	(— 6.7)	(.15)	(1.97)				
2. 1962/75		— 1.05	.40	1.80	.89	2.17	.12	—.18
		(—11.5)	(.54)	(2.72)				

4 — Concluding remarks

The problems in estimation, unwarranted constraints and data seem sufficient to cast serious doubt on the quality of the results and the claim that the monetary approach does explain well the evolution of exchange market pressure in Brazil. In view of these doubts our earlier estimates, in Table 1 with a more complete model and better data, and covering a more recent period restore some confidence in the monetary approach. The model does explain the behavior of the exchange market pressure in a satisfactory way, but it is quite clearly not a last word on the issue. Not only are their problems left in matching coefficient estimates with the theory, there are also the serious issues related to the endogeneity of some of the right-hand side variables that have been noted in the literature.

As an exercise in the monetary approach the Brazilian case is of particular interest because it draws attention to the role of nonmonetary liabilities. To use the monetary approach for financial programming in Brazil it would actually be necessary to control *net* domestic credit. But that implies estimates of nonmonetary liabilities ranging from bank debt to import deposits. It is these nonmonetary liabilities, and not only the difficulty of predicting nominal money demand, that make financial programming particularly hard.

Appendix: The data

This appendix describes the data used in estimating the equations in Table 2. The series for domestic credit and net foreign assets, as well as the data on the US price index, were obtained from the IMF, *International Financial Statistics* (*IFS*), Yearbook 1979, and July 1978. The remaining data came from *Conjuntura Econômica* (*CE*) and the *Boletim do Banco Central do Brasil* (*BCB*) as indicated below.

1. *Net Domestic Credit Creation* (d):

Net domestic credit is defined as the difference between domestic credit (line 32 in *IFS*) and nonmonetary liabilities (lines 36b + 37a + 37r in *IFS*). To obtain midyear estimates we averaged the end of year data for the current and the preceding year. The growth rate, as a fraction of lagged money, is defined as:

$$d = \Delta DC/M_{t-1}$$

where M is the money stock (lines 34 + 35 in *IFS*) obtained as an average of the current and preceding and of year data. The revisions are bridged by using the previous series, reported in *IFS* July 1978 through 1972 and the new series in *IFS* 1979 for the remaining years.

2. *Net Foreign Asset Changes* (r):

As net foreign assets we use line 31n less line 36cl in *IFS*. The data are end of year, hence we use the averaging described above. Changes in net foreign assets, as a fraction of lagged money are defined as:

$$r = \Delta NFA / M_{t-1}$$

The series on net foreign assets starts in 1955. Hence our first observation, as defined here is for 1957.

3. *Real Income Growth* (y):

Real income growth is formed as a three year moving average of the growth rates of real GDP. Denoting the real GDP growth rate by $\bar{y}$, the variable is calculated as:

$$y = \frac{1}{3} (\bar{y} + \bar{y}_{-1} + \bar{y}_{-2}) \tag{3}$$

Through 1966 we used the data from *CE*, April 1977. Growth rates from 1966 on were calculated on the basis of the new series series reported in *IFS*, 1979. (The old series corresponds to that reported in *IFS* until 1964, the new series starts in 1965; the break in the series is not indicated). The preliminary estimate of 1979 real growth comes from *BCB*, March 1979.

4. *Foreign Inflation* (p^*):

Foreign inflation is measured by the rate of inflation of the annual average US wholesale price index: *IFS*, 1979, line 63.

5. *Exchange Depreciation* (e):

The exchange rate data are obtained from the annual average exchange rate reported in *CE*, April 1977 and July 1979. The rate of depreciation, e, is the percentage rate of change of the anual average.

6. *Change in the Alternative Cost of Holding Money* (Δp):

The alternative cost of holding money is measured as the rate of inflation of the "general price level". The rate of inflation is calculated from the index reported in *CE*. The variable p is the change in the inflation rate thus calculated. The *CE* issues were April 1977 and July 1979.

	r	e	p^*	d	y	Δp	$\bar{y}$
1957	0.00	0.0188	0.0270	0.27	0.069	−0.057	0.0520
1958	−0.04	0.7081	0.0150	0.29	0.077	−0.012	0.0550
1959	−0.06	0.2104	0.0018	0.37	0.056	0.248	0.0673
1960	−0.04	0.2115	0.0018	0.43	0.097	−0.086	0.0767
1961	0.00	0.4362	−0.0037	0.44	0.103	0.078	0.0853
1962	−0.35	0.4238	0.0018	0.90	0.053	0.146	0.0843
1963	−0.64	0.4883	−0.0037	1.27	0.015	0.238	0.0570
1964	−0.72	1.2029	0.0019	1.48	0.029	0.146	0.0323
1965	−0.35	0.4880	0.0203	1.14	0.027	−0.332	0.0237
1966	0.00	0.1718	0.0344	0.42	0.038	−0.188	0.0313
1967	0.02	0.2012	0.0018	0.33	0.048	−0.097	0.0377
1968	0.04	0.2748	0.0245	0.36	0.112	−0.041	0.0660
1969	0.15	0.1996	0.0392	0.25	0.100	−0.034	0.0867
1970	0.08	0.1272	0.0361	0.22	0.088	−0.010	0.1000
1971	0.07	0.1521	0.0333	0.24	0.133	0.012	0.1070
1972	0.16	0.1224	0.0445	0.08	0.117	−0.034	0.1127
1973	0.21	0.0324	0.1307	0.24	0.140	−0.019	0.1300
1974	0.04	0.1084	0.1883	0.35	0.098	0.136	0.1183
1975	0.06	0.1968	0.0929	0.43	0.056	−0.010	0.0980
1976	0.01	0.3131	0.0460	0.38	0.090	0.136	0.0820
1977	0.02	0.3250	0.0612	0.38	0.047	0.014	0.0650
1978	0.08	0.2776	0.0784	0.39	0.060	−0.040	0.0660

NOTE. For definitions see text $\bar{y}$ is a three-year moving average of y.

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Brazilian wage policiës, 1964/81 *

Livio de Carvalho **

1 — Introduction

An understanding of the wage policies implemented from 1964 to 1979 is crucial to an analysis of the current policy in its historical context. Item 2 will summarize the principles of pre-1979 wage policies. In Item 3, we well comment on the wage policy applied from December 1979 to December 1980. In Item 4, we will analyze the changes recently approved by Congress.

2 — Wage policies: 1964/79

The basic objective of post-1964 economic policies centered on stabilization, a goal which was to orient all instruments of both financial policy and wage policy. [1]

The four basic principles of wage policies during this period were:

a) minimum of one year between adjustments;

b) reposition of the real average wage of the 24 months preceeding the month of adjustment;

c) the average wage should also be readjusted in accordance with the rate of increase of productivity;

Editor's note: Translation not revised by the author.

* Some of the ideas in this article were discussed with several people to whom I am grateful: Paulo Roberto Furtado de Castro, João Agostinho Telles, Hamilton Bizarria, Mauricio Galinkin, Fernando Werneck, Mozarte Foschete, Arno Meyer, Ramonaval Augusto Costa and Ricardo Lima. Any errors and all opinions expressed, of course, are the exclusive authority of the author. I would also like to thank Ana Maria de Resende for her valuable aid in organizing and computing the data. Originally published in *Revista Brasileira de Economia, 36* (1), January-March 1982.

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[1] Ministry of Planning and General Coordination (1965, pp. 15-6).

d) the adjustment should also include half of the inflation projected by the government's financial programming (residual inflation).

These principles were not uniformly applied over time,[2] and the law was modified several times. Nearly all the changes were made to correct distortions which arose in the course of applying the law (inappropriate interpretations or lack of uniformity in the criteria applied) and to broaden its reach.

Four major changes were made to the law:[3]

1) Calculation of the adjustment of the average real salary on the basis of indices published monthly by the government (Executive Order n.º 15; July 29; 1966), in order to eliminate the multiplicity of criteria, since various institutions had been furnishing this information to the Labor Justice Department.

2) In 1968 (Law n.º 5.451; June 12), a coefficient of correlation was introduced to allowfor residual inflation. This modification was intended to prevent the deterioration of salaries in one period from stretching into subsequent periods by using the actual (rather than projected) inflation rate to calculate the average real wage of the twelve months preceeding the adjustment. It should be noted, though, that this change was only made when the difference between the projected and actual inflation rates had dropped to nearly insignificant levels.[4] In addition, the

[2] When it was first introduced, the wage law covered only public sector wages; the scope of its application was extended over time.

[3] For a detailed description of the evolution of wage policies, see Carvalho (1973) e DIEESE (1975).

[4] Residual inflation and increase of the cost-of-living index (Rio de Janeiro): 1964/77.

Period	Residual	Cost of Living (%)
07.64 a 07.65	25	74,3
08.65 a 12.65	0	9,4
01.66 a 07.66	10	28,6
08.66 a 07.67	10	30,4
08.67 a 07.68	15	21,1
08.68 a 07.69	15	21,1
08.69 a 07.70	13	22,1
08.70 a 07.71	12	21,4
08.71 a 07.72	12	16,3
08.72 a 07.73	12	13,0
08.73 a 07.74	12	27,8
08.74 a 07.75	15	29,0
08.75 a 07.76	15	38,9
08.76 a 07.77	15	40,4

indexing mechanism consisted of applying a coefficient to the nominal wages of the last twelve months preceding the adjustment (months 13 to 24)[5] to make them equivalent to the wages that should have been paid if there had been no error in the inflation forecast for the same period. The following adjustment would take into account the nominal wages actually paid during those months (which then had become months 1 to 12 in the formula). This procedure amounted to assuming an error in projected inflation for the first adjustment, only to deny the error in the next adjustment.

3) In 1974 (Law n.º 6.147; November 19), new modifications were introduced to correct two distortions still present in the precedures for calculating the adjustment. The first change was to use the average wage of the twelve months preceding the adjustment, thereby correcting the error mentioned in the last paragraph. Second, in the calculation itself, the rate of increase of productivity is now *multiplied* by the average real wage (which itself is multiplied by the preceding residual inflation), which had not been the case before.[6] Though the difference between the two procedures is small, when productivity is added rather than multiplied the wage raise is diminished.

[5] $$Sr = 1/24 \left\{ \sum_{i=1}^{12} S_i I_i + \sum_{i=13}^{24} S_i I_i C \right\}$$

where Sr = real average wage in the 24 months preceding the adjustment; S_i = wage in month i; I_i = wage index in month i and C = correction coefficient of residual inflation $(1 + R'/2)(1 + R/2)$, where R' and R are, respectively, actual inflation and residual inflation in the 12 months immediately preceding the adjustment.

[6] The previous formula for obtaining the rate of adjustment was:

$$T = S'r/S_{13} + P - 1$$

where P = rate of increase of productivity;

$$Sr = Sr(1 + R/2) \text{ e } Sr = \frac{1}{24} \left\{ \sum_{i=1}^{12} S_i I_i + \sum_{i=13}^{24} S_i I_i C \right\}$$

(See footnote 5 for meaning of the symbols.) Following 1974, the rate of adjustment was obtained directly using the following formula:

$$S/S_t = 1/12 \sum_{i=1}^{12} I_i (1 + R_t/2)(1 + P) \left(\frac{1 + R'_t/2}{1 + R_{t-1}/2} \right)$$

where:

S = new salary following correction.
S_t = salary in each of the preceding 12 months.
R_t = residual inflation in the 12 months preceding the present adjustment.
R_{t-1} = residual inflation used in the previous year's adjustment.
R'_t = actual inflation rate in the 12 months preceding the adjustment.
P = rate of increase of productivity.

4) In 1976,[7] a policy statement signed by three minister recommended the inclusion of a new factor in the calculation of wage adjustments: the correction of the productivity coefficient by a terms-of-trade index to reflect the gains or losses of urgan wage-earners' purchasing power[8] caused by variations in the terms-of-trade indices in Brazil's foreign trade and between the economy's urban and rural sectors.[9]

If the terms of trade were unfavorable to the country or to the urban sector, productivity would be corrected downwards, and if they were favorable, productivity would be adjusted upwards.

This change calls for a few comments. First of all, it reveals that many of the wage law's distortions grew out of the manner in which its principles were being applied, rather than out of the law itself. We thus have a ministerial policy statement modifying the calculation procedures of one of the factors implicit in the law's wage formula. Yet their modification was not illegal, since no law (with the exception of n.º 6.147; November 20, 1974) had presented an explicit formula for calculation of adjustments. Nor was there an explicit presentation of the procedure for computing the elements taken into account by the adjustment formula. Even in the case of productivity, the indicator to be used was never mentioned.[10]

[7] Exposição de Motivos n.º 115, June 2, 1976, signed by the Ministers of the Planning Secretariat, Treasury and Labor and published in the *Diário Oficial* of June 21, 1976.

[8] Wage policies apply basically to urban wage earners.

[9] See, in this respect, DIEESE (1976) e Doellinger (1979, p. 21). For an adjustment carried out in period t, the change implies that the rate of increase of productivity, p, is multiplied by a coefficient $A_{t-1}/A_{t-2} \cdot B_{t-1}/B_{t-2}$, where A_t and B_t are, respectively, the index of the country's terms of trade with other countries and the index of the terms of trade between the economy's rural and urban sectors, both in period t.

[10] Brazil: Productivity increases granted and rate of growth of *per capita* GDP

Period	Productivity for Calculation of Adjustment	Growth of *Per Capita* GDP
1964/65	1,0	0,0
1965/66	0,0	0,3
1966/67	2,0	2,0
1967/68	2,0	2,0
1968/69	2,0	6,0
1969/70	3,0	6,5
1970/71	3,5	6,0
1971/72	3,5	8.9
1972/73	3,5	6,7
1973/74	3,5	8,3
1974/75	4,0	6,7
1975/76	4.0	1.2

This change also made it possible to diminish the productivity index, which at the time was around 4%, well above the 1.2% rate of per capita GDP growth, at a time when inflation was once again on the rise. In any case, the growing inflation rate reduced the relative impact of a cut in the productivity index from, say, 4% to 3%, since the most ungent goal was to recover what had been eaten away by inflation, which was running at about 36%.

The first wage policy law, therefore, went through a number of phases. In the phase 1964/68, during which it was first established, its application was broadened (1964/65) and then consolidated, despite its evident insufficiencies; during the second phase (1968/74), the policy was softened to correct its main distortion (underestimation of future inflation); the third phase (1974/76) brought the correction of the remaining distortions and revealed a certain official concern with the problem of income distribution; and during the fourth phase (1976/79), a resurging of inflation led policy makers to ease up on their wage containment policies. [11]

The philosophy behind the wage policy was that if wages grew no faster than the economy's average productivity, [12] they could be improved without increasing their weight in the aggregate product, that is, without altering prices. [13] The idea is actually very simple, and in the early 1960s was successfully applied in several European countries and in the US. [14]

However the policy's practical results diverged from theoretical forecasts due to the innumerable distortions contained in the law. Indeed, without these distortions, especially the systematic underestimation of residual inflation, the average real minimum wage between adjustments would never have fallen below its

[11] Doellinger (1979, p. 26) gives real rates of wage increases in collective bargaining of 0.6%, —3.0% and 0.9% for the years 1976, 1977 and 1978, respectively.

[12] Actually, this growth would take place due to the initial compression caused by the reposition of the average real wage of the last 24 months before the adjustment and not because of the reposition of the peak.

[13] If we take the share of labor in the product as $S_L = WL/PY$ where W = wage rate, L = employment, P = price level and Y = product, it is easy to see that labor's share remains constant if real wages grow at the same rate as the growth of average labor productivity. By the same token, $P = W - Y/L$, that is, P will remain constant if the growth of nominal wages (W) is equal to the growth of average labor productivity (Y/L).

[14] Some of these policies were mandatory and others not. In the case of the US, this policy was known as "Wage-price guideposts," during the Kennedy government. See, in this respect, *The Economist*, 226(7014):100.1, Dec. 1978.

1965 level.[15] The nominal wage in May 1973 would have been 38.1% higher than that established by decree (Cr$ 431.00, as opposed to Cr$ 312.00).[16]

The real minimum wage fell 42.5% from February 1964 to February 1968. Even if we make the comparison for the months immediately following the adjustment (March 1968), the drop is still 28.3%.

The public civilian employees, the retired and those living on social security (INPS) seem to have been the hardest hit, since their incomes were adjusted at rates even lower than the minimum wage.[17]

As for the wages of professional employees, a study by DIEESE[18] revealed the following behavior: a drop in real wages through 1968, a slight recovery in 1971 and another drop from 1972/74. Overall, from 1964/74, 46% of the professional categories analyzed were set back by 30% or more in their real wages.

Even in industry as a whole, average real wages fell by 9.1% a year from 1964/67, the period when the wage policy was applied most drastically, and only recovered their 1964 levels in 1970.[19] Average wages in industry gres at rate of 8.7% annually from 1967/76. Thus, by 1976, they were 59% higher than in 1964. The reasons for this behavior of average wages in industry, at a time when the real minimum wage was falling substantially, are to be found in the fact that, while the wages of unskilled and smi-skilled labor (always closer to the minimum wage) either fell or grew only slightly, the wages of skilled workers and managers grew at very high rates.[20] Thus the difference between the wages of skilled and unskilled labor grew substantially from 1964/75.[21]

This evidence, and the fact that the inequality of income distribution had increased in the 1960s and throughout the 1970s,[22]

15 With or without the 13th wage (additional wage paid in December), or whether one takes the productivity rate used in the adjustments or the rate of growth of the *per capita* GDP. See, in this respect, Carvalho (1975, pp. 52-60).

16 Carvalho (1975, p. 60).

17 Macedo (1976, pp. 84-7).

18 DIEESE (1975, pp. 57-65).

19 Doellinger (1979, pp. 28-9).

20 During the 1966/72 period, the real wages of unskilled workers fell; skilled workers' wages rose by approximately 2.6% per year and managers' wages by 8.1% per year. See Bacha (1975, pp. 124-58).

21 The ratio between the wages of a companyes general manager and of a construction helper in São Paulo was 65:1 in 1969; 81:1 in 1973 and 90:1 in 1975 (not counting the additional benefits earned by the manager). See Suplicy (1977, p. 11).

22 Malan (1978).

gave rise to heated controversy over the effects of wage policies on income distribution and on the overall role of wage policies in the Brazilian development model.

There were two extreme positions in the debate. Several studies guaranteed that a more generous wage policy (which would have fed inflation and hindered savings) could improve the position of the middle strata on the spectrum of income distribution, but there would be no guarantee that this improvement would take place at the expense of the richest strata. Therefore, the distribution of income could evolve towards a position where losses would result in the share of the poorest 40%. [23] This current also argued that the structural transformations of the Brazilian economy since 1964 had also brought changes in the employment structure, major rural-urban migrations, etc. which led to a temporary increase in inequalities, [24] mostly due to the fact the labor supply's profile of skills did not match the demand for skilled labor, which was growing faster than the educational system could keep up with. [25]

The other argument presented by this current was that a fall in the minimum wage bore no relation to a more unequal distribution of income. One reason was that formal employment relations apply to only one sector of the labor market, leaving rural workers, urban occasional workers, the self-employed and domestic servants outside the reach of wage policies. In addition, though the minimum wage had fallen, the percentage of workers earning the minimum wage had also decreased. [26]

On the other side of the controversy, some economists held that wage policies were directly responsible for the deterioration of income distribution, since the wages of the unskilled and semi-skilled urban workers were the ones hardest hit. [27] A variation of this argument is that the combination of wage containment policies with rapidly growing production (following 1967) increased the companies' profitability and made it possible to pay higher wages to higher-ranking technicians. [28]

23 Simonsen (1975, p. 19).

24 This is the Kuznets effect [Kuznets (1955)] used by Langoni (1973).

25 Langoni (1973).

26 Macedo (1976). Macedo's article is reviewed by Bacha and Taylor (1977). The authors show that the median wage is influenced by the minimum wage; thus the latter is not irrelevant.

27 Fishlow (1972) e Hoffmann (1973, pp. 7-17).

28 Bacha (1975).

The interpretation that compressions of wages had been a deliberate component of the government's economic policy was even supported by official publications and articles by authors identified with official positions. The Government Economic Action Program, for example, by blaming institutional wage hikes above the growth of productivity for the inflation it was trying to fight, said between the lines that earlier governments had been too benevolent in their administration os wage policies.[20]

3 — The wage policy of december 1979

3.1 — The new principles

a) Adjustments were not to be made every six months;

b) Adjustments were linked to the variation of the national consumer price index (INPC)[30] of the six months preceding the month of the readjustment;[31]

c) Adjustments came to be differentiated by wage strata, in relation to the country's highest minimum wage (MW);[32]

d) The rate of growth of productivity was to be negotiated between employers and employees during each professional category's yearly bargaining session, and could not be passed on to prices.

[20] Ministry of Planning and General Coordination (1965, p. 28).

[30] In practice, although a national INPC was never mentioned, this had been done since 1966 when, to standardize criteria, the government began to publish 24 coefficients every month in order to calculate the average real salary over the 24 months preceding the adjustment and, even more clearly, following 1974, when it began the monthly publication of the rate of adjustment of wages for the following month. The difference is that it never made explicit what index was bein used, its methodology, etc. Indeed, the same can be said of all the other components of the adjustment formula prior to Law n.º 6.708.

[31] In other words, wage earners recovered the peak wage aerned in the six months preceding the adjustment, as opposed to the average real wage of the twelve months preceding the adjustment. Actually, there is a two-month lag, that is, the CPI for the adjustments to occur in November take into account the cost-of-living indices from April to September.

[32] Up to 3 MW, 110% of the CPI; the 3-10 MW bracket would receive an adjustment equal to the CPI and those receiveing more than 10 MW would receive 80% of the CPI. In the old wage policy, all wages in a given company were readjusted at the same rate. The original proposal of the Labor Ministry was that the third bracket would cover 10-20 MW and would receive 80% of the CPI, while the fourth — over 20 MW — would receive 50% of the CPI.

Our evaluation of the 1964/79 experience can shed some light on the 1979 reorientation of wage policies. With rising inflation and greater freedoms for the labor movement, several unions had already won anticipated wage adjustments before the expiration of their annual contracts.[34] The application of adjustments on a semester basis to all wage earners was thus an extension of a benefit already achieved by some. In addition, the differentiation of raises by wage levels was aimed at ending the previous tendency of a real decline in the minimum wage and an increase of the lower wage levels far below that of higher wage brackets. It remained to be seen how this would be done without further contributing to the growing inflation rate. It would seem that the policy makers' expectation was that the measure would not be inflationary, since the cost of indexing salaries up to three MW at 10% above the CPI would be taken out of the upper wage brackets, thus maintaining the level of total labor costs and the companies' overall cost structure.

One frequent criticism of the previous wage policy was that productivity growth rates were applied to all sectors and companies indiscriminately, were considered to be very low and lacked an explicit methodology for their calculation. By making them subject to collective bargaining agreements. differences could be established between sectors and companies without affecting inflation, since this part of the raise could not be passed on to prices. It seems to us that the expectation was that this would minimize the importance of the methodological problem. In short, as an attempt to cope with some of the tendencies that critics had attributed to the previous policy, the new law held promise. It remained to be seen to what extent the redistributive and antiinflationary goals would be achieved and to what extent they would conflict with each other. In addition, economic policy measures conceived to achieve one result can generate effects which compromise other objectives and even the specific desired result.

3.2 — The redistributive impact

Even assuming no turnover problem[35] and effective price-control measures, the law's redistributive impact is difficult to analyze,

[33] This was the only negotiable part of the adjustment, and would guarantee a reasonably similar reposition of wages for all sectors.
[34] These anticipated adjustments were compensated for, that is discounted, at each yearly bargaining session.
[35] The possibility that companies which make intensive use of unskilled labor might adapt to the wage policy by increasing labor turnover. See, in this respect, Souza (1980).

even when actually applied. It is true that the distribution of wage income should improve, but this is not enough to say what would happen to income distribution as a whole, whether we look at it as functional income distribution (which would depend on adjustments in employment, on the transfer of income from small business and the behavior of other income sources not affected by wage policies) or as personal income distribution (which would depend on the impact of wage policies on the informal sector, once again on the transfer of income from small business and on the diminishing of higher wages).

To evaluate the impact of the 1979 wage policy, if actually applied, on the wage structure, we have simulated two alternatives for continuous application of the wage policy through the year 2000,[36] on the basis of wage brackets established in relation to the highest minimum wage prevailing in November 1980. The alternatives are two hypothetical projections of inflation (both optimistic): in the first, the six-month CPI variation ending in May 1981 is 30% and falls five percentage points in each of the following semesters[37] until it hits 20%, a rate which holds constant through the following semesters until the year 2000; the second alternative is that the first INPC increase is 35%, falling five percentage points per semester until it hits 10%, and then remains constant through the year 2000.[38]

A look at Tables 1 and 2 shows that:

1) The growth of wages in the lower brackets and the fall of the higher wages is highly sensitive to the rates of variation of the INPC. We could conclude that the growth of the lower wages is directly proportional to the inflation rate, which contradicts what is routinely taken to be the effect of inflation on the lower wage groups. In any case, the increase of lower wages and the compression of higher wage levels would reduce the proportion between the two extremes from 50 to 41.25 already in 1982, to 32.48 in 1985 and finally to 13.32 in the year 2000. In the second bracket down to 19.69 MW in the year 2000.

Since the minimum wage used to establish the adjustment brackets is adjusted at a rate above the INPC, there is an expansion

[36] We have no illusions as to the continuity of application of any given wage policy over such a long term. This is simply one means of illustrating its effects.

[37] May 81, 30%; Nov. 81, 25%; May 82, 20%; Nov. 82, 20%; May 83, 20%, etc., etc.

[38] May 81, 35%; Nov. 81, 30%; May 82, 25%; Nov. 82, 20%; May 83, 15%; Nov. 83, 10%; May 84, 10%, etc., etc.

Table 1

Simulation of the Real Wage Structure Produced by Application of the Present Wage Policy Through the Year 2000, at Nov. 1980 Prices (Alternative 1) *

Wage Brackets in Relation to the Highest Min. Wage of Nov. 1980	Period					
	Nov. 1980	Nov. 1982	Nov. 1985	Nov. 1990	Nov. 1995	Nov. 2000
1	5793,8	6213,9	6935,1	8133,9	9591,0	11320,8
3	17369,4	18731,6	20685,2	24401,8	28788,0	33962,4
5	28911,0	30103,2	32263,1	35979,1	40355,4	45539,8
7	40521,6	41896,2	43840,5	47555,8	51042,2	57116,6
10	57899,0	59253,2	61208,0	64922,4	69309,9	74483,3
12	69165,6	69333,9	70376,2	74023,5	78410,1	83584,8
15	86932,0	84193,7	82501,4	83476,1	87750,5	92925,0
17	98409,6	94100,2	90584,7	89235,5	92600,5	97774,9
20	115776,0	109960,0	102709,0	97873,3	98871,6	106501,6
23	133142,4	123819,8	114935,0	106511,6	105026,3	111331,6
25	144720,0	133726,3	122918,4	112270,5	109129,3	114274,2
30	173664,0	158492,6	143126,0	126667,0	119387,0	121592,5
40	231552,0	208025,3	183541,0	155462,2	139001,9	136198,8
50	289440,0	257557,9	223981,0	184256,6	160417,8	150815,8

*CPI: May 81, 30%; Nov. 81, 25%; May 82 on, 20%. Does not include increase of productivity.

Table 2

Simulation of the Real Wage Structure Produced by Application of the Present Wage Policy Through the Year 2000, at Nov. 1980 Prices (Alternative 2) *

Wage Brackets in Relation to the Highest Min. Wage of Nov. 1980	Period					
	Nov. 1980	Nov. 1982	Nov. 1985	Nov. 1990	Nov. 1995	Nov. 2000
1	5789,8	6299,8	6879,2	7311,0	8003,6	8761,8
3	17368,4	18899,6	20037,7	21933,0	24010,9	26275,4
5	28911,0	30174,0	31616,7	33510,7	35589,7	37863,4
7	40521,6	42040,4	43195,8	45089,4	47166,6	49441,3
10	57889,0	59112,5	60561,4	62454,9	64533,3	66808,2
12	69165,6	69315,8	69016,2	71588,0	73667,4	75942,4
15	86932,0	83917,8	82869,4	82401,2	83204,1	85366,8
17	98409,6	96932,7	91503,2	89583,6	89275,0	90406,1
20	115770,0	109205,1	104455,4	100367,1	98249,5	97977,6
23	133142,4	122777,8	117407,6	111146,8	107221,2	105346,2
25	144720,0	132492,4	128013,3	118333,1	113202,8	110325,1
30	173661,0	156779,6	147820,0	136290,1	128157,2	122772,6
40	231532,0	205354,1	193303,2	172231,0	158033,9	147667,6
50	289410,0	253929,7	233977,2	208163,0	187074,6	172582,6

*CPI: May 81, 35%; Nov. 81, 30%; May 82, 25%; Nov. 82, 20%; May 83, 15%; Nov. 83 on, 10%. Does not include increase of productivity.

of the lower wage brackets. In other words, after a few adjustments, ten MW represent much more than they did at the beginning. On the other hand, since for each readjustment 11.5 MW is the cut-off above which wages are adjusted at rates below the INPC, there is a compression of wages in the intermediate bracket (12-20 MW), which progressively drop into the lower brackets, whose adjustment rates are greater. This fact leads to the observations mentioned in points 3 and 4 below.

3) Applied over time, even wages above 15 MW will enjoy real increases. This because the 11.5 MW cut-off for wages to be readjusted at or above the CPI is static, held constant for all adjustments. In the first simulation, wages equal to 15 MW in November 1980 (Cr$ 86,832.00) fall gradually until the May 1987 adjustment, at which point they have lost 5.2% in real terms, but beginning with the next adjustment (November 1987) they begin to grow until they reach Cr$ 92,925.00 in the year 2000, again of 6.6% in real terms in relation to their value in November 1980.

4) As a result of the effect observed in point 3, over the long term the losses taken by salaries in the 17-20 MW bracket in November 1980 would be negligible. In the forst simulation, a wage earner making 20 MW in November 1980 would lose 2.99% yearly until November 1982, but the loss would fall to 0.42% per year until the year 2000. Even the wages-earners located initially in the 20-30 MW bracket would take small losses in annual terms. Thus, if we include productivity increases in our calculations, say 2.5% per year, the real value of wages in the 20-30 MW bracket in November of wages would be improved through the growth of the lower wages (1-15 MW), the maintenance of the intermediate wage leves (15-30 MW) and a reduction of the highest wages (above 30 MW).

The real degree of these wage increases would also depend on other factors, among them:

a) the relation between the INPE and the inflation rate;

b) the weight given to some specific items, such as food and transportation in the cost-of-living index.

Item *a* involves two aspects. The INPC, which is a national average of cost-of-living indices, may vary from other inflation indices. Secondely, structural changes in the economy may not have been reflected in the cost-of-living indices established by out-of-date family budget surveys.[30]

[30] This point was made by Sabóia (1980).

In fact, although the causes are not very clear, such variations seem to be taking place,[40] following the implementation of the wage law.

One probable explanation, although only partial, lies in the significant changes that have occurred on the economy's price structure, changes that have not been picked up by the cost-of-living indices because the family budget surveys which established these indices' weighting systems were carried out many years ago.[41] Two items whose prices have risen disproportionately in recent times are food and transportation. In a cost-of-living index drawn up today, therefore the weight of these two items would be much greater and, as a logical extension, the variations given by most current cost-of-living indices do not actually reflect the cost of living experience by the consumer.

Item *b* refers to the fact that even when the indices do reflect the changes that have occurred in the price structure, the price rises of some items covered by the cost-of-living index has different weights depending on the income level of the consumers.[42] The National Family Spending Study (ENDEF) carried out by the IBGE for 1974/75 revealed that for up to 3.5 MW of expenses, foods accounts for 50% of total spending; for the 10 to 15 MW bracket, this item only accounts for about 20% of expenses and, for the segment over 30 MW, only 6% of the total.[43] Thus, a 30% hike in food prices in general, if consumption habits do not change, would mean a 15% increase in food expenses for the first group, 6% for the second and a mere 1.8% for the highest income group. This observation is of particular importance considering the fact that food and public service prices have systematically risen faster than the increase in the cost-of-living index.[44]

This and the preceding observations suggest that wage adjustments at rates higher than the CPI are simply compensatory and by no means represent a real 10% increase above the INPC.

40 Compare the INPC, for example, with the wholesale price index (domestic availability) — total and foodstuffs.

41 The most recent survey carried out by the IBGE was in 1975.

42 See Homem de Melo (1979).

43 However, the stratified cost-of-living indices, that is, each income bracket with its own cost-of-living index, showed no significant variations over the 1971-79 period. See DIEESE (1979, p. 11).

44 Compare the variations of the components of the cost-of-living index published by *Conjuntura Econômica*, for the period following 1977.

3.5 — The impact on company payrolls

To discuss the inflationary impact of wage policies we must look at the impact of the adjustments established by law on the payrolls of companies. Sarlier studies have shown that this impact varies according to the size of the establishments.[45] The government's RAIS study for 1976 shows that there is indeed a sharp decrease in the share of lower wage-earners as the size of the establishment increases. This decrease is accompanied by an increase in the share of the second wage bracket (3-10 MW), which is also affected by adjustments greater than the INPC. This softens the differentiation of the impact of the wage adjustments. To measure this impact, we have used the RAIS 1976 data to calculate the average wages for each wage bracket.[46] Using these averages, then, we calculated — for each wage bracket — the ratio between the the rate of adjustment and the INPC, and finally pondered each ratio by the share of each wage bracket in the total payroll of companies grouped into different size categories. The final results are shown in Table 3 and the intermediate calculations on Tables 2 and 5, in the appendix.

Table 3

Wage Policy (Nov. 79 — Dec. 80): Impact of Wage Adjustments on Payrolls (in Relation to INPC) by Size of Establishment — Industry, Commerce, Services

Size of Establishment (No. of Employees)	Impact of Wage Adjustments on Payrolls (in Relation to INPC)		
	Industry	Commerce	Services
Micro (0—5)	1,0697	1,0860	1,0739
Small (5—20)	1,0645	1,0718	1,0518
Medium (20—250)	1,0454	1,0463	1,0306
Large (250—1000)	1,0326	1,0200	1,0197
Very Large (1000 or more)	1,0214	1,0409	1,0170
Total	1,0353	1,0540	1,0290

SOURCE: Raw data, RAIS, 76; processed data, Tables 1 and 2 of appendix.

[45] Camargo (1980).

[46] The averages were obtained by pondering the wage brackets presented in the raw data by the weight of employment in each size group, before aggregating the companies by the size groups used in this study. In other words, using

Table 3 indicates that there is in fact a greater impact on the smaller establishments.[47] However the impact is not extremely different from that obtaining in the larger establishments.[48] With a 40% CPI, the payrolls of the smallest establishments would rise, respectively, by 42.79%, 43.44% and 42.96% in industry, commerce and services, while the corresponding adjustments for the largest establishments would by, respectively, 40.86%, 41.64% and 40.68%. In other words, the difference in the average rate of adjustment for the two extreme size groups would be 1.93% for industry, 1.8% for commerce and 2.28% in services.

It would seem that the impact is as differentiated by sector as it is by size. This is even visible at the two-digit level. And if we disaggregate the data to a three-digit level, the effect is even clearer. For example, while 55.06% of exployees in the pharmaceutical and veterinary industry earn less than 3 MW, 82.32% of food industry workers, 41.03% of financial institution employees and 92.69% of hotel and restaurant service workers fall into this same wage bracket. This indicates that inter-sectoral differences can often be even more significant than size differences. To illustrate this possibility, we have classified the companies of the clothing, footwear and softgoods sector and of the motor vehicles

the wage brackets from 3-4, 4-6, 6-8, 8-10, 10-20, 20-30 and 30 and more MW, and weighting them by the employment in each size groups presented in the raw data (less than 1 employee, 1-5, 5-10, 10-20, 20-50, 50-100, 100-250, 250-500, 500-1000, 1000 and more employees), we were able to arrive at the averages for the wage brackets from 0-3, 3-10, 10-20 and 20 and more MW in each of the five size groups we have selected. In the highest wage bracket presented in the raw data (30 and more MW), we have assumed that the highest wage was 50 MW.'

47 Since in all size groups there is a certain percentage of undeclared wages, our procedure implicitly implies dividing the share of undeclared wages among the wage brackets proportionally to the weight of each bracket. To the extent that the undeclared wages fall into the higher wage brackets, the payroll impact would be less than what is shown in Table 3. In other words, if we assume that the sum of the share of each bracket is not equal to the total, the sum of our weights to calculate the average is differente from one. In Table 4 of the appendix, we have calculated the impact placing the undeclared wages in the highest wage bracket, for the sake of comparison. Table 3 thus gives the maximum impact and Table 4 the minimum. Table 4 shows that this procedure significantly changes the results. However, it is best to recall that the shares of the highest wages brackets (20 or more) in the payrolls become absurdly high, ranging from 32.4% to 41.5% in industry; 31.7% to 62.9% in commerce and 33.2% to 43.1% in services. These results seem to indicate that not all undeclared wages fall into the highest wage bracket and that the most reasonable procedure would be to divide the undeclared wages proportionately among all wage brackets, when obtaining the wage averages.

48 Especially if we consider that wages are a small share of total costs.

sector by size and calculated the impact of the wage policy on each using the same method as in Table 3.

Table 4 shows that the impact of the wage policy is differentiated at least as much by sector as it is by size. Depending on the sector, the size differences are much less significant than the sectoral difference. With a 40% INPC, the smallest companies in the clothing sector would have to increase their payrolls by 43.63%, while the largest companies in the same sector would see their payrolls grow by 42.98% (a difference of only 0.65%). Yet the adjustment for the largest companies in the motor vehicle sector would be 40.86% (2.12% lower than the same sized companies in the clothing sector). It should be explicitly noted that, if this wage policy were actually applied in the long term, and if the companies were unable to pass the adjustments on to prices and/or make lower adjustments (through turnover), the difference in impact between small and large employers and between more and less labor-intensive sectors would be an incentive for the adoption of increasingly capital-intensive technologies requiring skilled labor, thus creating one more complicating factor in the country's already discouraging prospects in terms of employment generation.

Table 4

Wage Policy (Nov. 79 — Dec. 80): Impact of Wage Adjustments on Payrolls (in Relation to INPC) by Size of Establishment — Clothing, Footwear and Softgoods and Motor Vehicles Sectors

Size of Establishment (No. of Employees)	Payroll Impact in Relation to INPC	
	Clothing, Footwear and Softgoods	Motor Vehicles
Micro (0–5)	1,0908	1,0400
Small (5–20)	1,0907	1,0526
Medium (20—250)	1,0774	1,0444
Large (250—1000)	1,0676	1,0202
Very Large (1000 or more)	1,0622	1,0185
Total	1,0746	1,0216

SOURCE: Raw data, RAIS 76; processed data, Tables 2 and 6 of appendix.

3.4 — Regional cost-of-living differences and wage policies

If there continue to be systematic regional disparities — with some metropolitan regions' cost-of-living indices above and others below the national INPC average — the wage adjustment process will just as systematically grant adjustments above the cost-of-living indices in some regions and below the same indices in others.
In this case, depending on the magnitude of the difference, even the wages indexed by a 1,10 INPC coefficient may not be recovering the purchasing power they have lost. And this phenomenon could take on extremely negative connotations from the point of view of regional imbalances, if the fastest rising cost-of-living indices were precisely those of the less developed regions, where wages are lowest.

Table 5 shows that Fortaleza and São Paulo's cost-of-living indices were systematically lower than the national INPC; thus the wage adjustments for the lower wage brackets were always above the cost-of-living index. Recife, Belo Horizonte and Rio de Janeiro, on the other hand, had cost-of-living increases systematically higher than the CPI. For them, even the wages adjusted by 110%

Table 5

CPI and 6-Month Cost-of-Living Indices in Some Metropolitan Areas — Nov. 79 — Sept. 80

	1979		1980								
	Nov.	Dec.	Jan.	Feb.	Mar.	Apr.	May.	Jun.	Jul.	Ago.	Sep.
INPC	26,6	28,2	33,2	38,7	40,9	39,9	37,7	37,0	36,8	34,4	33,4
Belém	23,7	28,4	35,1	36,2	39,0	39,1	40,6	36,2	33,7	30,9	39,3
Fortaleza	25,3	25,7	30,1	36,8	36,8	39,8	37,7	39,1	38,2	35,7	32,3
Recife	25,3	30,9	36,0	40,8	47,6	47,8	43,9	37,4	37,5	36,4	33,6
B. Horizonte	28,0	31,0	40,0	43,8	40,3	42,7	40,4	37,9	33,0	32,5	34,3
R. Janeiro	27,7	30,0	35,0	41,4	44,8	43,7	40,0	38,3	30,5	36,4	33,2
São Paulo	25,2	24,2	28,5	34,2	35,1	34,3	34,0	36,0	37,6	33,6	33,7
Salvador	28,0	20,8	35,3	42,2	41,4	41,0	37,3	34,3	31,8	25,9	29,8
Curitiba	32,0	33,1	41,3	43,5	42,1	37,7	30,8	29,2	28,2	30,0	32,0
Porto Alegre	28,9	30,3	37,0	39,7	43,1	42,1	39,1	34,5	31,9	32,0	32,4
Brasília	25,9	22,0	26,4	36,8	39,2	37,6	40,2	43,0	42,2	30,8	38,6

SOURCE: IBGE.

40 Some regions not included in the national INPC calculations have had cost-of-living indices much higher than the INPC. In Manaus, for example, the six-month cost-of-living variations from November 1979 to September 1980 were, respectively: 30.1, 29.9, 32.8, 38.5, 43.3, 42.9, 43.1, 51.2, 53.2 and 52.3.

of the INPC in some cases took real losses.[49] Nevertheless, there seem to have been metropolitan regions that gained and others that lost both in the richer and in the poorer regions.

4 — The new wage policy

4.1 — Introduction

Less than a year after the December 1979 wage policy came into effect, the government sent a bill to Congress to modify it.[50] While maintaining all the principles of the previous policy, it modified the structure of the wage brackets (in relation to the highest minimum wage) on which the various INPC coefficients are applied:

Up to 3 MW — 110% of the INPC
3 — 10 MW — 100% of the INPC
10 — 15 MW — 80% of the INPC
15 — 20 MW — 50% of the INPC
Over 20 MW — adjustment to be negotiated between employer and employees.[51]

According to the government's economic policy makers, one justification for the proposal was based on the high turnover rate occurring among highly skilled workers. This justification, however, does not seem to be true, since if one of the objectives of both the present wage policy and its predecessor is to contain in the turnover of highly skilled labor would be right on target.

Another more significant justification refers to the impact of the wage policy on the state-owned companies' payrolls and, in a broader sense, to the conflict between anti-inflation and wage policies in terms of the largest state-owned companies.

Stated-owned companies occupy a leading role in the Brazilian economy and predominate in several key sectors among them basic inputs and the energy industry. As part of the government's antiinflation policies, policy makers have gove beyong traditional measures (credit, government spending, monetary emissions, etc.)

50 This proposal, an Executive Order, was approved by the fact that Congress had not been able to vote on it within the established time limit, shortly before the end of the legislative session of 1980.

51 In other words, for those earning more than 15 MW in November 1980 there was no change in the wage policy.

to using public companies as instruments in the fight against inflation (due to their leading role in the economy). In a general attempt to limit these companies' expansion stricter controls were placed on their investments and prices. But, dependind on the stiffness of the price controls, there came to be a gap between the nominal growth of their revenues and of their labor costs, with negative effects on the companies' profitability. Since these companies have a high percentage of their employees in the upper wage brackets and, moreover, since the specific objective of the wage policy is to grant higher raises to the lower wage brackets, the profitability problem could be reconciled — that is, total labor costs could be contained — by means of an even more drastic containment of the higher wages. This seems to be a plausible explanation; however it is hard to say to what extent it was a determining factor. The compression of higher wages has now become much more drastic. To give an idea of just how fast they are to be compressed, suffice it to saw that in just the first two years (through November 1982) the losses range from 6.3%, for the 17 MW bracket, to 40.4% for the 50 MW bracket (with a higher INPC at the beginning of the period, which is the alternative closest to reality).

Real wage losses in the 15-20 MW bracket would run between 3-5% per year at the beginning, but would decrease as these wages dropeed into lower brackets, to between 0.2-1%. These losses would therefore be compensated by the inclusion of productivity increases. This compensation would be only partial for the 20-25 MW bracket and barely significant for the higher wage brackets, given the magnitude of the losses.

4.2 — The redistributive impact

The same observations made in item 3.2 still hold for the new law, in reference to the difficulty of determining the redistributive impact in a general sense. It is easier to try to determine the results in terms of the wage structure, if the new wage policy is applied continuously over a long period of time. We have therefore simulated its application through the year 2000, using the same procedures as in item 3.2. The real salaries resulting from this exercise are shown in Tables 6 and 7.

These tables show the continuing validity of our earlier observation, that the increase of the lower wages and the decline of the higher wages are highly sensitive to variations in the CPI.

Table 6

Simulation of the Real Wage Structure Produced by Application of the New Wage Policy Through the Year 2000, at Nov. 1980 Prices (Alternative 1) *

Wage Brackets in Relation to the Highest Min. Wage of Nov. 1980	Period					
	Nov. 1980	Nov. 1982	Nov. 1985	Nov. 1990	Nov. 1995	Nov. 2000
17	98409,6	92709,0	86216,7	88123,1	90075,1	95240,5
20	115776,0	104326,9	97863,3	94420,5	96411,6	101571,0
23	133142,4	113890,7	103731,6	98500,0	99382,2	104337,1
25	144720,0	119411,8	107007,3	100774,2	100938,5	105703,0
30	173664,0	131794,0	113960,0	105030,5	103971,0	108216,0
40	231552,0	156532,5	125014,8	111016,6	108236,0	111477,3
50	289440,0	181271,0	134078,3	115378,7	111343,8	113706,0

*CPI: May 81, 30%; Nov. 81, 25%; May 82 on, 20% does not include increase of productivity. Up to 15 MW, real wages are the same as in Table 1.

Table 7

Simulation of the Real Wage Structure Produced by Application of the New Wage Policy Through the Year 2000, at Nov. 1980 Prices (Alternative 2) *

Wage Brackets in Relation to the Highest Min. Wage of Nov. 1980	Period					
	Nov. 1980	Nov. 1982	Nov. 1985	Nov. 1990	Nov. 1995	Nov. 2000
17	98409,6	92227,5	90254,1	88548,1	88410,7	89689,0
20	115776,0	103256,6	99308,2	96083,4	94682,9	94909,7
23	133142,4	112161,1	105906,6	101256,6	98088,9	98493,9
25	144720,0	117294,4	109710,6	103980,0	101256,5	100381,4
30	173664,0	128589,6	117946,1	109387,3	105756,6	104127,1
40	231552,0	150578,2	131372,1	117817,9	112226,0	109512,8
50	289440,0	172560,8	143246,6	124745,8	116072,4	113462,8

*CPI: May 81, 35%; Nov. 81, 30%; May 82, 25%; Nov. 82, 20%; May 83, 15%; Nov. 83 on, 10%. Does not include increase of productivity.
Up to 15 MW (Nov. 80), real wages are the same as Table 2.

4.3 — The impact of payrolls

We have adopted the same procedure as in item 3.3 to determine the differential impact of the new wage policy by size and by sector, using the 1976 RAIS data. The final results are shown in Table 8, and the intermediate calculations in Table 3 of the appendix. These calculations are simply an approximation, since the 1976 RAIS data we have used [52] aggregate wages between 10 and 20 MW in a single stratum. For our calculations we placed 2/3 of this stratum in the 10-15 MW group of total employees in each size category, and the other third in the 15-20 MW group.

Table 8 shows that the effects of the new wage policy, as were those of the previous policy, are felt strongest by the payrolls of the smaller firms. With the CPI at 40%, the payrolls of the smallest establishments would be adjusted, respectively, by 41.92%, 43.19% and 43.35% in industry, commerce and services, while in the largest establishments the respective adjustments would be, respectively, 39.58%, 41.03% and 30.38%. The new wage policy produces differences in impact on the same order of magnitude between the smallest and largest establishments (previously, the differences in average wage adjustments varied between 1.8% and 2.3%; they now vary from 2.2% in commerce to 2.8% in services). The difference lies in the fact that the new wage policy makes the salary adjustments — both as a total (industry, commerce and services together) and for all the medium to large companies — lower than the CPI.

Comparing Table 8 to Table 3, we find that the new wage policy will have very little effect on the magnitude of average adjustments in comparison to its predecessor, at least if we lood at the aggregate data. [53] And it is not hard to see why. Bot the smallest and the largest establishments have 90% of their employees in the first two wage brackets (0-3 MW and 3-10 MW). Since these are precisely the brackets granted adjustments above the INPC, the results for each category should not differ drastically.

As was mentioned in reference to the previous wage policy, intersectoral differences under the new wage policy can often be greater than the differentiation by size of establishments. Using the same procedures as in item 3.3, we have calculated the new policy's impact on the payrolls of the clothing, footwear and softgoods sector and of the motor vehicles sector.

52 Ministry of Labor (1980*a*).

53 For the largest companies which make intensive use of highly skilled labor, the situation may be different.

Table 8

New Wage Policy: Impact of Wage Adjustments on Payrolls (in Relation to CPI) by Size of Establishment — Industry, Commerce, Services

Size of Establishment (No. of Employees)	Payroll Impact in Relation to CPI		
	Industry	Commerce	Services
Micro (0—5)	1,0481	1,0797	1,0587
Small (5—20)	1,0439	1,0606	1,0099
Medium (20 - 250)	1,0092	1,0158	0.9892
Large (250—1000)	0,9823	0,9490	0,9598
Very Large (1000 or more)	0,9656	1,0023	0,9651
Total	1,9895	1,0259	0,9845

SOURCE: Raw data, RAIS 76; processed data, Table 1 and 3 of appendix.

Table 9, as did Table 4, indicates that the difference in impact by size is less significant than the sectoral differentiation, especially for the large and very large establishments. By the same token, a comparison between the results of Tables 4 and 9 shows that the new wage policy will make very little difference in the total magnitude of average adjustments in comparison to the previous policy — whether by sector or by size of establishment.

As for the drastic compression of higher wages which could result from the new wage policy, we should consider the following points:

a) Private companies will hardly be able to diminish the wages of their workers whose skills are scarce on the market. As for non-scarce skilled labor, through a growth in the turnover rate. The new wage policy will therefore be of greater relevance for public companies, which have a sector of employees whose wages would be lower if they were not set institutionally.

b) If a drastic and generalized compression of higher wages were actually implemented as a medium — and long-range policy measure, this could seriously discourage the formation of human resources.

c) The compression of the highest wage levels, in the form described in item *b*, would undoubtedly lead the most specialized sectors of the labor to organize in defense of their interests through unions, the media, etc.

Table 9

New Wage Policy: Impact of Wage Adjustments on Payrolls (in Relation to INPC) by Size of Establishment — Clothing, Footwear and Softgoods and Motor Vehicles Sector

Size of Establishment (No. of Employees)	Payroll Impact in Relation to CPI	
	Clothing, Footwear and Softgoods	Motor Vehicles
Micro (0—5)	1,0845	0,9972
Small (5 – 20)	1,0888	1,0282
Medium (20 – 250)	1,0651	1,0132
Large (250 — 1000)	1,0459	0,9639
Very large (1000 or more)	1,0412	0,9712
Total	1,0597	0,9746

SOURCE: raw data, RAIX 76; processed data, Tables 3 and 6 of appendix.

5 — Final considerations

While it is not possible to precisely determine the redistributive effects of wage policies, they can certainly bring a more equal distribution of wage income, when applied over the long term. This redistributive effect would take place without wages contributing to inflation, since the inflationary impact of the present policy is very reduced. At the same time, the fact that adjustments take place every six months makes it very difficult for wage policies to be used, as they seem to have been in the past, as an easy road to containing inflation.

For the redistributive effects to take hold, however, a more comprehensive approach must be adopted, going beyond the scope of wage policies to broader changes required for the attainment of overall income redistribution:

a) heavier taxation of non-wage earnings above a certain level;

b) measures to inhibit the turnover of the labor force as a means of getting around the wage law;

c) economic measures to guarantee lower food prices;

d) updating of family budget surveys, in order to assure that cost-of-living indices reflect structural changes in the economy;

e) regionalization of the consumer price index used as a basis for wage adjustments, in order to minimize the difference between local price indices and those used for wage adjustments.

Appendix

Table 1

Brasil: Percentage of Each Wage Group (in Relation to Highest Minimum Wage) by Size of Establlishment

Size of Establishment (by n.º of Employees)	Sector of the Economy											
	Industry (%)				Commerce (%)				Services (%)			
	Wage Brackets				Wage Brackets				Wage Brackets			
	0—3	3—10	10—20	20 or More	0—3	3—10	10—20	20 or More	0—3	3—10	10—20	20 or More
Micro (0—5)	89,04	7,85	0,67	0,34	92,66	5,05	0,26	0,08	90,14	6,74	0,69	0,22
Small (5—20)	84,22	11,68	1,04	0,39	85,53	9,77	0,87	0,21	78,65	14,63	2,07	0,53
Medium (20—250)	77,37	15,92	2,09	0,87	74,30	17,34	2,26	0,76	69,28	20,75	3,68	1,24
Large (250—100)	72,23	19,12	3,85	1,41	72,00	16,24	3,10	1,86	69,68	19,44	4,21	1,87
Very Large (1000 or More)	59,18	30,71	4,58	2,08	70,63	14,66	2,28	1,00	56,53	31,18	6,12	2,14
Total	72,03	20,13	2,84	1,28	80,72	12,68	1,56	0,58	69,88	20.54	3,78	1,35

SOURCE OF RAW DATA: RAIS 76.

OBS.: Totals are not equal to 100% because for all size groups there is a percentage of undeclared salaries.

Table 2a

Average Wages (in Multiples of Highest Minimum Wage) and the Ratio Between Rate of Adjustment/INPC, to Discern Impact of Wage Adjustments on Payrolls — by Size of Establishment

Sector and Size of Establishment					Average Wage in Multiples of MW	Ratio Between Adjustment Rate/CPI for Each Wage Bracket in Multiples of MW			
	0—3	3—10	10—20	20 or More		0—3	3—10	10—20	20 or More
Industry									
Micro (0—5)	1,581	5,027	15,0	31,094	3,090	1,10	1,0597	0,9533	0,8740
Small (5—20)	1,290	4,632	15,0	31,716	1,757	1,10	1,0648	0,9533	0,8725
Medium (20—250)	1,380	4,748	15,0	31,118	2,048	1,10	1,0632	0,9533	0,8739
Large (250—1000)	1,504	4,977	15,0	31,120	2,639	1,10	1,0603	0,9533	0,8739
Very large (1000 or more)	1,607	5,065	15,0	31,246	3,135	1,10	1,0592	0,9533	0,8736
Cloth., Foot., Softg.	1,795	5,071	15,0	30,986	4,092	1,10	1,0592	0,9533	0,8742
Micro	1,339	4,824	15,0	31,522	1,733	1,10	1,0622	0,9533	0,9730
Small	1,281	4,389	15,0	31,667	1,433	1,10	1,0684	0,9533	0,8726
Medium	1,288	4,557	15,0	29,286	1,482	1,10	1,0658	0,9533	0.8785
Large	1,332	4,814	15,0	31,900	1,698	1,10	1,0623	0,9533	0,8721
Very Large	1,314	4,873	15,0	31,087	1,772	1,10	1,0616	0,9533	0,8740
	1,488	4,978	16,0	31,071	2,150	1,10	1,0603	0,9533	0,8740
Motor Vehicles	1,918	5,093	15,0	30,545	4,683	1,10	1,0589	0,9533	0,8753
Micro	1,126	4,959	15,0	32,017	2,141	1,10	1,0605	0,9533	0,8718
Small	1,524	4,680	15,0	30,611	2,589	1,10	1,0641	0,9533	0,8751
Medium	1,951	4,940	15,0	30,579	3,366	1,10	1,0607	0,9533	0,8752
Large	1,858	5,173	15,0	30,617	4,140	1,10	1,0580	0,9533	0,8751
Very Large	2,086	5,100	15,0	30,500	5,287	1,10	1,0588	0,9533	0,8754

Table 2b

Average Wages (in Multiples of Highest Minimum Wage) and the Ratio Between Rate of Adjustment/ INPC, to Discern Impact of Wage Adjustments on Payrolls — by Size of Establishment

Sector and Size of Establishment	Average Per Wage Bracket (in Multiples of MW)				Average Wage in Multiples of MW	Ratio Between Adjustment Rate/CPI For Each Wage Bracket in Multiples of MW			
	0—3	3—10	10—20	20 or More		0—3	3—10	10—20	20 or More
Commerce	1,384	4,920	15,0	30,948	2,255	1,10	1,0610	0,9533	0,8743
Micro	1,204	4,653	15,0	31,818	1,443	1,10	1,0645	0,9533	0,8723
Small	1,336	4.814	15,0	30,000	1,874	1,10	1,0623	0.9533	0,8767
Medium	1,487	4,986	15,0	30,625	2,685	1,10	1,0602	0,9533	0,8751
Large	1,513	5,094	15,0	32,063	3,195	1,10	1,0589	0,9533	0,8717
Very Large	1,571	5,060	15,0	29,950	2,815	1,10	1,0593	0,9533	0,8768
Services	1,535	5,201	15,0	30,667	3,267	1,10	1,0577	0,9533	0,8750
Micro	1,270	4,694	15,0	32,059	1,672	1,10	1,0639	0,9533	0,8717
Small	1,411	5,095	15,0	30,346	2,427	1,10	1,0589	0,9533	0,8758
Medium	1,518	5,256	15,0	30.650	3,238	1,10	1,0571	0,9533	0,8751
Large	1,573	5,191	15,0	31,088	3,485	1,10	1,0578	0,9533	0,8740
Very Large	1,819	5.209	15,0	30,537	4,401	1,10	1,0576	0,9533	0,8753

SOURCE OF ORIGINAL DATA: RAIS 76.

OBS: a) The average were obtained by weighting each bracket in the raw data by the share in employment of each size group; the establishments were the aggregated in the five size groups used in this study.

b) Maximum wage in the highest wage bracket: 50 MW.

Table 3a

Average Wages (in Multiples of Highest Minimum Wage) and the Ratio Between Rate of Adjustment/INPC, to Discern Impact of Wage Adjustments on Payrolls — by Size of Establishment

Sector and Size of Establishment	Average Per Wage Bracket (in Multiples of MW)*		Share of Employment of Size Groups**		Adjustment/CPI per Wage Bracket (x MW)*			% of Payroll per Wage Bracket (x MW)	
	10—15	15—20	10—15	15—20	10—15	15—20	20 or More	10—15	15—20
Industry	12,5***	17,2***	1,89	0,95	0,984***	0,8886***	0,3403	7,65	5,38
Micro (0—5)			0,45	0,22			0,5297	3,20	2,19
Small (5—20)			0,69	0,35			0,5399	4,21	2,99
Medium (20—250)			1,39	0,70			0,5398	6,58	4,74
Large (250—1.000)			1,90	0,95			0,5377	7,58	5,30
Very Large (1.000 or More)			3,05	1,53			0,5422	9,32	6,54
Cloth., Foot., Softg.	12,5***	17,5***	0,44	0,22	0,984***	0,8886***	0,5330	3,17	2,22
Micro (0—5)			0,07	0,03			0,5305	0,61	0,37
Small			0,13	0,06			0,5337	1,10	0,71
Medium			0,37	0,19			0,5267	2,72	1,96
Large			0,57	0,28			0,5404	4,02	2,77
Very Large			0,95	0,48			0,5407	5,52	3,91
Motor Vehicles	12,5***	17,5***	3,78	1,89	0,984***	0,8886***	0,5500	10,09	7,06
Micro			1,33	0,66			0,5247	7,77	5,40
Small			1,45	0,73			0,5488	7,00	4,93
Medium			2,09	1,04			0,5494	7,76	5,43
Large			1,29	1,64			0,5487	9,93	6,93
Very Large			4,43	2,22			0,5508	10,47	7,35

Table 3b

Average Wages (in Multiples of Highest Minimum Wage) and the Ratio Between Rate of Adjustment/ INPC, to Discern Impact of Wage Adjustments on Payrolls — by Size of Establishment

Sector and Size of Establishment	Average Per Wage Bracket (in Multiples of MW)*		Share of Employment of Size Groups**		Adjustment/CPI per Wage Bracket (x MW)*			% of Payroll per Wage Bracket (x MW)	
	10—15	15—20	10—15	15—20	10—15	15—20	20 or More	10—15	15—20
Commerce	12,5***	17,5***	1,04	0,52	0,984***	0,8886***	0,5428	5,77	4,04
Micro			0,17	0,09			0,5280	1,47	1,09
Small			0,58	0,29			0,5600	3,87	2,71
Medium			1,51	0,75			0,5486	7,03	4,89
Large			2,07	1,03			0,5240	8,10	5,64
Very Large			1,52	0,76			0,5609	6,75	4,73
Services	12,5***	17,5***	2,52	1,26	0,984***	0,8886**	0,5478	9,64	6,75
Micro			0,46	0,23			0,5240	3,44	2,41
Small			1,38	0,69			0,5536	7,11	4,98
Medium			2,45	1,33			0,5487	9,46	6,65
Large			2,81	1,40			0,5404	10,08	7,03
Very Large			4,08	2,04			0,5502	11,59	8.11

SOURCE OF ORIGINAL DATA: RAIS 76.

*For other wage brackets, see Table 2 of this appendix.

**The shares of other wage brackets are given in Table 1 of this appendix.

***Values are equal for all size groups, due to aggregation of original data.

Table 4

*Minimum Impact * of Wage Adjustments on Payrolls (in Relation to INPC), by Size of Establishment and by Sectors*

Size of Establishment (No. of Emp.)	Payroll Impact of Wage Adjustments					
	1979 Wage Policy			New Wage Policy		
	Industry	Commerce	Services	Industry	Commerce	Services
Micro (0—5)	1,0145	1,0206	1,0128	0,9030	0,9112	0,8906
Small (5—20)	1,0085	0,9985	0,9903	0,8951	0,8722	0,8019
Medium (20—250)	0,9914	0,9792	0,9786	0,8007	0,8319	0,8407
Large (250—1000)	0,9827	0,9573	0,9745	0,8420	0,7685	0,8287
Very Large (1000 or more)	0,9902	0,9460	0,9850	0,8751	0,7461	0,8707
Total	0,9900	0,9837	0,9822	0,8630	0,8365	0,8509

OURCE OF RAW DATA: RAIS 76; processed data, Tables 1, 2 and 3 of this appendix.
* Assuming that the percentage of undeclared wages all fall into the highest wage bracket.

Table 5

Brazil — 1976 Wage Brackets' Percentage Participation in Payrolls (in Multiples of Highest Minimum Wage)

Sector and Size of Establishments (by No. of Employees)	Percentage Share of Payrolls Corresponding to Each Wage Bracket (in Multiples of Highest MW)				
	0—3	3—10	10—20	20 or More	Sum of Shares*
Industry	36,85	32,75	13,79	12,88	96,27
Micro (0—5)	65,37	20,70	5,72	6,14	97,93
Small (5—20)	56,75	27,08	7,62	5,93	97,38
Medium (20—250)	44,09	30,02	11,88	10,26	96,25
Large (250—1.000)	37,03	30,89	13,64	14,05	95,61
Very large (1000 or more)	25,96	38,06	16,79	15,88	96,69
Cloth., Foot, Softg.	68,88	17,54	5,71	4,18	96,31
Micro	84,32	10,97	1,05	1,77	98,11
Small	80,26	14,51	1,92	0,59	97,28
Medium	69,82	17,95	4,95	3,38	96,10
Large	65,17	16,64	7,20	6,14	95,15
Very Large	60,03	22,00	9,98	6,07	98,08
Motor Vehicles	19,35	46,59	18,16	13,76	97,86
Micro	43,86	28,91	13,94	11,96	98,67
Small	45,50	32,99	12,63	7,21	98,33
Medium	40,70	33,29	13,95	9,09	97,03
Large	28,05	35,96	17,86	16,57	98,44
Very Large	14,20	51,04	18,87	13,85	97,96
Commerce	49,54	27,67	10,38	7,96	95,55
Micro	77,31	16,28	2,70	1,76	98,05
Small	60,98	25,10	6,96	3,36	96,40
Medium	41,15	32,20	12,63	8,67	94,65
Large	34,10	25,89	14,55	18,67	93,21
Very Large	39,42	26,35	12,15	10,64	88,56
Services	32,83	32,70	17,36	12,67	95,56
Micro	68,47	18,92	6,19	4,22	97,80
Small	45,73	30,71	12,79	6,63	95,86
Medium	32,48	33,68	17,05	11,73	94,94
Large	31,45	28,96	18,12	16,68	95,21
Very Large	23,36	36,90	20,86	14,85	95,97

SOURCE OF RAW DATA: RAIS 76; processed data: Tables 1 and 2 of this appendix.
* The sum is less than 100% due to the presence of undeclared wages.

Table 6

Brazil — 1976 — Percentage Share of Employment in Each Wage Bracket (in Relation to Highest Minimum Wage) for the Clothing, Footwear and Softgoods Sector and for the Motor Vehicles Sector)

Size of Establishment (N.º of Emp.)	Clothing, Footwear, Softgoods					Motor Vehicles				
	Wage Brackets in Rel. to Minimum Wage				Absol. N.º of Employ (%)	Wage Brackets in Rel. to Minimum Wage				Absol. N.º of Employ (%)
	0-3	3—10	10-20	20 or More		0-3	3—10	10—20	20 or More	
Micro (0-5)	94,32	3,58	0,10	0,08	13562 (4,74)	83,40	12,48	1,99	0,80	2107 (0,65)
Small (5-20)	92,35	4.72	0.19	0,03	39053 (13,66)	77,29	18,25	2,18	0,61	7785 (2,40)
Medium (20-250)	89,00	6,33	0,56	0,18	133659 (46,77)	70,21	22,68	3,13	1,00	48476 (14,92)
Large (250--1000)	87,89	6,05	0,85	0,35	65750 (23,00)	62,49	28,78	3,93	2,24	60074 (18,49)
Very Large (1000) or More	86,73	9,50	1,43	0,42	33799 (11,83)	35,49	52,91	6,65	2,40	206509 (63,54)
Total	89,15	6,30	0,66	0,23	285823 (100,0)	47,25	42,84	5,67	2,11	324951 (100,0)

SOURCE: RAIS 76.

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Inflation and relative prices: the Brazilian experience, 1970/79 *

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Décio K. Kadota **

1 — Introduction

In this paper we present a few measures of inflation, output growth and relative prices dispersion regarding the Brazilian experience in the 1970/79 period. We neither intend to propound a new theory to explain observed events, nor analytically and strictly establish procedures to distinguish between the causes and effects of inflationary process. Rather, ours is an empirical and descriptive analysis. In the first place, based on behaviour of output growth and of industrial sector prices, we shall try to define the several stages comprised by the cyclic trends of Brazilian economy.

We shall then analyze the behaviour of inflation rates and of relative prices. Under this head we show that there is a positive association between the measures of inflation rates and of their variability, as well as between both these measures and dispersion of relative price changes. Secondly, we will demonstrate the assymetric behaviour of relative price changes. The measure indicates price changes distribution skewed to the right, that is, price increases for most products comprised by the general index remain lower than general index growth in the 1971/79 period. Thirdly, we shall also show a differentiated systematica behaviour evidenced by variation of relative prices in price sub-groups within the general index. These measures particularly refer to the crice

* We thank João Sayad for his helpful suggestions and Luis Antonio Amarante for suggesting and carrying-out statistical tests of profile analyses.

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of farm produce and industrial products, and to price sub-groups within the industrial sector itself.

2 — Short-term and output variations

The implications of monetarist theories on fluctuations and on short-term interdependence between inflation and output growth must be explained, so as to provide a norm for a clearer understanding of the Brazilian experience.

Diagram 1 shows the short-term oscillations of inflation (P) and of output growth (Y), which are consistent with orthodox theories on inflation. On the upper half of Diagram 1 we see on plane (p, y) the course taken by y and by p impelled by the impact of demand, while everything else remains constant. On the lower half, we see the course in time of y and p, which is consistent with the adjustment process shown on plane (y, p). This type of formulation can be found in the works of Friedman (1977) and Almonacid (1971). As shown by Diagram 1, provided a natural rate of growth and from the point of departure of a stable inflation (or stable prices), any nominal income rise based on economic policies will be distributed between a higher rate of output growth and an increase of inflation rates. In Stage I, due to the gap between observed and expected inflation, growth derives from the fall of real wages and from the expansion of the real money supply. In Stage II, output growth falls and the inflationary spiral starts to show lower rates of increase as expected inflation converges towards observed inflation, thus ushering in a reversal of the real wage decrease process, on the one hand, and of the lower real money supply, on the other.

In Stage III, as a result of wage demands, additionally to a rise in prices that counterbalances accured wage losses from Stages I and II and to the need for a decrease of the real stock of money to a level consistent with the new inflationary conditions, output starts to grow at rates lower than the natural growth rate, and inflation starts to recede.

In Stage IV, the recovery process begins and inflation either becomes stable or decreases; the liquidity crisis from Stage III gives way to a milder liquidity condition and real wage escalation is checked by a low level of economic activity. [1] All in all,

1 This short explanation is not intended to advance all monetarist arguments. It can be seen, however, that if initial expansion is based on a steady, say, 5 to 10% growth of money supply, inflation shall of course be higher than that observed at the point of departure in Stage I.

Diagram 1

MONETARIST MODEL: SHORT-TERM INTERDEPENDENCE BETWEEN INFLATION AND OUTPUT GROWTH

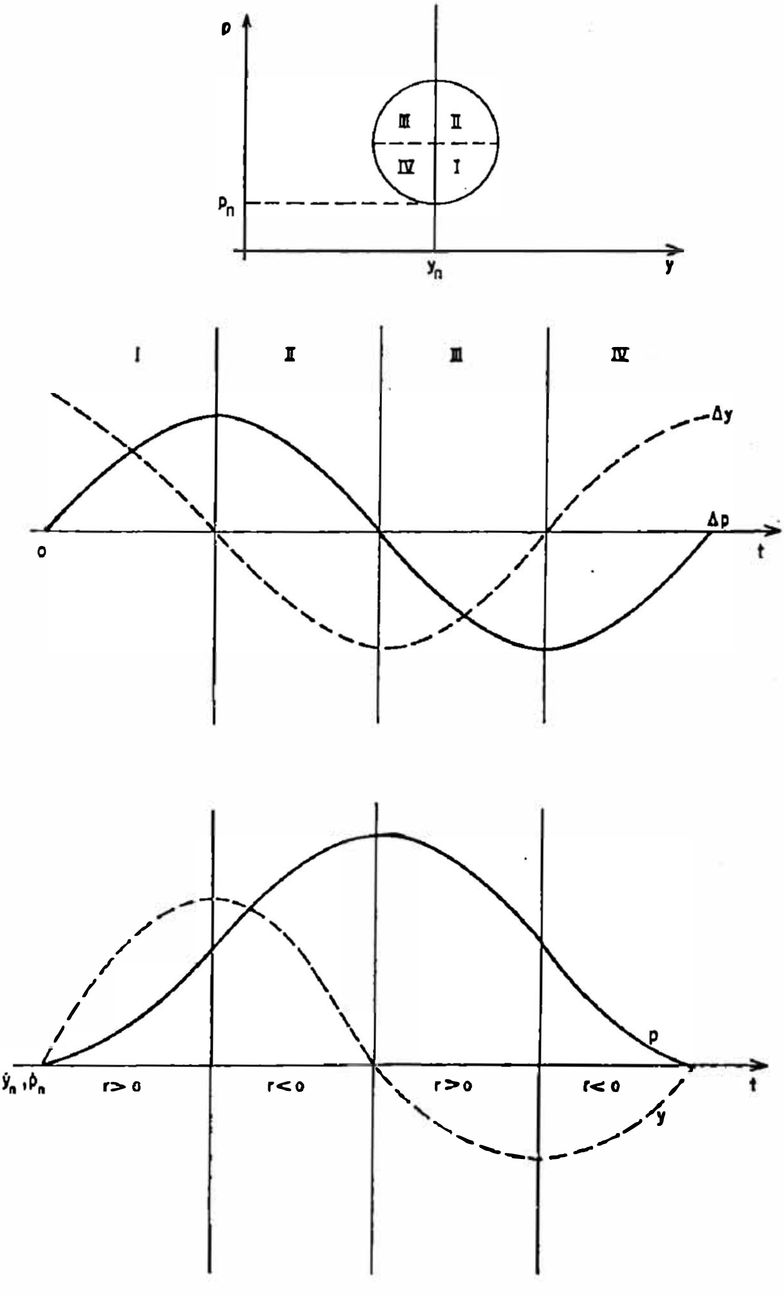


Diagram 1 contains the basic lessons of monetarist logic: an output growth in excess of natural rate that results from the management of aggregate demand (Stages I and II), is but a temporary event that occurs at the expense of future growth (Stages III and IV). The actual trade-off, therefore, is rather between a higher growth rate today and a lower rate in the future, than between inflation and unemployment, as taught by theories associated with the "Phillips Curve", which not always consider the expectations adjustment process. [2] According to this formulation, redistributive effects arising from adjustment processes fall short of significantly altering the mechanisms which lead towards an original equilibrium

Let us now see whether or not this lesson will provide a clear understanding of the Brazilian experience. Once we establish that the country's recent inflationary events predominantly relate to costs, we will be able to demonstrate that the behaviour of y and p in Brazil fails to conform with at least some stages of the monetarist model.

With this in mind we drafted Diagram 2, comprising the industrial sector (manufacture) growth rate and the rate of inflation assessments based on industrial sector prices. [3] The diagram's upper half registers the first differences of u and p, respectively.

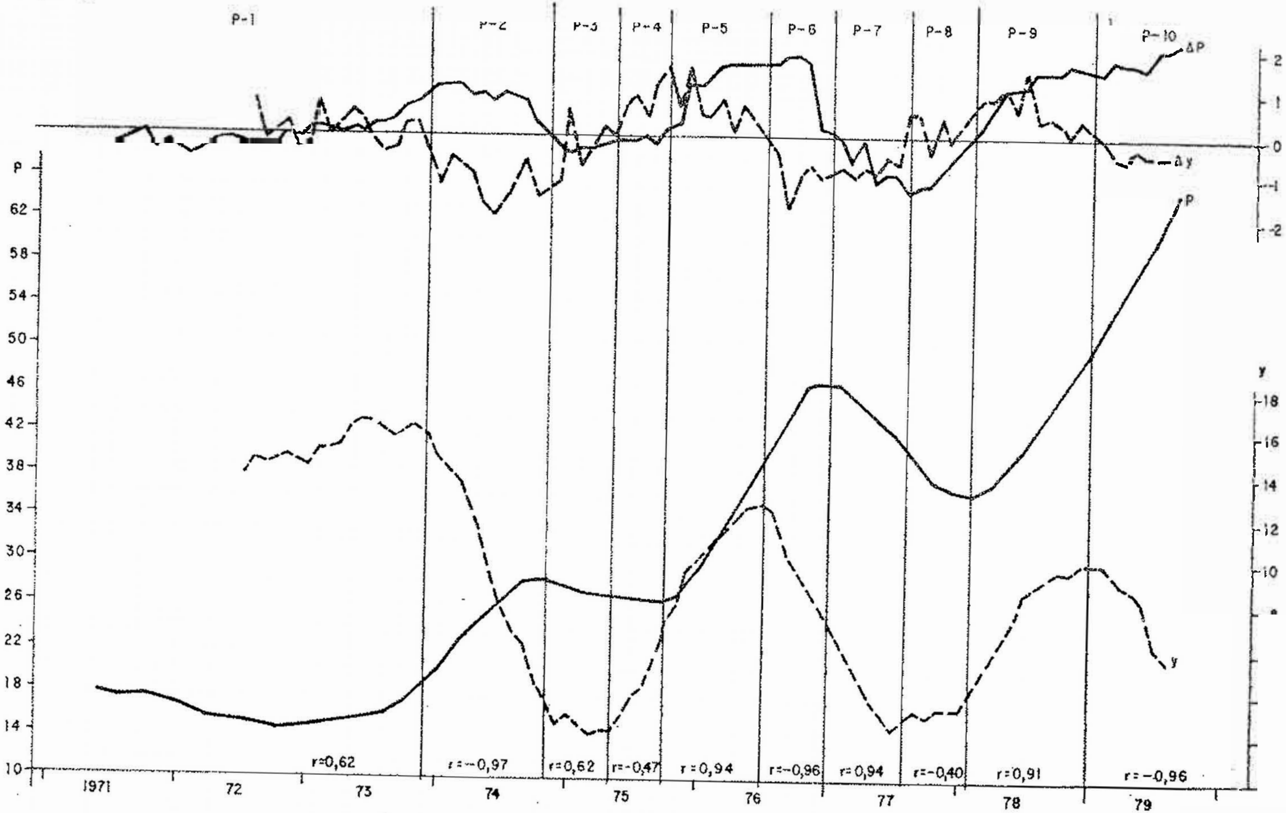
It must be noted that during the sampling period — 1972 to 1979 — three different surges of inflation escalation were recorded — 1974, 1975/76 and 1978/79 — as well as three different cycles of industrial growth slowdown. Is is quite clear that there exists an upward drift of the inflation oscillations and that there is a downward trend of output growth.

By observing such movements on a monthly basis we can verify the extent to which they fit the above model. No one of course expects that Brazilian events could precisely mirror the orthodox pattern described above in all its details, as shown in Diagram I. In the case of output growth rate is difficult to

2 In this type of model, the negative correlation (positive) between inflation rates and unemployment (output variation) occurs in face of a given inflation rate expectation. However, in the adjustment dynamics employed to adjust expectations, the correlation between inflation rates and output depends on the cycle's stage. As shown in the lower half of Diagram I, correlations are positive in Stages I and III, whereas they're negative in Stages II and IV.

3 Inflation rates (p) and output growth rate (y) are defined as 12-month averages, that is, p value in t is given by variation of price outputs rate averages assessed for the t-11 to t and t-12 periods, and similarly for the value of y in t.

Diagram 2



NOTE: P_t and y_t are defined as rate of growth of 12 months moving average at time t in relation to $t-12$.

pinpoint timely the precise moment at which alterations of rate behaviour took place.

However, when classifying stages according to the movements of inflation rates – which are relatively more stable – we see that except for Stage II, the output variation rate shows a profile that is almost consistent with that seen in Figure I. It can also be noted that the values of p and y correlation coefficients (lower half of Diagram 2) for the several adjustment phases also conform with those forecast by the monetarist model.

For comparison purposes, the sampling period was thus classified as follows:

a) Stage I: p and y drifts observed between 1972 and 1973, between September 1975 and June 1976 and, during 1978, mostly follow those of the orthodox model's Stage I;

b) Stage III: p and y oscilations observed between October 1974 and April 1975, and during the first half of 1977, correspond to those shown by Stage III; and,

c) Stage IV: p and y observed behaviour between the second and third quarters of 1975 and the second half of 1977 closely follow those indicated by Stage IV.

This is equivalent to say that in those periods y and p drifts were essentially governed by fluctuations of the variables that mostly affect aggregate demand. It must be seen, for instance, that the fall of industrial growth in early 1975 (Stage III) is followed by stabilization or decrease of inflation rates during 1975 (Stage III). It must also be noted that this latter period was preceded by a severe cut of real liquidity.[4]

Growth escalation (Stage IV) observed in the second half of 1975 and late in 1977 is equally followed by a stage of inflationary escalation (Stage I), that prevailed in both subperiods, early 1976 and early 1978. Moreover, both such subperiods were preceded by expansionist phases of the monetary policy.

Whereas these mini-cycles conform with monetarist norms, it must be remarked that in Stage II, as shows by Diagram 2, p and y drifts do not conform with the forecasts of orthodox system. It is even more important to observe that these divergent drifts occurred in 1974 and late in the 1978/79 period.

On the former date, despite a significant fall of output growth, inflation persisted in an accelerated escalation rhythm. According

[4] See Moura da Silva (1976).

to orthodox precepts, the peak of inflation rates ($\Delta p = 0$) should have occurred concurrently with the maximum speed of y decrease; this, however, was not so due to at least two simultaneous supply shocks that took place during the year: the rise of farm produce real prices and, even more important, the rise of oil prices — not to mention the liberalization of wage policies, that also took place in that year.

A similar event was observed in late 1978/79 with regard to y and p, once agains as the result of impacts caused by farm produced and oil prices. In that late 1979 sub-period came the enforcement of "cruzeiro" maxi-devaluation.

It is also important to note that Stage II, which prevailed over the second half of 1976, even if not quite conforming with orthodox norms, presents a feature difficult to interpret. The duration of the acceleration stage as well as occurrence of inflation rate peak ($\Delta p = 0$) were respectively shorter and less intense than those prevailing in 1974 and in 1978/79. In the latter instance, supply shocks (failure of harvests in 1975/76 and the subsequent increase of farm produce real prices, in addition to the rise of interest rates in 1976 had shorter lasting effects than those of 1974 and of 1978/79. Behaviour of farm produce prices (*vis-à-vis* industrial ones) and of petroleum products prices are shown in Diagrams 3 and 4, respectively.

In short, y and p drifts, as observed in 1974 and in 1978/79, fail to conform with orthodox norms. In these periods, supply shocks magnified the cyclic drifts of y an p; and, even more significantly, the 1974, 1976 and 1978/79 impacts precisely coincide with the three cycles of sharp inflation rate escalation observed along the sampling period. During such periods, Brazilian average inflation rate escalated from 20% p. a. in 1973 to 27% p. a. in 1974; from that level to 40% p. a. in 1976 and, finally, to 52% p. a. in 1979.

Whether a coincidence or not, it is hard not to ascribe the three inflationary surges observed in Brazil between 1972 and 1979 to supply shocks.

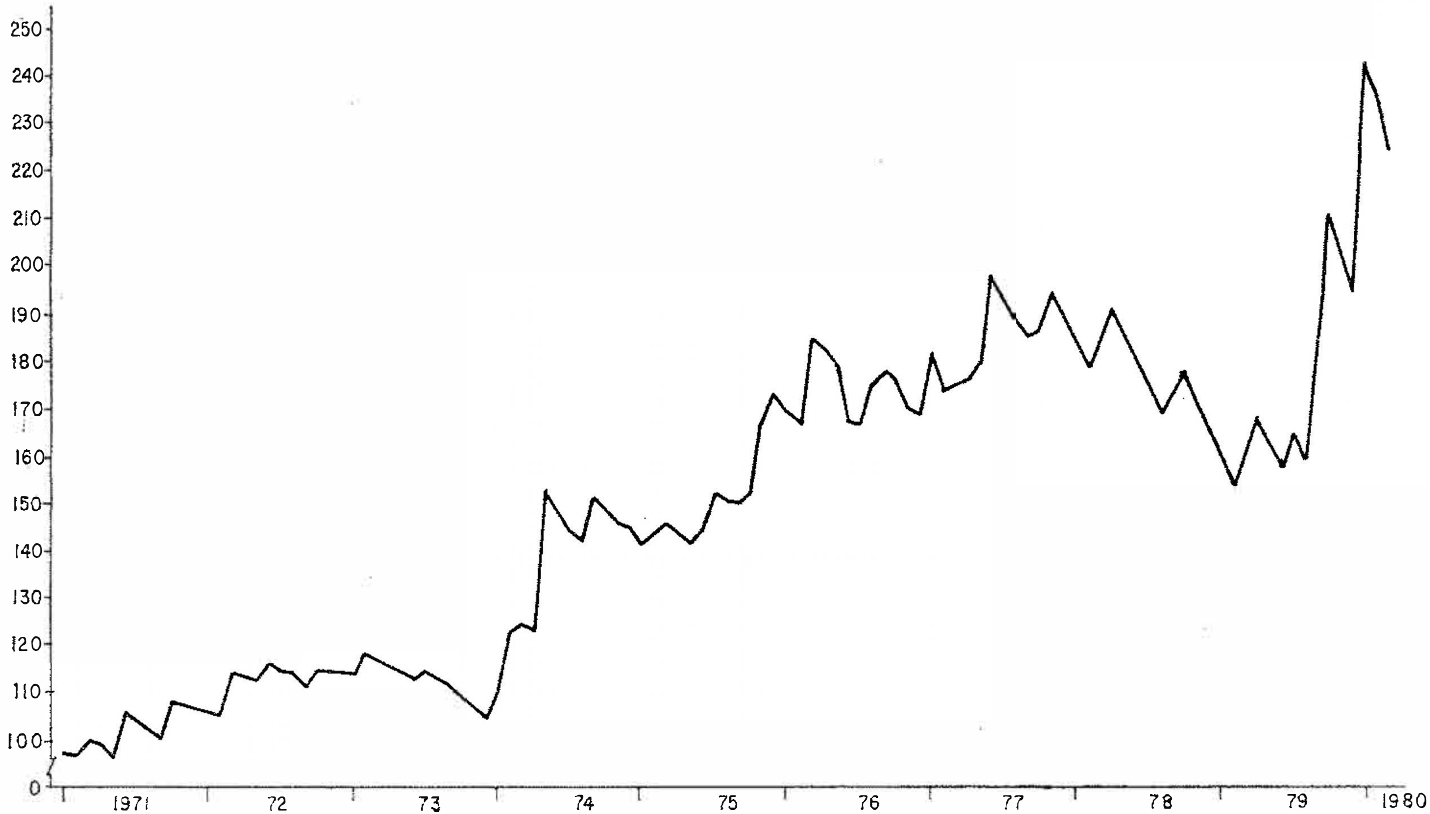
An advocate of the orthodox theory would doubtlessly argue that the impact of relative prices could only be translated into escalating inflation as a consequence of monetary policy liberalization arising from a deep slump in industrial output growth in the period. He might also contend that supply shocks would at best prolong the duration of Stage II, but they would never bring about a rise of inflation rate levels if these had not firstly been sanctioned by monetary policy.

Diagram 3

FARM PRODUCE PRICE INDEX
INDUSTRIAL PRICE INDEX
170
160
150
140
130
120
110
100
1971
72
73
74
75
76
77
78
79
1980

Diagram 4

$$\frac{\text{PETROLEUM PRODUCTS PRICE INDEX}}{\text{INDUSTRIAL PRICE INDEX}}$$



3 — Relative prices and inflation: empirical evidences

When formulating macroeconomic models, it is usual to advance the implicit — or even explicit — hypothesis proposing independence of relative prices dispersion from the inflation rates level.

Lucas (1973) formalized the hypothesis of independence. In his model, Lucas expresses the price of any "*i*" goods in economy as it relates to inflation rates as follows:

$$P_{it} = P_t + Z_{it}$$

where P_{it} and P_t respectively represent the logarithm of prices of "*i*" goods and of the general price level. The independence hypothesis is equivalent to saying that any eventual differences in behaviour shown by P_{it} and P_t occur at random, as stated by Z_{it}, which is normally distributed with zero means and constant variance u^2. It must be further pointed out that *a fortion* there occurs no systematic correlation between P_{it} and P_t, nor is there a differentiated and systematic behaviour involved among diverse price sub-groups that constitute the general index.

Complementarily, and considering P_t, too, as a normal random variable with mean equal to $\bar{P}_t$ and constant variance given by v^2 distributed independently of Z_{it}, Lucas also admits independence between the latter and relative prices dispersion. This hypothesis has been subjected to countless tests.

The hypothesis advancing the independence between variability and inflation rate levels has been almost universally rejected. In a cross-section analysis carried out internationally Logue and Willet (1976), using data collected in 45 countries in the period comprised between 1949 and 1970, found a positive association between the inflation reate standard deviation and the average inflation rate. Similar result was found by Jaffee and Kleiman (1977), which involved 17 countries of the OECD in the period comprised between 1951 and 1968, and 16 Latin American countries in the period comprised between 1950 and 1969. Fisher (1981) when analyzing time series with respect to the United States also found a relation between the standard deviation of inflation rates and the level of those rates. Standard deviations of yearly and quarterly rates were assessed with basis on non-overlapping, 5-year and 12 quarterly periods, respectively, and the positive correlation with average inflation rates for those periods was found.

With respect to the variability of relative prices, there is some evidence that it is closely linked with inflation rate levels, and with their variability. Glejser (1965) analyzed the main groups of goods comprised by Consumer Price Indices in 15 countries of the OECD, and found a positive correlation between relative prices dispersion and the average inflation rate between countries.

Jaffee and Kleiman argue that the above analysis presents more drawbacks. On the one hand, it does not allow for distinctions between the effects of year-by-year relative prices changes and the cumulative effects of these changes and on the other hand, relate diverse national experiences. So that 13 countries were individually analyzed as to the correlation between the standard deviation of the price changes of the main sub-group of goods, and inflation rates. [5] Results for most countries show that the coefficient of variation decreases proportionally to inflation rates. However, as the constant of the hyperbolic function adjusted did not significantly differ from zero, the authors conclude that the absolute value of relative prices dispersion is invariable in relation to the rate of inflation. [6]

On the other hand, a similar analysis involving approximately 1500 goods comprised by the Wholesale Price Indexes in the United States was undertaken by Vining and Elwertowski (1976) who found positive correlations between the standard deviation of the price changes and the rate of inflation and its variability in the 1948/74 period.

Cukierman and Wachtel (1979) and Parks (1978), show that the evidence found by Vining and Elwertowski can be explained with basis on a positive correlation of prices variability and the unanticipated inflation. While the first employed a variance factor out of Lucas' Model of Rational Expectation, the second used a traditional expectation model, that is, a model where the anticipated inflation rate is a function of rates observed in the past.

Positive correlations between dispersion of relative prices and current inflation rates, whether or nor anticipated, were likewise found by Fischer (1981) with respect to the United States.

The above mentioned evidence shows the existence of positive correlation between not only variability and inflation rate levels, but also between relative prices dispersion and the level and variability of inflation rates. However, no records exist with regard to controversies on the existence of differentiated behaviour of relative prices of product sub-groups.

[5] The periods under consideration differ among countries and disaggregation of the price index involving about 10 groups of goods were used.

[6] The adjusted function was of the following type: $SD(P_t)/P = a + b\ P^{-1}$. As $a = 0$, we have: $SD\ (P_t)/P = b\ P^{-1}$, or $SD\ (P_t) = b$.

4 — Variability measures or relative prices and inflation

The Wholesale Price Index as computed by the Getulio Vargas Foundation for the period comprised between 1970/80, was used in this study. The maximum disaggregation is given by the concept of Aggregate Supply, involving 50 groups of goods. In the case of analysis by goods' types involving the original 50 groups, 13 were not considered, basically because their prices were determined by foreign markets or controlled by the government. Of the remaining 37 groups, 7 were classed as agricultural groups, and 30 as industrial groups. Industrial products, on their turn, were subdivided into three groups comprised by 10 items each, classified by order of the increasing degree of sector concentration that produces these goods. Based on the above indices we defined both for general case and for each goods sub-groups the standard deviation of inflation rates and of relative prices that are usually employed in the papers mentioned above.

In view of the fact that inflation rate variability will be measured by time periods, whereas relative prices dispersion will be assessed by types of goods, at a given point in time, so as to match the time gaps involved in such measures, price index defined on a montly basis were adopted.

Variability of inflation rates was measured by use of the moving standard deviation on monthly rates change of the general index, with a 10 month amplitude, and the corresponding result was associated with the last month of that interval, as shows below:

$$SD(P)_t = \sqrt{\frac{\sum_{i=0}^{11} (P^m_{t-1} - \overline{P}^m_t)^2}{12}}$$

where P^m_t is the monthly rate of change of the general index and $\overline{P}^m_t$ is the arithmetic average of monthly rates over the t — 11 to t period.[7]

The dispersion of relative price changes was measured by use of the weighted standard deviation of the rate of change in 12 months, as to the index of each group of goods:

$$SD(P_i)_t = \sqrt{\sum_{i=1}^{n} w_i (P^a_{it} - P^a_t)^2}$$

[7] This measure was used by Klein (1976).

where P^a_{it} and P^a_t represent percentage change in the prices of goods groups "i", and in the general price index at t, with respect to the same month of the preceding year, and w_i the participation of value added of goods group "i" on the overall total.

The coefficient of skewness of individual prices changes P^a_{it} was measured as well:

$$S(P_i) = \frac{\frac{1}{n}\sum_{i=1}^{n}(P^a_{it} - \bar{P}^a_t)^3}{\left\{\frac{1}{n}\sum_{i=1}^{n}(P^a_{it} - \bar{P}^a_t)^2\right\}^{3/2}}$$

5 — Results

In this section we have tested the above mentioned hypotheses in the following sequence:

a) *Inflation and Inflation Rate Variability*

As shown by the studies already mentioned, there seems to be no doubt that a positive relation exists between inflation rates (P) and their variability ($SD(P)$. The positive correlation obtained by us was 0.73. [8] It must however be noted that the coefficient of variation is relatively stable, which means that inflation rates variability increases in equal proportion to inflation rates as shown by Diagram 5.

b) *Inflation and Relative Prices*

The most significant evidence, however, is found in the positive correlation between relative prices dispersion ($SD(P_i)$) and inflation rates (P), on the one hand, and between $SD(P_i)$ and $SD(P)$, on the other. The coefficients of correlation between $S(P_i)$ and P and between $SD(P_i)$ and $SD(P)$ were 0.77 and 0.47, respectively, both significant at the 1% level.

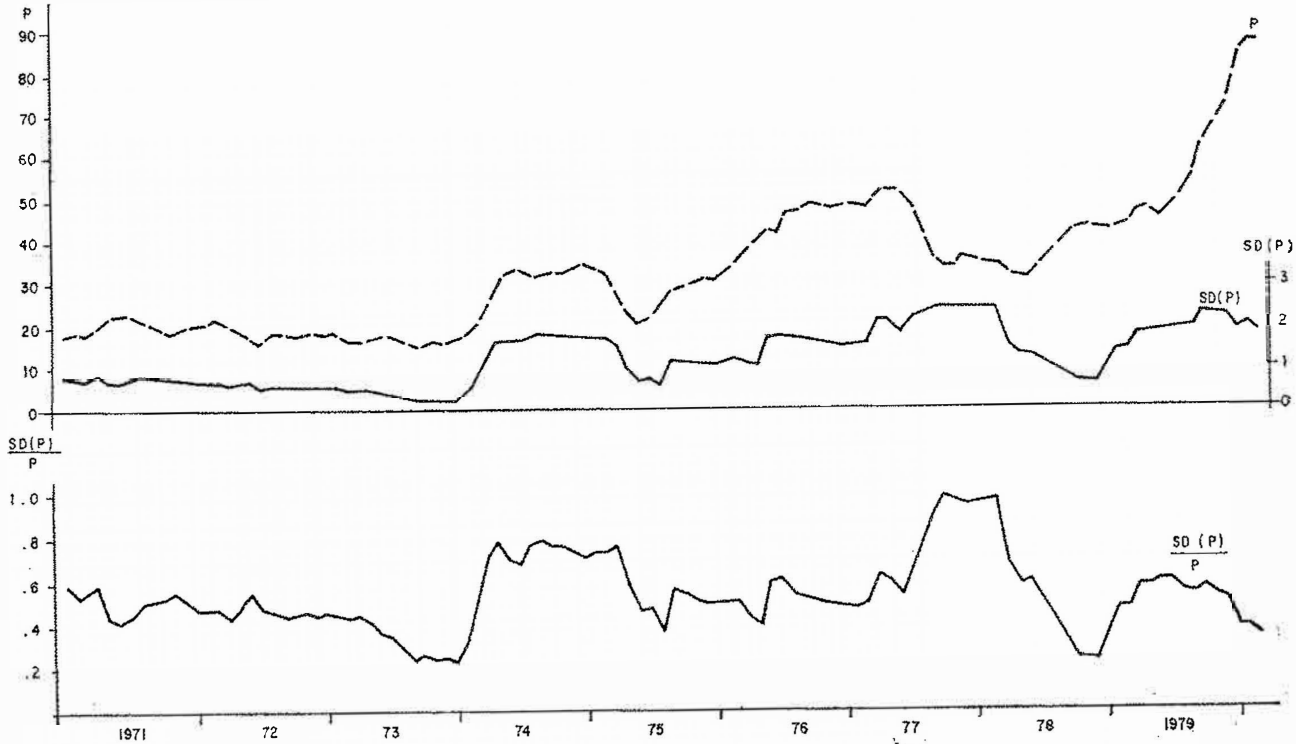
These results are similar to those found in the above mentioned papers with the exception of those obtained by Jaffee and Kleiman. However, carrying out the Jaffee and Kleiman test with respect to Brazil, we obtained the following results, using even more disaggregated data than they did:

$$\frac{SD(P_i)}{P} = \underset{(58.6)}{0{,}429} + \underset{(23.9)}{4.455}\; P^{-1} \qquad R^2 = 0.84$$

8 Significant at the level of 1%.

Diagram 5

VARIABILITY AND LEVEL OF INFLATION



NOTE: Coefficient of variation measur d out by monthly rate of inflation.

As evidenced by "t" ratios between parentheses, there exists an inverse correlation between the coefficient of variation defined by $SD(P_i)/P$ and inflation rate P, although notwithstanding this, the constant cannot be considered equal to zero, that is, that the $SD(P_i)$ value would not be invariant with respect to P. [9]

c) *Relative Price Distribution Profile*

Based on the coefficient of skewness as defined above, we may undertake a test of the hypothesis that the rate of change of individual prices P_i behave to a normal distribution.

The values of coefficient $S(P_i)$, as well as the confidence interval with 98% of probability that P_i is normal, that is, $S(P_i) = 0$ is found in Diagram 6. [10]

Pursuant to Vining and Elwertowski the hipothesis of symmetrical distribution is rejected for practically the entire period, and particularly with regard to subperiods of inflation rate increase.

Distribution regarding each moment in time presented a positive asymmetry, that is, the type of distribution where mode < median < average. According to Diagram 7, in that same period the arithmetic average of the rate of change of the prices P_i is approximately equal to the rate of change of the general index (weighted average).

Thus, over 50% of the rate of change of individual prices comprised by the general index is lower than inflation rate.

The listed results, as seen above, have really no meaning when considered as to the interim character, or not, of interdependence of inflation and relative prices. Authors such as Cukierman, Wachtel and Parks conciliate such results with the Lucas' hypothesis by alternative schemes of economic agents' expectation. For them, these correlations would therefore be maintained only throughout the adjustment process. In this light, the relative prices dispersion test that involves diverse price sub-groups becomes an important alternate test for Lucas' hipothesis.

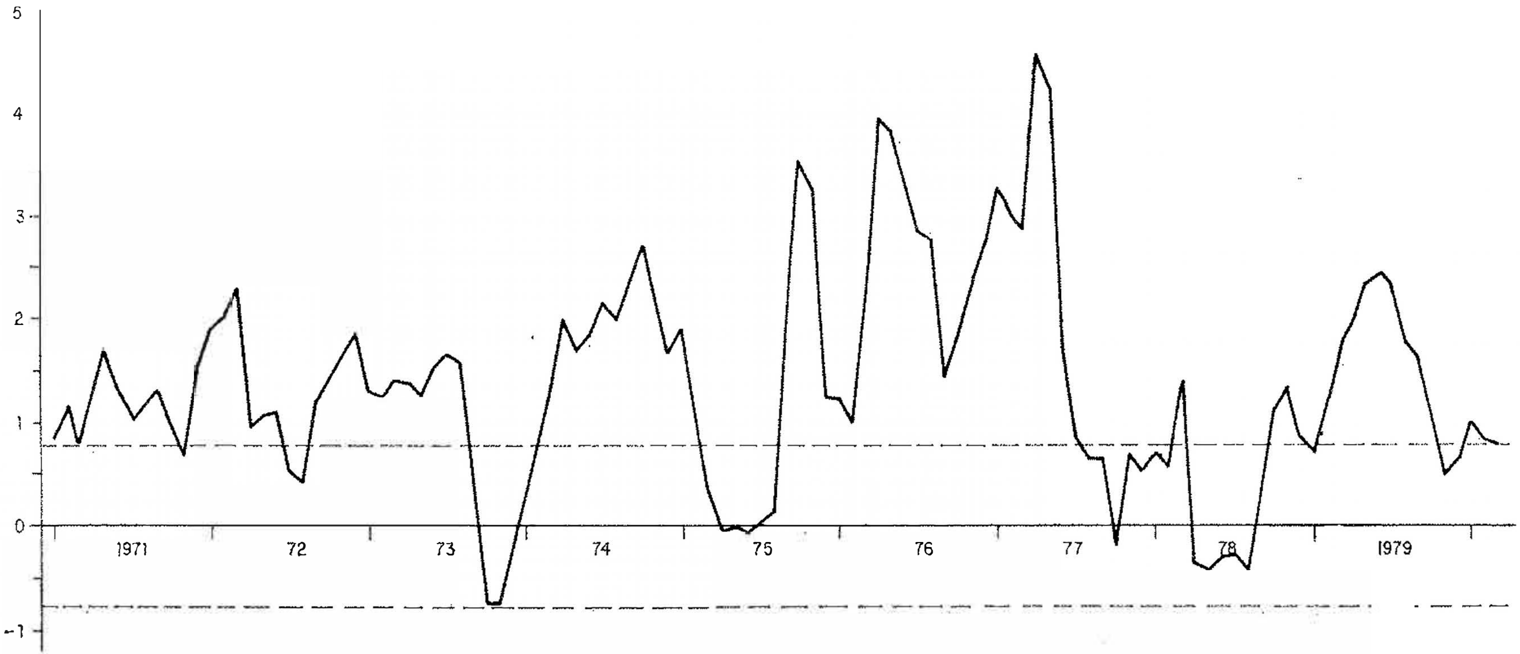
9 The result found by the authors with respect to Brazil using yearly data from the 1969/70 period, and 10 groups of goods, was as follows:

$$\frac{SD(P_i)}{P} = \underset{(0.38)}{0.09} + \underset{(3\ 07)}{4.60}\ P^{-1} \quad R^2 = 0.57$$

10 Further detail regarding the building up of confidence intervals can be found in Snedecor and Cochran (1967).

Diagram 6

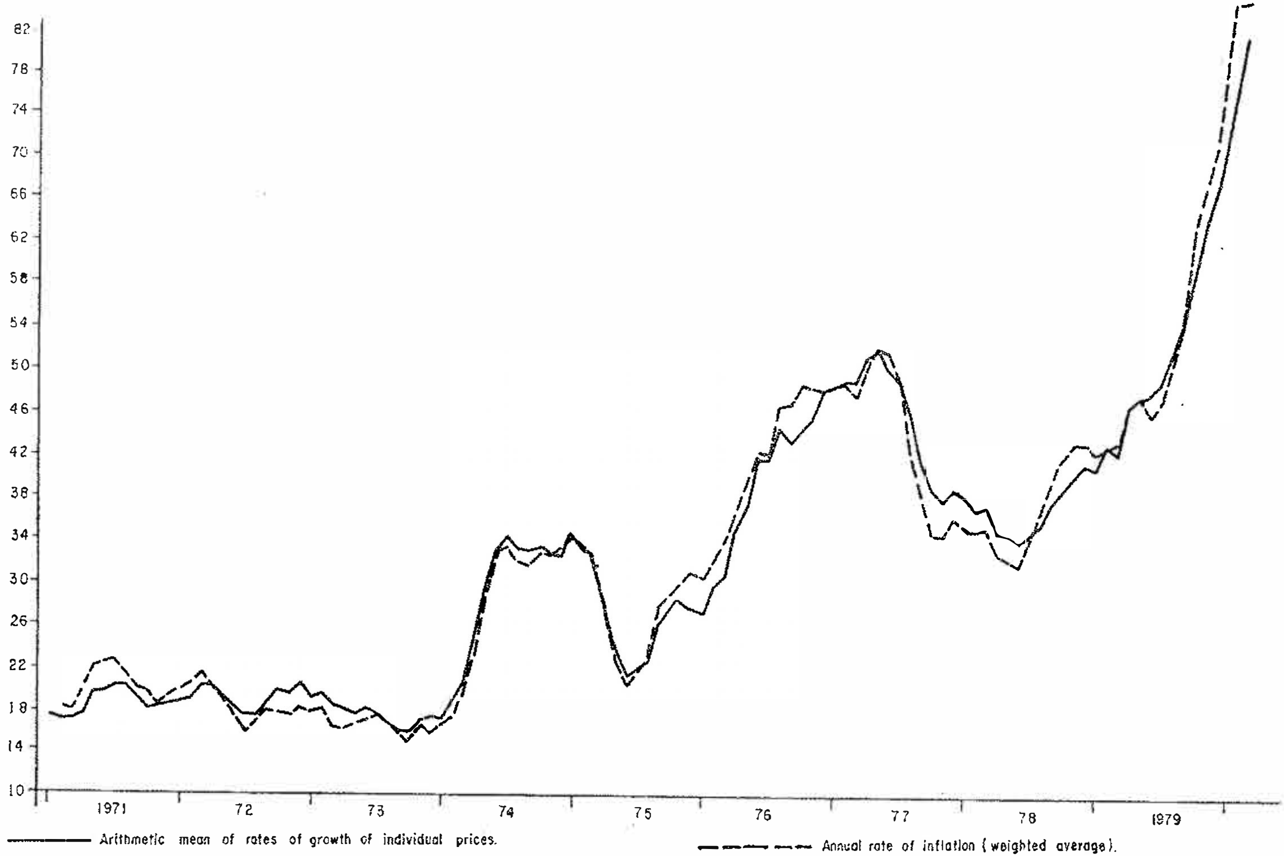
COEFFICIENT OF SKEWNESS



NOTE: The dotted lines show the confidence intervals at the 98 % level.

Diagram 7

RATE OF INFLATION AND ARITHMETIC MEAN OF RATES OF GROWTH OF INDIVIDUAL PRICES



d) *Relative Price Dispersion in Price Sub-Groups*

Measure of variability of the relative prices of agricultural, industrial and industrial sub-groups of products on an increasing sequence by degree of production concentration can be seen on Diagrams 8 and 9, and in Table 1, where the average values are show with reference to adjustment phases of industrial products and prices, as defined by the preceding section.

There apparently are behavioural diversities regarding the agricultural and industrial relative prices, since in the first case dispersion is higher than in the second.

The same phenomenon can be observed among sub-groups of industrial prices, and in this case, dispersion is lower in inverse proportion to the degree of production concentration. Presuming that standard deviation $SD(P_t)_{jt}$ as measured for each price sub-groups "j", at moment "t", has a normal multivaried distribution with mean vectors $s_j = s_{jp})$, where index "p" represents the adjustment phase of industrial products, we may undertake profile tests between the price sub-groups. [11]

In view of the restraint imposed by test requirements as to equality of the number of observations made in each "p" phase, a sampling of values $SD(P_t)_{jt}$ was made, with 6 observation samples for each price sub-group "j", in each "p" stage. [12]

Basically we perform two tests comparing time series of standard deviations of relative price changes calculated for each component – according to classification above mentioned – of the general price index: tests of coincidence and parallelism.

The null hypothesis of these two tests, considering the series i and j, are:

$$H_0\colon \Delta_{ip} = \Delta_{ip}, \text{ for all } p = 1, 2, \ldots,$$

$$H_0\colon \Delta_{ip} - \Delta_{jp} = K, \text{ for all } p = 1, 2, \ldots,$$

where Δ_{ip} and Δ_{jp} are respectively the mean standard deviation or relative price changes of sub-group prices i and j, at phase p, and K is a constant real value.

The results shown by Table 2 indicate that vectors s_j are neither equal, nor parallel, whether in comparing agricultural and industrial prices, or in the comparison with sub-groups of concentrated and non-concentrated industrial products.

11 Further details on multivaried tests can be found in Morrison (1967).

12 In phase 4, since we have only 5 observations, the 1st. value of the immediately subsequent phase was considered.

Diagram 8

VARIABILITY OF RELATIVE PRICES: $S(P_i)$

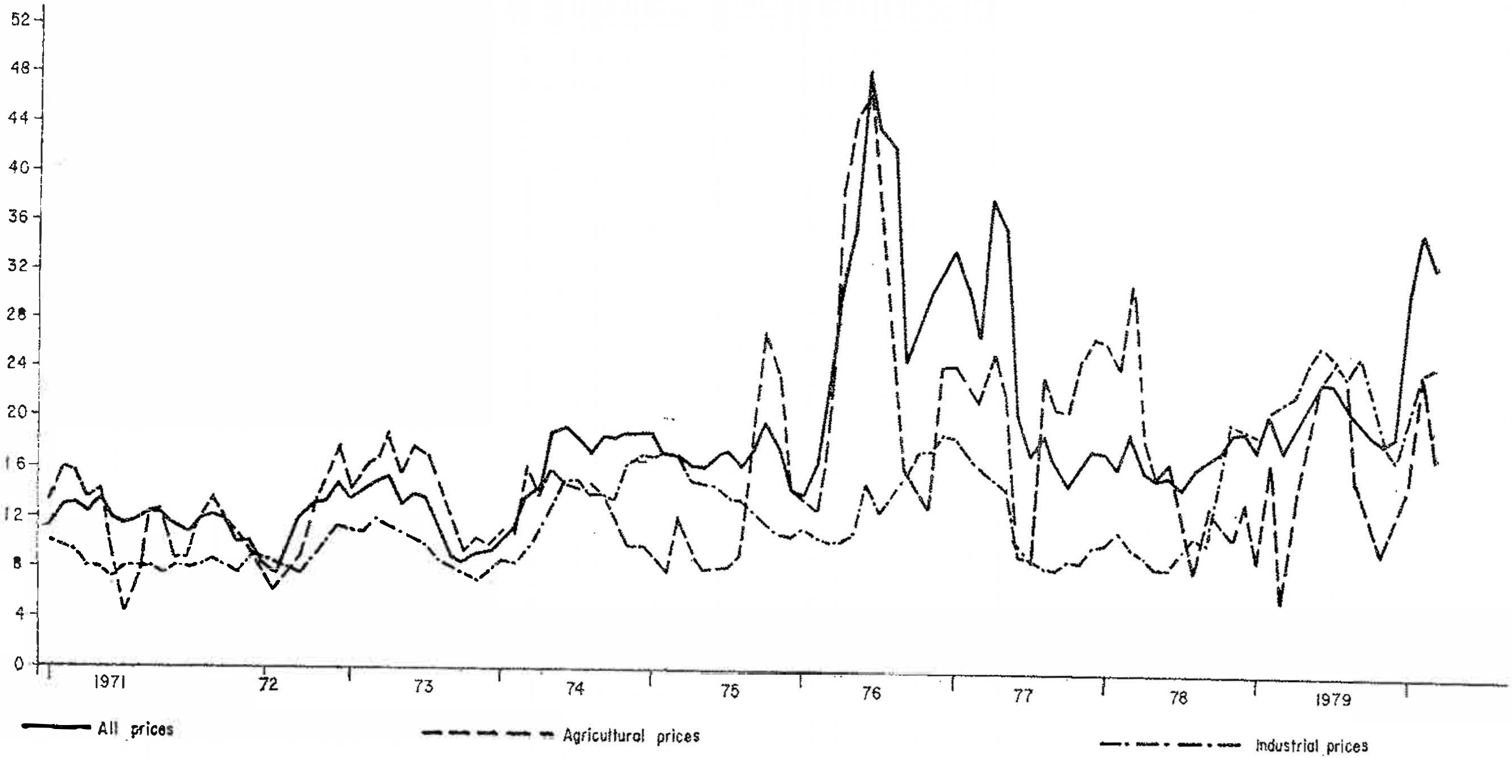


Diagram 9

VARIABILITY OF RELATIVE PRICES: $S(P_i)$ OF SUB-GROUPS OF INDUSTRIAL PRICES

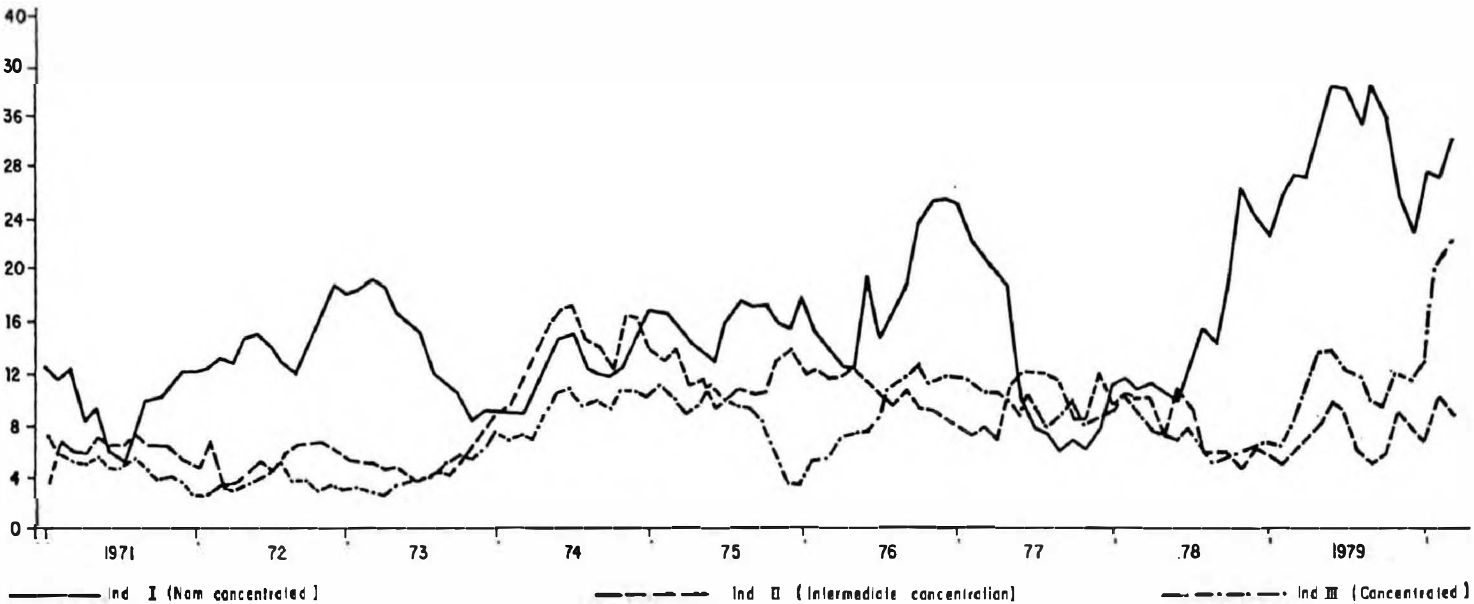


Table 1

Main Results by Phase of Adjustments in Industrial Prices and Product: Average Values

		Phase 1 (6/72-11/73)	Phase 2 (12/73-10/74)	Phase 3 (11/74-4/75)	Phase 4 (5/75-9/75)	Phase 5 (10/75-6/76)	Phase 6 (7/76-12/76)	Phase 7 (1/77-7/77)	Phase 8 (8/77-1/78)	Phase 9 (2/78-12/78)	Phase 10 (1/79-2/80)
	General	18,0	27,7	27,3	26,8	40,3	55,6	51,5	37,9	37,4	52,8
	Agric.	17,5	25,5	24,7	32,1	38,7	58,9	54,0	42,8	51,0	60,1
	Ind.	15,6	25,0	27,2	26,5	32,7	44,7	43,8	37,5	42,6	50,7
P	Ind. I	20,4	21,2	15,6	19,4	35,9	49,1	42,6	34,0	48,0	69,7
	Ind. II	14,5	35,1	39,8	30,9	26,4	34,2	41,8	40,1	37,7	45,8
	Ind. III	12,2	20,9	28,2	29,3	30,5	37,6	42,2	41,6	36,9	41,3
	General	12,2	16,6	17,5	17,9	27,0	31,8	27,0	16,8	17,3	23,4
	Agric.	13,5	13,6	9,8	14,6	28,1	20,5	19,7	24,4	15,0	16,9
	Ind.	9,7	12,8	16,7	13,7	11,9	17,4	13,4	10,1	13,3	23,1
$SD(P^a)$	Ind. I	14,4	11,5	15,0	15,8	14,7	21,7	14,7	7,7	15,4	28,1
	Ind. II	5,2	13,3	12,8	9,7	11,5	8,7	9,2	9,4	6,8	6,7
	Ind. III	3,8	8,6	9,7	9,3	5,6	11,1	0,3	8,5	6,2	11,7
	General	0,60	1,35	1,60	0,88	1,28	1,66	1,98	2,44	1,21	1,94
	Agric.	1,06	2,03	2,33	1,20	1,79	2,71	3,67	4,18	2,49	3,74
	Ind.	0,36	0,80	0,79	0,56	0,94	2,05	1,52	0,82	0,65	0,92
$SD(P)$	Ind. I	0,78	1,01	1,22	1,60	1,87	2,42	2,05	1,22	1,20	1,41
	Ind. II	0,43	1,12	1,24	1,03	0,70	0,73	0,85	0,96	0,88	1,28
	Ind. III	0,47	0,79	0,83	1,03	1,14	1,19	1,01	0,85	0,95	1,25

NOTE: P = rate of change of the indicated price index (annual rate).
$SD(P_i)$ = standard deviation of the relative price changes of the indicated sub-groups of price index (annual rates).
$SD(P)$ = standard deviation of the rate of inflation, as measured by the indicated price index (measured monthly).

Table 1

Main Results by Phase of Adjustments in Industrial Prices and Product: Average Values

		Phase 1 (6/72-11/73)	Phase 2 (12/73-10/74)	Phase 3 (11/74-4/75)	Phase 4 (5/75-9/75)	Phase 5 (10/75-6/76)	Phase 6 (7/76-12/76)	Phase 7 (1/77-7/77)	Phase 8 (8/77-1/78)	Phase 9 (2/78-12/78)	Phase 10 (1/79-2/80)
	General	18,0	27,7	27,3	26,8	40,3	55,6	51,5	37,9	37,4	52,8
	Agric.	17,5	25,5	24,7	32,1	38,7	58,9	54,0	42,6	51,0	60,1
	Ind.	15,6	25,0	27,2	26,5	32,7	44,7	43,8	37,5	42,6	56,7
P	Ind. I	20,4	21,2	15,6	19,4	35,9	49,1	42,6	34,0	48,0	69,7
	Ind. II	14,5	35,1	39,8	30,9	26,4	34,2	41,8	40,1	37,7	45,8
	Ind. III	12,2	20,9	28,2	29,3	30,5	37,6	42,2	41,6	36,9	41,3
	General	12,2	16,6	17,5	17,8	27,0	31,8	27,0	16,8	17,3	23,4
	Agric.	13,5	13,6	9,8	14,6	28,1	20,5	10,7	24,4	15,0	16,9
	Ind.	9,7	12,8	16,7	13,7	11,9	17,4	13,4	10,1	13,3	23,1
$SD(P^a)$	Ind. I	14,4	11,5	15,0	15,8	14,7	21,7	14,7	7,7	15,4	28,1
	Ind. II	5,2	13,3	12,8	9,7	11,5	8,7	9,2	9,4	6,8	6,7
	Ind. III	3,8	8,6	9,7	9,3	5,6	11,1	9,3	8,5	6,2	11,7
	General	0,50	1,35	1,60	0,88	1,28	1,66	1,96	2,44	1,21	1,94
	Agric.	1,06	2,03	2,33	1,20	1,79	2,71	3,67	4,18	2,49	3,74
	Ind.	0,36	0,80	0,79	0,66	0,94	2,05	1,52	0,82	0,65	0,92
$SD(P)$	Ind. I	0,78	1,01	1,22	1,60	1,87	2,42	2,05	1,22	1,20	1,41
	Ind. II	0,43	1,12	1,24	1,03	0,70	0,73	0,85	0,96	0,88	1,28
	Ind. III	0,47	0,79	0,83	1,03	1,14	1,19	1,01	0,85	0,95	1,26

NOTE: P = rate of change of the indicated price index (annual rate).
$SD(P_i)$ = standard deviation of the relative price changes of the indicated sub-groups of price index (annual rates).
$SD(P)$ = standard deviation of the rate of inflation, as measured by the indicated price index (measured monthly).

Additionaly, the hipothesis of a stationary series was tested as well; results show that standard deviations of relative prices in each of the price sub-groups are not constant from one adjustment stage to the other. The null hypothesis tested was: H_0: $\Delta_{p} = K$, for all $p = 1, 2, \ldots,$.

Therefore, there are now two additional evidences against Lucas' hypothesis, namely:

i) the existence of systematic behaviour diversity in the case of relative prices among price subgroups;

ii) the non-constancy of standard deviation of relative prices over time.

e) *Inflation, Relative Prices and Industrial Products*

Based on the measurement computed for industria prices as a whole, we shall briefly discuss the evidence found by Blejer and Leiderman (1980), regarding the negative impact of price dispersion increases on the rate of change of industrial product. By interrelating variability measures of inflation rates and relative prices, of inflation rates and product growth, we were able to obtain the following correlation matrix:

Table 2

Profile Analysis

Tests	F
•Agric and •Ind Parallel	$F_{9:2} = 544{,}79$
•Agric = •Ind	$F_{10:1} = 283{,}91$
•Ind I and •Ind III Parallel	$F_{9:2} = 431{,}11$
•Ind I = •Ind III	$F_{10:1} = 741{,}70$
Agric. Series is Stationary	$F_{9:2} = 217{,}65$
Ind Seiies is Stationary	$F_{9:2} = 344{,}59$
Ind I Series is Stationary	$F_{9:2} = 242{,}86$
Ind III Series is Stationary	$F_{9,2} = 208{,}40$

NOTE: In allcases the null hipoteses is rejected. The critical values of the F — distribution are: $F_{9:2} = 19.4$ and $F_{10:1} = 242.0$ at the 5% level.

	$SD(P)$	$SD(P_i)$	P	y
$SD(P)$	1	0.36*	0.55*	—0.21**
$SD(P_i)$		1	0.76*	—0.29*
P			1	—0.44*
y				1

NOTE: * — significant at 1% level.
** — significant at 5% level.

Even if unable to make any statement as to direction of causation, our results also point to the negative association of product growth rares with respect to relative price dispersion, as well as with the standard deviation of the inflation rate.

6 — Conclusions

The results of our analysis of the Brazilian case are consistent with the majority of existing evidence advanced by other analysis on other inflationary conditions. They indicate that there exists positive correlation between inflation rates, inflation rate variability, and relative prices dispersion. Particularly with respect to industrial prices, both latter factors show negative correlations with the rate of growth of industrial products.

We have also been able to demonstrate that the distribution shape of individual prices changes also presents a positive asymmetry. A new angle of the problem involves obtaining proof that price sub-groups within economy present differentiated and systematic behaviour. On the one hand, our results indicate that dispersion of agricultural prices is higher than that of industrial prices and, on the other, dispersion of prices of concentrated industrial products is less than that of the dispersion of prices of non-concentrated products. The main inferences arising from such results, with regard to macroeconomic anaylisis, are:

a) Stabilization policies based solely on the control of aggregated demand results in permanent losses for *some* competitive sectors, as to non-competitive ones; thus, stabilization policies based on one product models are inadequate in that they do not consider the implicit redistributive effects on the systematic and differentiated behaviour of relative prices under an inflationary process.

b) *Full* and *automatic* monetary re-adjustment defined with basis on a given inflation measure inflicts serious losses on a significant number of products. As recently was the case in Brazil, where *modal inflation* is positioned to the left of *average inflation* (the explicit hypothesis usually refers to an average equal to mode in the frequency distribution of prices changes), the correct policy would probably be to define a monetary readjustment policy in which automatic readjustments of contractual values are systematically enforced at lower levels than those indicated by the increase of general prices indexes. It would likewise be misleading to suppose that the level of nominal interest rates should perfectly mirror the inflation rate level in addition to the so-called real interest rates. That is, in Fisher's equation the inflation rate coefficient associated with the expected inflation rate should be lower than unity.

c) There are doubts as to validity of the orthodox view that the fundamental question is the stabilization of the inflation rate whatever its level. As long as a positive association exists between the level of inflation rates and its unstability it will be difficult to avoid the adverse effects of variable relative prices over resource allocation. The monetarist view could only be tenable if that association were an interim phenomenon, only mirroring the incidental effects of inflation escalation.

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Subsidy policies and the export of manufactured goods in Brazil

Alberto Roque Musalem *

1 — Introduction

Since the 1973 oil crisis, Brazil has experienced a chronic deficit in foreign trade. The country will need an economic policy of expenditure shifting, through a rise in the relative prices of tradeable goods and/or containment of imports, to achieve equilibrium in its balance of trade. An overall strategy for the sector will have to be based on an in-depth understanding of the response of manufactured exports to both relative price incentives and the cooling of aggregate demand.

This study is divided into three parts. The first presents a time series of the rates of all subsidies to manufactured exports. This was only possible thanks to the recent contributions of Cardoso (1980). Our own contribution here is to have corrected some of the data from conflicting sources and, especially, to have included all the financial subsidies and drawbacks. To make simpler the calculations of opportunity costs of financial resources obtained, we have assumed rational expectations by economic

Editor's note: Translation not revised by the author.

* Visiting professor in the Masters Program in Economics at the Federal University of Bahia. The author is grateful to his research assistant, the graduate student José Carrera Fernandez, for his dedication, efforts and efficiency. He also expresses his thanks to the graduate student Romeu Salaro, for his rapid collection and mailing of relevant statistical information from Brasília. J. L. Carvalho, of the EPGE/FGV, and M. Barata, of the FIPE, were very helpful collegues in answering requests for some of the statistical data series. All these joint efforts made it possible to carry out and to conclude this study as possible. Victor J. Elias, at the National University of Tucumán, Argentina, made valuable comments. The Rockeffeler Foundation gave financial support. Originally published in *Revista Brasileira de Economia, 35* (1), January/March 1981.

agents in relation to expected interest rates and expected rates of currency devaluations.

In the second part, the time series of the three basic categories of subsidies to manufactured exports — tax exemptions, fiscal incentives and financial subsidies — are used to look into the relative importance ofeach type of subsidy in explaining the behavior of manufactured exports. The model we have here was recently proposed by Cardoso and Dornbusch (1980*a*).[1]

The third part presents our conclusions, and the appendix contains the data which has been most relevant to our study.

2 — Time series of subsidies to the export of manufactured goods

Table 1 gives a summary of the estimated subsidy rates for each component of policies to promote the export of manufactured goods. These policies gained steam following 1964, with a deliberate reorientation of the previous approach to import substitution policies. This particular export-promotion policy was introduced gradually, as the table itself shows, and displayed an upwards tendency in the total rate of subsidies.[2]

The first policy instrument to be used was tax exemption; initially, in 1964, with exemption from payment of the tax on industrialized products (IPI, a federal tax), and later, in 1967, with the additional exemption from state sales taxes (ICM). The exemption policies were rounded out in 1969, with the introduction of drawbacks on import taxes for inputs from abroad used directly in the production of export goods.

The granting of fiscal incentives as a means of stimulating manufactured exports goes back to 1969, with the concession of IPI bonus credits (tax breaks on the unexported portion of an exporting company's production), followed the next year by the introduction of the ICM bonus. Fiscal incentives were rounded out in 1971, with the addition of income tax cuts.

Financial subsidies can be broken down into two major components: export credits at below-market interest rates, introduced in 1968, and drawbacks on compulsory deposits for imports of inputs used directly by the exporter, introduced at the same time as this deposit requirement in 1975.

1 Some of the other important studies in this area: Barata (1979), Carvalho and Haddad (1978), Coes (1979) and Tyler (1976).

2 For a detailed description of the legislation, see Cardoso (1980).

Table 1

Subsidies Granted Under Policies to Promote the Export of Manufactured Goods (as Percentage of the FOB Value)

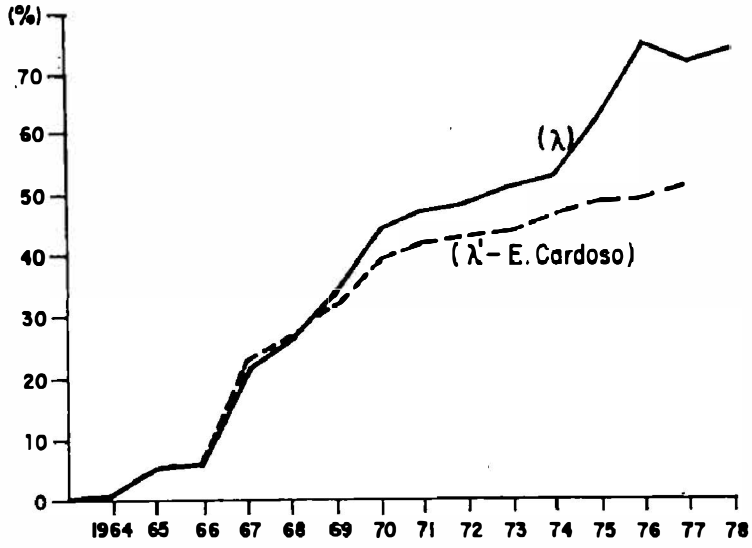
Years	Tax Exemptions			Rate of Tax Exemp-tions	Fiscal Incentives			Rate of Fiscal Incen-tives	Financial Subsidies		Rate on Fi-nancial Subsi-dies	Total Rate of Subsidies to Manu-factured Exports	(λ') Calculated by Eliana Cardoso
	(I) ICM	(II) IPI	(III) Import Tax draw back	(λ_1) (I+II+III)	(IV) IPI Bonus Credit	(V) ICM Bonus Credit	(VI) Income Tax Cuts	(λ_2) (IV+V+VI)	(VII) Credit Subsi-dies	(VIII) Back on Compul-sory Deposit	(λ_3) (VII+VIII)	(λ) ($\lambda_1+\lambda_2+\lambda_3$)	
1964	—	0,4	—	0,4	—	—	—	—	—	—	—	0,4	0,4
1965	—	5,0	—	5,0	—	—	—	—	—	—	—	5,0	5,0
1966	—	5,0	—	5,0	—	—	—	—	—	—	—	5,0	5,0
1967	16,1	5,2	—	21,3	—	—	—	—	—	—	—	21,3	21,9
1968	19,6	6,0	—	25,6	—	—	—	—	0,6	—	0,6	26,2	26,5
1969	20.5	6,8	0,7	28,0	4,3	—	—	4,3	1,7	—	1,7	34,0	31,8
1970	20.5	7,0	1,9	29.4	6,0	5,1	—	11,1	3,3	—	3,3	43,8	38,9
1971	19.8	7,5	2,4	29,7	6,4	5,9	1,3	13,6	4,2	—	4,2	47,5	41,3
1972	19,1	8,1	2,6	29,8	6,9	6,6	1,3	14,8	3,9	—	3,9	46,5	42,3
1973	16,3	9,8	3,5	31,6	7,0	7,0	1,3	15,3	3,6	—	3,6	50,5	43,4
1974	17,7	10,0	2,8	30,5	8,5	8,5	1,8	18,8	3,2	—	3,2	52,5	47,0
1975	17,0	10,0	4,6	31,6	10,1	10,1	1,7	21,9	5,6	3,2	8,8	62,3	49,1
1976	16,3	10,9	4,4	31,8	13,2	13,2	1,3	27,7	9,7	5,0	14,7	74,0	49,3
1977	16,3	12,0	2,9	31,2	11,2	11,2	1,5	23,9	12,3	4,1	16,4	71,5	50,6
1978	16,3	12,3	4,0	32,6	12,0	12,0	1,5	25,5	10,5	4,4	14,9	73,0	—

SOURCE: Explanations in text.

Of all these subsidies, tax exemptions were certainly the most important, in terms of both duration and in relative size. The second most important in magnitude were the fiscal incentives. Last, but actually not least, are the financial subsidies, which have become more important in recent years.

The final columns of Table 1 gives the evolution of the total rate of subsidies as calculated by Cardoso. The following figure compares over time our total and hers. While over the first few years the time series are practically equal, from 1969 on our data begin to diverge in an increasingly positive direction. This fact is due to Cardoso's not having considered the financial subsidies and tax drawbacks used to give new stimulus to exports. After 1975, the distance between the two broadens even further due to the intensification of incentive policies – this time with the increase in IPI and ICM bonus rates [8] – strengthened by the introduction of the compulsory deposit for imports and its

TOTAL SUBSIDY RATES FOR MANUFACTURED EXPORTS



8 Our original data sources reveal this increase. This point will be discussed extensively below.

counterpart in the form of a drawback on the import of inputs going directly into exports; and also due to the increase in credit incentives (a larger share of export deals could be financed, while nominal interest rates were held constant despite growing inflation).[4]

What follows is a description of the methodology used to estimate the respective subsidy rates.

2.1 — Tax exemptions

2.1.1 — IPI and ICM exemptions

Our calculation of the rates of IPI and ICM exemptions was basically equal to Cardoso's, except for the first ICM values for 1967 and 1968, mainly due to conflicting sources.[5]

2.1.2 — Import tax drawback

Drawback is defined as the suspension, exemption or restitution — either total or partial — of the import tax due on inputs

[4] Other expectations were not considered, such as those offered by the Befiex program, since they would be impossible to estimate. This program gained importance during the latter part of the period under study, when more emphasis was given to promoting manufactured exports. It offers additional subsidies in the form of tax exemptions for the import of machines and inputs utilized in domestic production by companies signing long-term export contracts. Other exemptions not considered here, which have been in effect ever since the promotion policies began, are: a) Minerals Tax (IUM); b) Financial Operations Tax (IOF); c) Fuel and Lubricants Tax (IUCL).

[5] Our source is the Department of the Treasury of the State of Bahia. The domestic rates it reported for domestic operations in the Central-South Region in these two years are: 15%, following February 1967, when the exemption was first granted; 15% from June 1, 1968 to March 31, 1968; 16% from April 1, 1968 to April 30, 1968; and 17% following May 1, 1968. These domestic rattes, applied to the FOB price, P_M, are converted into rates based on factory prices, P_F, in order to be expressed in terms compatible with the basis for applying the rates on domestic operations. The conversion follows the following formula: $P_M = P_F\ (1 + t_F)$ and $P_F = P_M\ (1 - t_M)$. Substituting the terms, we obtain the desired conversion, $t_F = t_M / (1 - t_M)$; since t_M is a known value, t_F can be obtained, which gave the following values for the respective periods: 17.7, 19.1 and 10.5%. The average rate for 1967 is thus:

$$100\ [1{,}177)^{11/12}\ .\ 1^{1/12} - 1] = 16{,}1\%;\ \text{and}$$

$$17{,}7^{3/12}\ .\ 19{,}1^{1/12}\ .\ 20{,}5^{8/12} = 19{,}6\%.\ \text{for 1968.}$$

imported directly by the exporter of the final product. The average exemption rate is:

$$\theta = \frac{\text{Value of import tax exemption}}{\text{Value of manufactured exports}}$$

To obtain the value of import tax exemption, we would have to know the various tariffs in effect for each product in each year, and apply each rate to all imports made under the drawback arrangement. Since this is not possible, due to the difficulty in obtaining each tariff rate and the subsequent value of goods imported, we have calculated — for each year — an average tariff for imports:

$$\zeta = \frac{\text{Calculated value of import taxes}}{\text{Value of total imports}}$$

The numerator of this ratio includes taxes that were calculated and paid as well as those that were calculated and exempted, in order to make a better estimate of the average tariff applicable to non-exempt imports, especially inputs.[6] The average rate of exemptions under the import tax drawback would thus be:

$$\theta = \zeta . \frac{\text{Value of imports subject to drawback}}{\text{Value of manufactured exports}}$$

or,

$$\theta = \zeta . \beta,$$

where β is a coefficent representing the percentage of imported inputs entering directly into manufactured exports.[7]

Column III of Table 1 shows the estimates we obtained. The behavior of the time series suggests the possibility of speculative stockpiling of imported inputs in 1975 and 1976, followed by a reduction of inventories in 1977. The same cycle seems to have occurred in 1973 and 1974, respectively.

[6] The numerator will include the taxes calculated both for imports subject to drawback and for the imports of the public sector (including state-owned companies). The average should thus be less than if calculated on the basis of tax revenues, where the incidence of final goods, which pay higher rates, is greater. Due to the lack of statistical information, we were only able to construct this series from 1971 on. The 1969 and 1970 values were extrapolated from the subsequent tendency. Our sources were: *Anuário Econômico-Fiscal*, 1972-77, and *Anuário Estatístico CIF-MF*, 1977-78.

[7] The data on the value of imports subject to drawback were taken from Relatório da CACEX, 1977 and 1978.

2.2 — Fiscal incentives

2.2.1 — IPI bonus credit

The rates at which this incentive was applied are taken from Cardoso's study. From 1974 on, they are calculated from the ratio between the IPI credits and the value of manufactured exports. [8] The largest average rates for this incentive were obtained in the later years, which could indicate they are the result of either a deliberate policy of increasing the bonuses or a change in the composition of exports, with an increase in the share of those exports enjoying the highest bonuses. [9]

2.2.2 — ICM bonus credit

The rates at which the ICM credits were calculated are generally the same as for the IPI. The procedure it thus nearly the same, except for one small difference: in this case there are deductions for the importation of inputs beginning in the year when this benefit was created (1970) until they were eliminated (1972). [10] The rate of this incentive, during this period, is thus obtained by correcting the IPI rate by the coefficient of the direct national value added $(1-\beta)$:

$$\text{Rate of ICM credit} = \text{Rate of IPI credit } (1-\beta)$$

2.2.3 — Income tax reduction

This series was taken entirely from Cardoso's study, which in turn got its data from the words of Varsano (1979) and Savasini, Kume and others (1979). Value added for 1978 is a repetition of the same value for 1977.

[8] The source for this data is *IPI — Informações Tributárias*, MF/SRF/CIEF, 1974-78.

[9] This latter effect could have been induced by the deliberate bias in favor of financial subsidies, which were significant in precisely these years.

[10] Until July 20, 1972, the IPI bonus credit included benefits on the portion of imported products, which were excluded after this date by Portaria No. 182. The ICM bonus credit never included this portion; thus the need for correction. The 1972 rate for this incentive is obtained by the geometric average of the IPI and ICM credits (after correction) weighted by the period in which each was in effect.

2.3 — Financial subsidies

This type of subsidy includes: a) subsidies granted to finance the export of manufactured goods, at nominal interest rates below the market rates; and b) compulsory deposits under the drawback arrangement.

2.3.1 — Credit subsidies

The credit systema at preferential interest rates granted to manufactured exports may be classified into three groups: a) financing of exportable production; b) financing of actual exports; and c) financing of export-oriented services.[11] Table 2 gives a summary of this credit program for exports in consignment, listing the various credit lines and their respective characteristics, user costs and terms of payment.

In order to simplify estimation of the subsidies given under each credit line, we will assume equal terms of payment for all loans. The period we have selected is one commercial year. The subsidy rates will be calculated for each of three basic processes: a) domestic financing in cruzeiros; b) domestic financing in dollars; and c) overseas financing in dollars. In each case, we must assume rational expectations by the economic agents in relation to the expected discount rate and the expected rate of currency devaluation. The following paragraphs present the methodology used in estimating the subsidy rate in each case.

2.3.1.1 — Domestic financing in cruzeiros

In this process, financing is contracted and payment is received in cruzeiros, that is, received and paid in cruzeiros:

$$I_1^{Cr\$} = I_0^{Cr\$}\,(1 + i)$$

where:

$I_1^{Cr\$}$ = nominal value of loan contracted in cruzeiros at the end of the period, assumed to be the commercial year.

11 Some of the credit lines in category *c* did not seem to apply to manufactured exports, but they are always intended to stimulate these exports. For example, in the third and fourth lines of category *c*, the priority projects for financing are those which will lead to the later export of manufactured goods. The second line of the same category is aimed at promoting the formation of overseas marketing facilities. In any case, as we shall see, these lines account for an insignificant percentage of the total.

$I_0^{Cr\$}$ = face value of the loan in cruzeiros; and
i = interest rate charged for each credit line evaluated by this proces.

By discounting the nominal value of the debt at the domestic market's average annual rate, we have:

$$\frac{I_1^{Cr\$}}{1+i_m} = I_0^{Cr\$}\,\frac{(1+i)}{1+i_m}$$

where i_m is the average annual market rate, identified as the cost of money to the taker of 360-day bills of exchange. The absolute value of the subsidies will be the difference between the face value and the discounted value of the debt, that is:

$$\text{Abs. Sub.} = I_0^{Cr\$} - \frac{I_0^{Cr\$}}{1+i_m} = I_0^{Cr\$}\left[1 - \frac{(1+i)}{1+i_m}\right]$$

$$\text{Abs. Sub.} = I_0^{Cr\$}\,\frac{(i_m - i)}{1+i_m},\ \text{ou}$$

$$\text{Abs. Sub.} = I_0^{Cr\$}\cdot\gamma$$

where $\gamma = \frac{i_m - i}{1+i_m}$, which is the present value of the subsidy per cruzeiro of credit taken. Since the value of the credit is not usually equal to the total value of exports it will be necessary to distinguish the percentage representing the ratio between the amount of credit and the value of manufactured exports, which will be called ϕ.

Taking that ratio into account, our subsidy rate (Π') is:

$$\Pi' = \frac{\text{Abs. Sub.}}{\text{Value of manufactured exports in Cr\$ } (VME^{Cr\$})} =$$

$$= \frac{I_0^{Cr\$}}{VME^{Cr\$}}\cdot\gamma$$

therefore:

$$\Pi' = \phi\cdot\gamma$$

The value of ϕ is obtained by dividing total loans under the respective credit line, centered in the middle of the year, by the value of manufactured exports, both in cruzeiros.[12]

12 The data on each mode of financing are from the Relatório da CACEX, 1978 and 1979; and from the Boletim do Banco Central, Mar. 1976, Aug. 1977,

Table 2

Summary of Programs to Finance Manufactured Exports

	Credit Lines	Nature	Interest Rate to User (% yearly)	Terms
	Credit line established by Central Bank Resolution No. 71, Nov. 11, 67	Covers loans for production of manufactured goods to be exported	8	1 year
A)	Credit line established by Banco do Brasil (14—11)	Complements work in capital of firms producing exportable manufactures with a production cycle of up to 180 days.	16.8	180 days
	Special credit line from CACEX with FINEX resources for capital goods (Resolution 318, Feb. 22, 1975)	Finances current production costs of production of capital goods with high unit price and production cycle of over 180 days	12	Production Time plus 60 days
	Direct export financing	Directly finances exports through discount of overseas sales vouchers for exports by Brazilian firm	7 to 8, plus exchange correction	More than 180 days
B)	Direct financing of foreign importer (Concex Resolution 68)	Allows foreign importer to pay cash in Brazil for Brazilian manufactured goods.	7 to 8, plus exchange correction	1 to 8 years, or more
	Financing of exports on consignment (Concex Resolution 43)	Covers shipping costs of consignment exports of capital and durable consumer goods, and others, exceptionally	16.8	180 to 360 days
	Export financing established by Central Bank Resolution 352 (Dec. 22, 1975)	Incentive for exports of capital and durable consumer goods	7 to 8	1 to 5 years
	Credit for overseas promotion and marketing	For current expenses with overseas promotion and marketing	15	1 year
	Credit for overseas investments	Finances establishment of companies outside Brazil, or participation in the capital of foreign firms	15	Max. 3 years
	Credit for Sale of technical-economic and engineering studies and projects	Finances the export of services by companies of this sector, to enable the sale of finished or nearly finished Brazilian products	7 to 8	Variable
(C)				
	Credit for purchase of equipment	Allows purchase, in Brazil, of machinery and equipment to be used in overseas projects by Brazilian engineering firms	12	Variable (limited to duration of project)
	Credit for production of export projects	Financial support for preparation of bid proposals and/or projects for overseas sale of services by Brazilian companies	12	3 years
	Entrepôt financing (Central Bank Resolution 230, July 16, 1976)	Covers vouchers for manufactured exports stored in bonded customs warehouses, presented by trading companies	12	180 days

We have used this process to estimate the following credit lines: all the Group A lines (financing of exportable production); financing of exports on consignment in Group B (financing of actual exports); and the Group C lines (financing of export-promotion services), except for the credit line for sale of technical-economic and engineering studies and projects for overseas undertakings.

2.3.1.2 — Domestic financing in dollars

Although these are domestic credits, financing is contracted in dollars converted to cruzeiros at the exchange rate of the day:

$$I_1^{Cr\$} = I_0^{US\$} \cdot E_0 (1 + \delta)\,(1 + i')$$

where:

$I_1^{Cr\$}$ = value of the debt contracted at the end of the period, in cruzeiros;

$I_0^{US\$}$ = face value of the loan in dollars;

E_0 = exchange rate on the day the loan was contracted;

σ = annual rate of currency devaluation, including differential between buying and selling rate; [13] and

i' = annual interest rate charged in this mode of financing.

The discounted value of the debt, at the average annual market rate, will be:

$$\frac{I_1^{Cr\$}}{1 + i_m} = I_0^{US\$} \cdot E_0 \, \frac{(1 + \delta)\,(1 + i')}{1 + i_m}$$

Sept. 1979. Since the figures refer to the end of each year, we had to calculate the average of the year-end balances of two successive years to center the amount of financing in the middle of each year. i_m is the monthly geometric average centered in the middle of each year. The domestic interest rate is that of 360-day bills of exchange, as published in *Conjuntura Econômica*, Jan. 1975, Sept. 1977, and *Boletim do Banco Central*, Sept. 1979.

[13] The currency exchange rate for credit takers is obtained by dividing the exchange rate for sale at the end of each year by the purchasing rate at the beginning of the year, less one. Both taken from *Conjuntura Econômica*, Nov. 1972 (appendix), Jan. 1975 and *Boletim do Banco Central*, Sept. 1979.

The absolute value of the subsidies will be equal to the difference between the face value and the discounted value of the loan:

$$\text{Abs. Sub.} = I_0^{US\$} \cdot E_0 \left[1 - \frac{(1+\delta)\,(1+i')}{1+i_m} \right]$$

$$\text{Abs. Sub.} = I_0^{US\$} \cdot E_0 \left[\frac{(1+i_m) - (1+\delta)\,(1+i')}{1+i_m} \right]$$

$$\text{Abs. Sub.} = I_0^{US\$} \;.\; E_0\, \gamma',$$

where γ' is the present value of the unit subsidy in this mode, represented as:

$$\gamma' = \frac{i_m - (\delta + i' + \delta\, i')}{1 + i_m}$$

The subsidy rate for exports financed domestically in dollars, Π'', is obtained as the product of γ' by the share of these loans in total manufactured exports, ϕ'. Thus,

$$\pi'' = \phi' \gamma'$$

This process is used to calculate the subsidy rates of the following credit lines; direct export financing and export financing under Resolution n.º 352 of the Central Bank, both in Group B; and credit for the sale of technical-economic and engineering studies and projects for overseas jobs, in Group C.

2.3.1.3 — Overseas financing in dollars

This process makes a direct calculation of subsidies to foreign importers which stimulate exports of Brazilian manufactured goods. Here financing is contracted in dollars. The dollar value of the debt at the end of the period, at interest rate i'', is:

The present value of the debt, discounted at the overseas market interest rate, i_e, is: [14]

$$\frac{I_1^{US\$}}{1 + i_e} = I_0^{US\$}\, \frac{(1 + i'')}{1}$$

where the discount rate, i_e, is the average annual interest rate on the Eurodollar market.

14 The overseas interest rate is the annual rate on the Eurodollar market, taken from *International Financial Statistics*.

The absolute subsidy is:

$$\text{Abs. Sub.} = I_0^{\text{US\$}} \left[1 - \frac{1 + i''}{1 + i_e} \right] = I_0^{\text{US\$}} \frac{(i_e - i'')}{1 + i_e}$$

The subsidy rate under this process, Π''', is obtained by multiplying the unit subsidy $\gamma'' = \dfrac{i_e - i''}{1 + i_e}$ by the fraction of these loans over the total value of manufactured exports, ϕ'':

$$\pi''' = \phi'' \cdot \gamma''$$

This process is used to estimate the subsidy rate of the direct credit line to foreign importers in Group B.

2.3.1.4 – Credit subsidy rates

Table 3 presents a time series of the credit subsidy rates, Π, for each credit line designed to promote manufactured exports.

The direct credit lines for exporters and for foreign importers are taken together, since available data on the amount of credit granted does not discriminate between the two. Thus the subsidy rate on these credits is a geometric average of the subsidy rates obtained for each credit line.

The same table shows that the three types of loans to exportable production receive more subsidies than any other. It also shows that credits for exports and for foreign imports represent one percent or less of all credit subsidies. Compared to these, the other credit lines are insignificant.

The aggregate subsidy rate is the sum of the individual rates, and is given in the last line of Table 3.

2.3.2 – Drawback on the compulsory deposit for imports

To estimate the subsidy generated by the exemption from the compulsory deposit, we must calculate the opportunity cost of these resources in the economy (or absolute subsidy), which will be the difference between the face value, D, and the discounted value at the average annual market rate. That is:

$$\text{Abs. Sub.} = D \left[1 - \frac{1}{1 + i_m} \right] = D \cdot \frac{i_m}{1 + i_m}$$

Table 3

Credit Subsidy Rates (Π) for Export of Manufactured Goods (in Percent of FOB Value)

	Credit Lines / Years	1968	1969	1970	1971	1972	1973	1974	1975	1976	1977	1978
(A)	Credit Line Established by Central Bank Resolution 71	0,59	1,63	2,47	2,91	2,72	2,38	2,05	3,76	6,02	8,04	7,02
	Credit Line Established by Banco do Brasil (14—11)	—	—	0,52	0,88	0,08	0,57	0,50	0,72	1,04	1,22	0,97
	Special Credit Line From CACEX with FINEX Resources for Capital Goods	—	—	0,06	0,05	0,03	0,05	0,15	0,47	1,14	1,72	1,34
(B)	Direct Export Financing	—	0,08	0,23	0,20	0,00	0,53	0,30	0,35	0,32	0,81	1,02
	Direct Financing of Foreign Importer	—	—	—	0,01	0,01	...	...	...	...	...	0,02
	Export Financing Established by Central Bank Resolution 352, with FINEX Resources	—	—	—	—	—	—	—	—	...	0,01	0,01
(C)	Credit for Overseas Promotion and Marketing	—	—	0,01	0,01	0,02	0,02	0,02	0,10	0,21	0,15	0,04
	Credit for Overseas Investments	—	—	—	—	—	—	...	...	...	0,02	0,05
	Credit for Sale of Technical-Economic and Engineering Studies and Projects for Overseas Jobs	—	—	—	—	—	—	—	—	...	...	0,01
	Credit for Purchase of Equipment	—	—	0,04	0,04	0,02	0,05	0,13	0,18	0,40	0,24	...
	Credit for Production of Export Projects	—	—	—	—	—	—	—	—	—	0,01	...
	Entrepôt Financing	—	—	—	—	—	—	—	—	0,01	0,06	0,03
	Credit for Production of Export Projects	—	—	—	—	—	—	—	—	—	0,01	...
	Total	0,59	1,71	3,33	4,19	3,68	3,60	3,24	5,58	9,74	12,28	10,61

SOURCE: explanations in text.
OBS.: — = credit line not in operation
... = insignificant estimates

The subsidy rate, Π, is obtained by dividing the absolute subsidy by the value of manufactured exports. Thus:

$$\bar{\pi} = \frac{D}{VEM} \quad \frac{i_m}{(1 + i_m)}$$

But, since the total of compulsory deposit exemptions is equal to the value of drawback requests, the preceding expression becomes:

$$\bar{\pi} = \beta \cdot \gamma$$

where β is once again the percentage of drawback requests out of total exports, and $\bar{\gamma} = \frac{i_m}{1 + i_m}$, the present value of the unit subsidy per cruzeiro on compulsory deposits.

Initially, compulsory deposits had to be repaid within six months (Resolution n.º 331, Central Bank), under rules that were in effect from June 16 to December 2, 1975; thereafter, this was extended to one year (Resolution n.º 354, Central Bank). Rate $\bar{\gamma}$ for 1975 is thus a weighted geometric average:

$$\gamma_{75} = \left(1 + \frac{i_m/2}{1 + i_m/2}\right)^{11/24} \cdot \left(1 + \frac{i_m}{1 + i_m}\right)^{2/24} \cdot 1^{11/24} - 1$$

The first parenthesis corresponds to the six-month capitalization, weighted by 11 of the 24 forthnights in the year: a period approximately equal to that in which Resolution n.º 331 was in effect. The second parenthesis corresponds to the yearly capitalization rate during the two fortnights of the same year in which Resolution n.º 354 was in effect. The unitary term ($1^{11/24}$) weights out the rest of the year, when these deposits were not required, and during which the subsidy rate is zero.

Comparing columns III and VIII in Table 1, we can see that, with the exception of 1975, more stimulus was given to exports through the compulsory deposit drawback than through the import tax drawback.

3 — Subsidies and the behavior of manufactured goods exports

3.1 — Supply function

In evaluating economic policies, it is crucial to be able to know the operational efficiency of the various trade policy instruments used to promote exports. To this end, we will use Cardoso & Dornbusch's general supply model, specified as:

$$x = a + bp_j + cy + u \qquad (1)$$

where x is the natural logarithm of the ratio between the quantum index of manufactured exports and the quantum index of industrial production; p_j, the natural logarithm of the international price index of manufactured exports in cruzeiros, relative to the domestic wholesale price index for manufactures; y measures cyclical excess of demand; and u is the error.

While deriving this supply model, we assumed that Brazil is a price taker in the international market for manufactures, that is, that price is exogenously set. The relative price in cruzeiros, P_j, will be:

$$P_j = \frac{P^* \, E_j}{P}$$

where:

P^* = International price index in dollars of our exports of manufactured goods;

E_j = Exchange rate in cruzeiros for dollars for each type of subsidy;

P = Domestic wholesale price index for manufactures.

In order to test the significance of the subsidy policies, we have defined six alternate exchange rates, tied to six relative prices:

1) $PA = \dfrac{P^* E}{P}$;

2) $PB = \dfrac{P^* E (1 + \lambda_1)}{P}$;

3) $PC = \dfrac{P^* E (1 + \lambda_2)}{P}$;

4) $PD = \dfrac{P^* E (1 + \lambda_3)}{P}$;

5) $PE = \dfrac{P^* E (1 + \lambda_1 + \lambda_2)}{P}$;

6) $PF = \dfrac{P^* E (1 + \lambda_1 + \lambda_2 + \lambda_3)}{P}$;

and for each relative price we will make the respective adjustment. We will first consider the relative price dictated by the behavior of international prices and by exchange policies, *PA*. This will then be replaced by the relative price which includes both the exchange rate and tax exemptions, *PB*. The next step is the relative price defined by the exchange rate and only the fiscal incentives, *PC*. Next we introduce the relative price with only the credit subsidies, *PD*. After taking each subsidy by itself, we turn to the relative price with the exchange rate and all fiscal subsidies, *PE*. Finally, we calculate the relative price defined by the effective exchange rate, that is, including all subsidies granted under the export promotion system, *PF*.[15]

The cyclical variable is measured by the difference between real industrial output, set by demand, and projected industrial output, identified as the tendencial behavior of industrial output itself.[16]

Table 4 gives the results obtained from fitting Equation 1 to the annual data from the sample, for the period 1959/78. Besides the respective original equations, we have included the equations corrected by autocorrelation of the residuals, using the Hildreth-Lu procedure; except for the final adjustment, the value of the constant is not given, since it is a sevendigit number that is statistically insignificant.

The first estimated equation reveals the effect of the relative price dictated by the behavior of the international price of manufactures and by exchange policies practiced in each period. Cardoso and Dornbusch (1980*b*) and Balassa (1979) have emphasized the upward tendency of this relative price in determining the continued growth of the competitiveness of Brazilian manufactures on the world market. This equation, however, presents a determination coefficient of only 50% and the presence of autocorrelation of the residuals. Once the residuals are corrected, the fit becomes unstable, that is, the constant explodes and the relative price variable loses its significance. The coefficient of autocorrelation is $p = 1.0$.

Estimating the equation with the relative price that includes tax exemptions brings an appreciable improvement in the determination coefficient — which rises to 75% — as well as

15 In the estimations with price variables *PD* and *PF*, the respective errors in the model should include the "white noise" resulting from the hypothesis of rational expectations introduced in the financial subsidy estimates.

16 We also attempted to estimate planned industrial output through an autoregression process. But the results of the cyclical variable in Equation 1 were improved with the introduction of a trend.

Table 4

Results of Fitting Equation 1 to Different Relative Prices

	Equations (for Each Exchange Rate)	Constant	Coeficient P_1	Coeficient y	ρ	DW	R^2	SER
1	Exchange Rate Alone	−6,5	1,33	−2,1		0,965	0,5	0,24
		(−4,16)	(4,09)	(−2,9)				
	E	Explosive	0,11	−1,87	1,0	1,76	0,8	0,15
			(0,42)	(−2,33)				
2	Exchange Rate Including Tax Exemptions	−5,05	1,[illegible]5	−1,53		1,51	0,75	0,17
		(−7,34)	(7,18)	(−3,45)				
	$E(1+\lambda_1)$	Explosive	0,14	−1,87	1,0	1,78	0,8	0,15
			(0,52)	(−2,4)				
3	Exchange Rate Including Fiscal Incentives	−5,34	1,1	−2,13		1,19	0,68	0,19
		(−6,14)	(6,0)	(−3,85)				
	$E(1+\lambda_2)$	Explosive	0,13	−1,91	1,0	1,76	0,8	0,15
			(0,48)	(−2,33)				
4	Exchange Rate Including Financial Subsidies	−5,95	1,22	−2,12		1,08	0,6	0,21
		(−5,14)	(5,05)	(−3,36)				
	$E(1+\lambda_3)$	Explosive	0,09	−1,85	1,0	1,75	0,8	0,15
			(0,33)	(−2,39)				
5	Exchange Rate Including Fiscal Exemptions and Incentives	−4,18	0,87	−1,52		1,5	0,79	0,15
		(−8,2)	(8,0)	(−3,77)				
	$E(1+\lambda_1+\lambda_2)$	Explosive	0,15	−1,9	1,0	1,78	0,8	0,15
			(0,56)	(−2,41)				
6	Effective Exchange Rate Including Fiscal Exemptions and Incentives and Financial Subsidies	−3,9	0,81	−1,51		1,5	0,8	0,15
		(−8,46)	(8,24)	(−3,84)				
	$E(1+\lambda_1+\lambda_2+\lambda_3) = E(1+\lambda)$	−3,8	0,79	−1,57	0,25	1,68	0,81	0,15
		(−6,29)	(6,10)	(−3,01)				

OBS: t statistic in parentheses.

diminishing the problem of autocorrelation of the residuals. This leads us to conclude that the tax exemptions granted to promote Brazilian manufactured exports were important in increasing their international competitiveness. However the adjustment to correct the autocorrelation of the erros still gives problems with instability.

The third adjusted equation certainly gives better results than the first, but not as good as the second. We can thus conclude that fiscal incentives helped improve the competitiveness of these exports, but not as much as the tax exemptions. We still have the problem of instability in the equation corrected by autocorrelation of the residuals.

Equation 4 also fit the data batter than the first equation, revealing that financial subsidies also contributed to the expansion of manufactured exports. But, of the three types of export promotion policies, this one had the least relative importance. The problem of instability of the equation corrected by autocorrelation of the residuals remains.

The fifth estimated equation shows that tax exemptions taken in tandem with fiscal incentives produced better results than if they had been taken separately. But the problem of instability of the equation corrected by autocorrelation of the residuals still persists.

Finally, the last estimated equation — which takes into account the relative price in terms of the effective exchange rate, that is, which includes all three types of subsidies — does not offer results that differ significantly from the fifth; but it does give us stability in the equation that corrects the series correlation. From which we can conclude once again that financial subsidies are important in the behavior of manufactured exports.

There are some purely methodological considerations to be dealt with at this point. The fifth estimated equation — which leaves financial subsidies out — gives two *minima* for the function of standard error of the regression (SER) in relation to the autoregression of the errors: the relative minimum is $p = 0.25$ and the absolute minimum is $p = 1.0$. The same function for the final estimated equation which includes all the subsidies — displays the opposite behavior: a relative minimum at $p = 1.0$ and an absolute minimum at $p = 0.25$. [17] Therefore, if we had used the Cochrane-Orcutt procedure to correct the autocorrelation of the errors, this method's limitation would have produced stability in the estimation of the fifth equation. We might thus

[17] See Table 2, in the Statistical Appendix.

have erroneously concluded that the financial subsidies did not make a significant contribution to explain the behavior of manufactured exports.

In short, in addition to the favorable behavior of international prices and the indexation of the exchange rate, all three kinds of subsidies played significant roles in the expansion of manufacture exports. While Table 4 suggests that the tax exemptions were relatively more effective than the fiscal incentives, this is actually not the case. The three subsidies are not being appraised under equal conditions. We saw in Table 1 that over the entire sample period tax exemptions were the largest and most long-lasting of all the subsidies, followed by fiscal incentives and then by financial subsidies.

If we respecify the supply function, taking into account the lagged adjustment mechanism, similar to the one proposed by Cardoso and Dornbusch (1980a) – with the relative price logarith associated with the effective exchange rate P_F, and correction of autocorrelation of the errors – we obtain results very similar to theirs; now for the period of 1960/78:

$$x_t = -\underset{(-3,14)}{2,21} + \underset{(3,2)}{0,465P_{Ft}} - \underset{(3,4)}{1,175y_t} + \underset{(3,1)}{0,53x_{t-1}}$$

$$p = -0,25; \quad R^2 = 0,865; \quad \text{e} \quad SER = 0,13 \qquad (2)$$

Comparatively, our estimation reveals a fall of approximately 10% in the relative price logarithm and 2% in the coefficient of the cyclical variable. Table 5 summarizes the estimates for price-elasticity or supply-substitution of manufactured exports.

Table 5

Elasticities: Substitution of Supply of Exports

Equation		Short Term	Long Term
Ours	1	0.81	0.99
	2	0.465	
Cardoso and Dornbusch	1	0.83	
	2	0.51	1.09

4 — Conclusions

This study has advanced in quantifying the massive programs that subsidize Brazilian manufactured exports. We have demonstrated that each of the subsidies and all of them together have been used by the export sector, which reacted by expanding exports as a share of total sectoral output. However it was impossible to reach a conclusion as to the relative effectiveness of each of the policy instruments.

It is clear that all the components of the relative price of manufactured exports — the world market price, exchange policies and all of the subsidies — had decisive effects on the behavior of the sector's exports. And the exports, in turn, reacted positively and significantly in the same proportion to all stimuli to their relative price.

Statistical appendix

Table 1

Variables Used

Years	Indices (base: 1970 = 100) — Quantum: Manufactured Exports	Quantum: Industrial Production	P^* Overseas Price Manufactures	E Nominal Exchange Rate of Exports	P Domestic Price of Manufactures	Relative Price of Manufactured Exports at each Exchange Rate: PA	PB	PC	PD	PE	PF	y Cyclical Variable (%)
1959	35,0	48,0	83,0	3,40	2,00	142,3	110,0	128,1	137,8	101,3	99,0	8,60
1960	35,0	52,0	74,0	4,20	2,00	104,6	80,7	91,0	101,1	74,4	72,6	9,51
1961	39,7	58,2	77,6	6,00	4,20	110,8	85,7	99,8	107,3	78,9	77,1	11,40
1962	32,3	62,8	77,1	8,50	6,10	107,4	83,0	96,7	104,0	76,5	74,7	10,80
1963	32,8	62,0	86,9	12,0	11,1	98,6	76,2	88,8	95,5	70,2	68,0	2,67
1964	34,8	68,1	95,0	27,9	20,4	130,7	101,4	117,7	126,6	93,4	91,3	—0,61
1965	56,2	63,0	85,8	41,4	32,8	109,3	87,9	97,6	104,8	80,9	79,1	—13,70
1966	58,6	69,2	89,5	48,0	43,5	100,0	81,1	90,0	96,8	74,7	72,0	—12,50
1967	65,7	71,3	87,7	58,3	54,6	93,0	87,8	84,3	90,7	80,9	79,0	—17,80
1968	69,9	80,8	90,4	73,9	71,2	93,8	91,1	84,5	91,4	83,9	82,3	—13,50
1969	85,5	90,0	91,7	88,7	85,6	98,1	97,1	92,1	96,6	92,4	91,4	—10,30
1970	100,0	100,0	100,0	100,0	100,0	100,0	100,0	100,0	100,0	100,0	100,0	—8,60
1971	104,0	114,3	111,0	115,0	118,0	108,2	108,4	110,6	109,1	110,3	111,0	—3,51
1972	155,0	129,6	124,0	129,0	136,0	117,6	118,0	121,5	118,3	121,1	121,5	0,78
1973	177,0	150,1	178,0	133,0	156,0	149,2	151,7	151,8	149,6	150,0	150,2	7,23
1974	194,0	164,9	240,0	148,0	202,0	180,2	181,8	192,7	180,1	191,5	191,1	8,39
1975	206,0	175,2	244,0	177,0	202,0	164,8	167,6	180,9	173,6	180,1	186,0	6,20
1976	214,0	193,9	233,0	233,0	387,0	159,1	154,7	174,8	168,0	172,4	184,0	8,10
1977	265,0	201,4	261,0	308,0	497,0	161,7	164,0	180,4	182,3	178,6	192,9	3,06
1978	340,0	217,8	240,0	394,0	673,0	144,0	147,6	162,7	160,2	162,1	173,3	3,24

SOURCES: 1. Quantum indices, P^*, E and P: *Conjuntura Econômica*, Nov. 1972 (appendix), Jan. 1976, Apr. 1977, July and Dec. 1979; *Boletim do Banco Central*, Sept. 1979.

2. PA, PB, PC, PD, PE, PF: construed from their respective definitions in the text and information presented in Table 1.

3. y is identified by the residuals of the equation that fits the tendency of the natural logarithm to the index of industrial output, that is:

$$m = \underset{(82)}{3.7} + \underset{(22)}{0.08}\,T, \quad R^2 = 0.90.$$

Table 2

Relations Between RHO (ρ) and the Standard Error of the Regression *

(ρ) RHO	Standard Error of the Regression — Equation 1	
	Exchang Rate Including Fiscal Subsidies $E\ (1 + \lambda_1 + \lambda_2)$	Effective Exchange Rate $E\ (1 + \lambda_1 + \lambda_2 + \lambda_3)$
0,00	0,1580	0,1543
0,05	0,1570	0,1530
0,10	0,1560	0,1520
0,15	0,1550	0,1513
0,20	0,1548	0,1509
0,25	0,1547[a]	0,1508[b]
0,30	0,1549	0,1510
0,35	0,1550	0,1514
0,40	0,1560	0,1520
0,45	0,1564	0,1530
0,50	0,1570	0,1540
0,55	0,1580	0,1545
0,60	0,1582	0,1550
0,65	0,1584	0,1560
0,70	0,1582	0,1562
0,75	0,1580	0,1561
0,80	0,1565	0,1560
0,85	0.1550	0 1550
0 90	0.1540	0,1540
0,95	0,1530	0,15283
1,00	0,1525[b]	0,15281[a]

NOTES: *Negative values of ρ have been disregarded.
[a] Relative minimum.
[b] Absolute minimum.

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On the causes of the recent inflationary acceleration *

André Lara Resende **
Francisco Lopes **

1 — The econometrics of Brazilian inflation

Brazilian inflation has usually been analysed on the basis of the model of the Phillips curve which postulates an inverse relationship between inflation and product gap, as in the following equation:

$$\hat{P} = -a\ (y^r - y) + \hat{P}^e;\ a > 0 \qquad (1)$$

where y^r is the natural logarithm of the potential product, y the logarithm of the product, $(y^r - y)$ approximately the product gap, $\hat{P}$ the rate of inflation and $\hat{P}^e$ the expected rate of inflation.

The theoretical interpretation of this trade-off and the confidence in its stability over time have suffered great changes since Phillips' original work appeared. However, such an equation, or other similar formulation, was estimated with data on the Brazilian economy in various works with apparently very satisfactory results. [1]

Equation (1) can be deducted from three basic equations:

$$u - \bar{u} = a_1\ (y^r - y)\ ;\ a_1 > 0 \qquad (2)$$

$$\hat{P} = \hat{w} \qquad (3)$$

$$\hat{w} = -a_2\ (u - \bar{u}) + \hat{P}^e;\ a_2 > 0 \qquad (4)$$

Editor's note: Translation not revised by the author.

* Originally published in *Pesquisa e Planejamento Econômico, 11* (3) :599-616. December 1981.

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1 See, for example, the works of Lemgruber (1973 and 1974) and Contador (1977).

The first is Okun's law, which associates the deviations in the rate of unemployment, u, in relation to the natural rate of unemployment, $\bar{u}$, to the product gap. The second, assuming a constant mark up, states that prices grow according to costs of production, that is to say, wages, $\hat{w}$. And, finally, the third is a relation of the expected real wage adjustment to excess demand in the labour market, an equation, however, which assumes implicitly that wages are fixed in a totally independent way from the wage policy. One should remember that over the last 15 years there have been compulsory wage readjustments in Brazil, which could only be avoided by companies adopting the costly expedient of increasing the turnover of their labour force. In this sense, the Brazilian experience is unique, but its implication for the evolution of wages is totally ignored by equations (4) and (1).

Without assuming *a priori* whether the wage policy is irrelevant or not, one can formulate a model which allows such a hypothesis to be tested, dividing the labour market into two sectors.

In the first, called market sector, for which the adjustment equation is equivalent to the model of the Phillips curve, the salary, therefore, is not affected by the wage policy.

$$\hat{w}_1 = -b \ (y^r - y) + \hat{P}^e; \ b > 0 \tag{5}$$

On the other hand, in the second, called the institutional sector, wages depend on the legal minimum wage, $\hat{w}_{min}$, here being used as a proxy for the legal readjustment, and possibly also on excess demand.

$$\hat{w}_2 = \hat{w}_{min} - c \ (y^r - y) ; \ c > 0 \tag{6}$$

The growth rate of the average salary in the economy is obtained by:

$$\hat{w} = \hat{w}_1 + (1 - \alpha) \ \hat{w}_2; \ 0 < \alpha < 1 \tag{7}$$

where α is the relative weight of the market sector in the labour market.

Industrial prices are considered here to be obtained by applying a mark up rule on costs in accordance with:

$$P_I = (1 + m) \left[\frac{w}{g} + \frac{cP^*_m}{d^*} + \frac{Q}{d} \right] \tag{8}$$

where P_I is the industrial price, m the mark up, w the nominal wage, g the output/labour ratio, P^* the price in foreign currency of imported inputs, e the exchange rate, d^* the output/imported input ratio, d the output/domestic input ratio and Q the price of the domestic input.

This formulation, including imported inputs in industrial costs, allows one to consider the impact of external shocks (S), which seem to have played an important part in recent Brazilian inflation. Taking d, d^* and the mark up m as constant and expressing equation (8) in terms of rates of variation, we have:

$$\hat{P}_I = \lambda_0 \, (\hat{e} + \hat{P}^*_m) + \lambda_1 \, (\hat{W} - \hat{g}) + \lambda_2 \hat{Q} \tag{9}$$

Substituting (5) and (6) in (7), and then (7) in (9), one obtains:

$$\hat{P}_I = \lambda_0 \, (\hat{e} + \hat{P}^*_m) + \lambda_1 \alpha \, \hat{P}^e - \lambda_1 \hat{g} + \lambda_1 \, (1 - \alpha) \, \hat{w}_{min} -$$
$$- [\lambda_1 \alpha \, b + \lambda_1 \, (1 - \alpha) \, c] \, (y^p - y) + \lambda_2 \hat{Q} \tag{10}$$

It is presumed, furthermore, that the expected inflation is simply past inflation and the growth rate in the prices of domestic inputs is the average rate of growth of current industrial prices and past inflation:

$$\hat{P}^e = \hat{P}_{-1} \tag{11}$$

$$\hat{Q} = \delta \hat{P}_I + (1 - \delta) \, \hat{P}_{-1} \tag{12}$$

It is convenient, in order to obtain the foreign inflationary pressure variable in terms of shocks and to test the relative importance of the labour market sector, that the variables be defined as $\hat{\delta} = \hat{e} + \hat{P}^*_m - \hat{P}_{-1}$ and $\hat{w} = \hat{w}_{min} - \hat{g}$.

Thus, equation (10) can be rewritten as follows:

$$\hat{P}_I = \gamma_0 \hat{\delta} + \gamma_1 \hat{\omega} + \gamma_2 \hat{P}_{-1} + \gamma_3 \hat{g} + \gamma_4 \, (y^p - y) \tag{13}$$

where

$$\gamma_0 = \frac{\lambda_0}{1 - \delta\lambda_2}$$

$$\gamma_1 = \frac{\lambda_1 \, (1 - \alpha)}{1 - \delta\lambda_2}$$

$$\gamma_2 = \frac{\lambda_0 + \lambda_1 \alpha + \lambda_2 \, (1 - \delta)}{1 - \delta\lambda_2}$$

$$\gamma_3 = \frac{\alpha\lambda_1}{1 - \delta\lambda_2}$$

$$\gamma_4 = \frac{[\lambda_1 \alpha b + \lambda_1 (1 - \alpha) c]}{1 - \delta\lambda_2}$$

The results of the estimation of equation (13) are shown in Table 1.

As the rate of growth of industrial prices can be one of the explanatory variables both of the nominal wage and of the rate of exchange – especially the latter, due to the system of mini-devaluations, more or less in accordance with the parity of purchasing power, followed by Brazil during the greater part of the period of the sample – these two variables are not exogeneous to the model. To overcome this difficulty, the model was estimated using instrumental variable method, the results of which are shown in equations (1) and (2) of Table 1. The variables P^*_m, $\hat{P}_{-1}$, $\hat{g}$, $(y^p - y)$, the constant, the dummy for 1963 and the rate of growth of the import quantum served as instruments. A dummy was introduced for the year 1963 when it was observed that the error in this year is systematically about three times the standard error of the regression. Structural change, still not fully understood, seems to have occurred in 1963.

Both the product gap as well as the *per capita* income growth rate have insignificant coefficients (and with opposite signs to those expected). Therefore, y_3 and y_4 are statistically not different from zero (note that if $\gamma_3 = 0$, then $\alpha\lambda_1 = 0$). As $\lambda_1 \neq 0$ as one can see from the fact that $y_1 \neq 0$, one can conclude that $\alpha = 0$, that is to say, the market sector is negligible in the labour market.

If $\alpha = 0$ and $y_4 = 0$, then $\lambda_1 c = 0$, which means that $c = 0$, that is to say, the response of wages in the institutional sector to demand pressures is insignificant.

The restricted model, without a constant and with $y_3 = y_4 = 0$, was estimated and the results are in equation (2). In equations (3) and (4), estimated by ordinary least squares, the exaggerated values of the coefficients of the external shock variable confirm the suspicion that the exchange rate is not an exogenous variable in the model. [2]

[2] Note that the coefficients of the variables $\hat{\delta}$ and $\hat{\omega}$ are not exactly those of imported inputs λ_0 and wages λ_1 in the total costs, but are expanded by the factor $\frac{1}{1 - \delta\lambda_2}$.

Table 1

Equation of Industrial Prices — 1960/78 *
(Dependent Variable: $\dot{P}_I$)

	Constant	Independent Variables					
		$\dot{i}$	$\dot{\omega}$	$\dot{P}_{-1}$	$\dot{g}$	(y^P-y)	Dummy*
Equation (1) R^2 = 0.97 SE = 0.04 DW = 2.30	0.0267 (0.45)	0.2770 (2.58)	0.6034 (2.68)	0.3513 (1.44)	0.1719 (0.33)	0.0738 (0.33)	0.2828 (5.32)
Equation (2) R^2 = 0.97 SE = 0.03 DW = 1.59	— —	0.3803 (4.98)	0.4219 (3.78)	0.5545 (5.05)	— —	— —	0.2927 (7.22)
Equation (3) R^2 = 0.98 SE = 0.03 DW = 2.08	0.023 (0.50)	0.388 (5.93)	0.465 (6.64)	0.454 (4.46)	0.073 (0.17)	0.036 (0.29)	0.254 (7.43)
Equation (4) R^2 = 0.98 SE = 0.03 DW = 1.95	— —	0.431 (7.29)	0.426 (6.77)	0.552 (10.03)	— —	— —	0.200 (7.91)

The values in brackets are the statistic t. The questions (1) and (2) were estimated using the instrumental variable method and equations (3) and (4) y using ordinary least squares.

Refers to 1963.

The apparently favourable results obtained in other work for the model of the Phillips curve, where the gap appears significantly among the inflation determinants, disappear when a more complete model is estimated. The inclusion of variables that capture the shock and wage policy effects in the equation of industrial prices makes the trade-off between inflation and product gap disappear. This apparent trade-off has been utilized to justify the need for recessive policies in order to successfully fight the current Brazilian inflation. Contador,[3] for example, bases his argument on a graph (like Graph 1, which follows), interpreting it on the basis of the interaction between gap and expectations according to the traditional model of Phillips' curve, but Graph I can be interpreted according to the alternative model of equation (3). The 1964/67 period, when inflation was greatly reduced, coincides with the wage policy of the first post-1964 government, which, as has been widely studied, exercised strict control over the minimum wage, while the 1973/74 years, when inflation accelerated, coincides with a period of external shock, due to the increase in the prices of petroleum. These are precisely the two periods in which the gap and inflation are moving in the direction prescribed by the Phillips curve. In fact, small alterations in the gap in these two periods seen to be associated with great variations in the rate of inflation. If the estimated model does not take into account the two important variables related to external shocks and to wage policy, there will clearly be a statistical fabrication which will make the inverse relation between gap and inflation significant despite periods such as 1967/73, when the gap had a much greater variation and inflation remained stable or slightly on the decline.

2 — Simulation for 1979/80

Equation (2), besides providing us with a satisfactory explanation for the evolution of the industrial price index in the 1960/78 period, was also utilized to simulate the behaviour of industrial prices in 1979 and 1980, in an attempt to identify the factors responsible for the recent inflationary acceleration (the basic data used in the simulations are shown in Table 2).
We have, then, for 1979:

$$\begin{aligned} \hat{P}_I &= 0.3803\ (34.5\%) + 0.4219\ (51.3\%) + 0.5545\ (38.9\%) \\ &= \underset{\text{(external shock)}}{13.1\%} + \underset{\text{(wages)}}{21.6\%} + \underset{\text{(other inputs)}}{21.6\%} = 56.32\% \end{aligned}$$

[3] See Contador (1980).

and for 1980:

$$\begin{aligned}\dot{P}_I &= 0.3803\ (74.20\%) + 0.4219\ (84.0\%) + 0.5545\ (55.4\%)\\ &= \underset{\text{(external shock)}}{28.2\%} + \underset{\text{(wages)}}{35.4\%} + \underset{\text{(other inputs)}}{30.7\%} = 94.4\%\end{aligned}$$

It can be seen that the value of $\dot{P}_I$ estimated for 1979 is exceptionally close to the actual figure, which, as Table 2 shows, was 55.6% (the error of the estimate is only 0.72 percentual points).

Furthermore, of the 20% increase in the dependent variable between 1979 and 1978, 13 percentual points are explained by the external shock component, while the other seven are explained by the wages component; reflecting the only moderate impact on the annual average wage for 1979 of the change in the wages policy at the end of that year. It should be noted, however, that the real minimum wage (deflated by the wholesale price index-domestic supply) remained constant.

The result of the simulation for 1980 is much less satisfactory than for 1979: the equation explains industrial price inflation of 94.4%, which remains, however, 9.1 percentual points below the observed value (estimated) of 103.5%. This error is almost three times greater than the standard error of the equation (3.4%), which suggests the occurrence of a structural change in the dynamics of industrial prices. The following section will try to show that this discrepancy is a consequence of the alteration in the wage policy in November 1979.

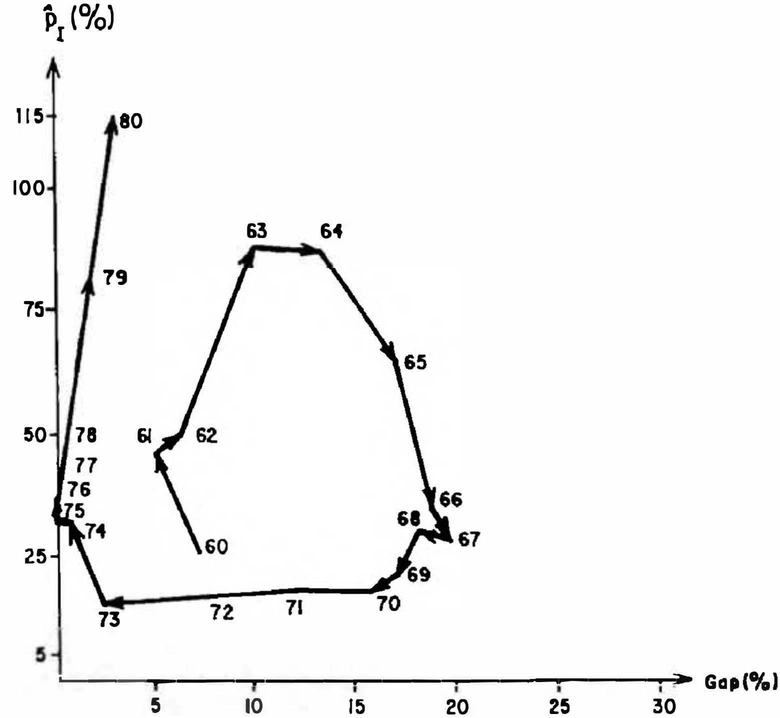
It is interesting to note that the external shock component is responsible for more than half the increase of about 48 percentual points in the observed value (estimated) of $\dot{P}_I$, which clearly shows — together with what was found for the previous year — the importance of this factor in explaining the recent acceleration in the inflationary process.

One can use the model to assess the importance of the 30% devaluation which occurred in December 1979, within this external shock component of the inflationary acceleration.

Assuming that the traditional practice of minidevaluations had remained in force throughout the period and that it was equivalent to:

$$\tilde{e} = \dot{P}_I - \dot{P}^* \tag{14}$$

Graph 1



where $\hat{P}^*$ is the external rate of inflation. It follows that:

$$\hat{S} = \hat{e} + \hat{P}^*_m - \hat{P}_{-1} = \hat{P}_I + (\hat{P}^*_m - \hat{P}^*) - \hat{P}_{-1} \qquad (15)$$

Assuming that the external inflation was $\hat{P}^* = 10\%$ and utilizing the data from Table 2, we have $\hat{S} = \hat{P}_1 - 15.2\%$. Therefore:

$$\hat{P}_I = 0.3803\,(\hat{P}_I - 35.4\%) + 35.44\% + 30.7\% \qquad (16)$$

which results in $\hat{P}_I = 84.9\%$. Comparing this value with the estimate of 87.7%, produced by the equation when the variation in the exchange rate which effectively occurred is included, shows that abandoning the parity rule of equation (14) resulted in almost 10 percentual points in the 1980 inflation.

Table 2

Basic Data Used in the Simulations
(Annual Average Rates, in %)

Years	Variables							
	$\hat{P}_I$	$\hat{P}_{-1}$	$\hat{\omega}_{min}$	$\hat{g}$	$\hat{w}$	$\hat{c}$	$\hat{P}^{*}_{m}$	$\hat{S}$
1978	35.3			3.1			5.0	
1979	55.6	38.9	55.1	3.8	51.3	48.4	25.0	34.5
1980	103.8	55.4	87.1		84.0	99.6	30.1	74.20

$\hat{g}$: growth of GDP *per capita*, with growth of GDP of 6% and demographic growth of 2.8%
$\hat{P}^{*}_{m}$: 0.35 (Δ petroleum price) + 0.65 (Δ non-petroleum price) = 0.35 (75%) + 0.05 (6%).
$\hat{S}$: external shock

3 — The inflationary impact of the change in wage policy

In November 1979, a new Brazilian wage policy came into force, replacing the previous practice of annual wage readjustments for one of half-yearly readjustment. In this section we will try and determine the magnitude of the inflationary impact resulting from this change.

Graphs 2 and 3 show that a reduction in the interval between wage readjustments results in an acceleration of inflation in an economy where profit margins are kept constant. Up to the moment T the economy remains in inflationary equilibrium with annual wage readjustments; after T the readjustments become half-yearly. In the graphs, which show the evolution over time of the real wage logarithm of a representative class of workers, we see that the real wage decreases at a constant geometric rate (therefore, linear in log) in the 12 month period between the two readjustments of the nominal wage, with the rate equal to the rate of inflation, and that the annual (geometric) average real wage remains constant up to T.

Graph 2 shows that, if the inflation rate does not change after T, the annual average real wage then increases due to the greater frequency of the readjustments, which, however, means a reduction in profit margins, which contradicts our initial assumption.[4] In fact, the only way of making more frequent wage

[4] That is to say $p = (1 + m)\ bw$, where p is the price, m the profit margin, b the labour/output ratio and w the nominal wage. Then, we have $(1 + m)\ b\,(w/p) = 1$, which shows that when b is constant there exists an inverse relation between real wage and profit margin.

readjustment compatible with unaltered profit margins is through an increase in the rate of inflation after T, as is shown in Graph 3. In this case, the average real wage does not alter, despite the reduction in the interval between readjustments, but the change in the wage policy produces, on the other hand, an inflationary shock.

It is possible to obtain a formal derivation of this result, that allows us also to determine the extent of this inflationary impact. Let us assume that wage readjustments are made annually and that the economy can be divided into 12 productive sectors with equal shares in the aggregate product, each sector renegotiating its collective labour contract in a different month of the year. Let us also assume that each sector readjusts its price only once a year, immediately after the wage readjustment of its workers. Let p_t be the logarithm of a general price index and $p_t\ (k)$ the logarithm of the price in the month k. Let us accept as a simplication that the general price index is a geometric average of sectoral prices, so that:

$$p_t = \frac{1}{12} \sum_{k=1}^{12} p_t (k) \tag{17}$$

The variation in month t of any variable x_t, will be shown by dx_t. As each productive sector of the economy only readjusts its price once a year, we have $dp_t\ (k) = 0$ if $k \neq t$, which means that the monthly rate of inflation is obtained by:

$$dp_t = \frac{1}{12}\ dp_t\ (t) \tag{18}$$

The economy operates with constant profit margins, so that the increase in the price of each sector is a weighted average of the wage readjustment and the increase of the average price of the intermediate goods:

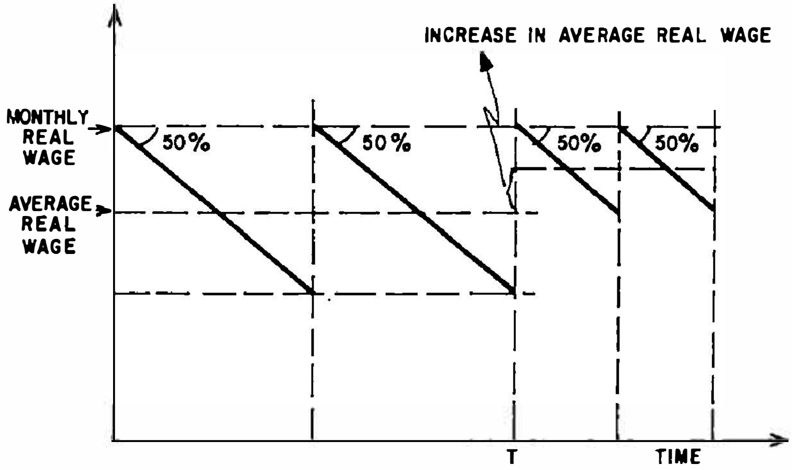
$$dp_t (t) = \gamma dw_t (t) + (1 - \gamma)\ dq_t (t) \tag{19}$$

and we can also accept that the annual price increase of the intermediate goods incorporated in the price readjustment is equal to the inflation accumulated in the previous 12 months, that is:

$$dg_t (t) = \sum_{j=1}^{12} dp_{t-j} \tag{20}$$

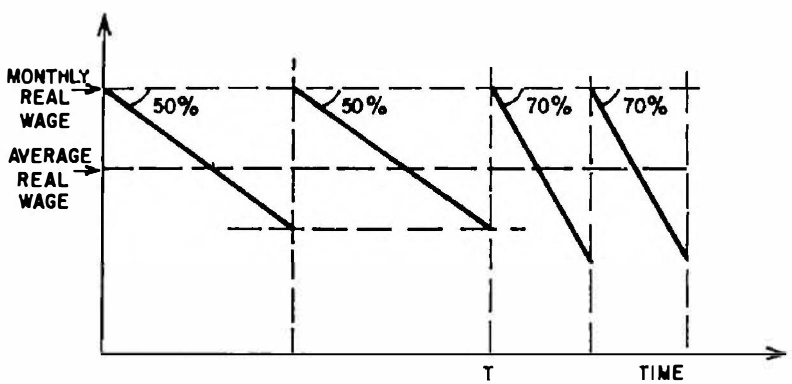
Graph 2

CHANGE TO THE RULE OF HALF-YEARLY WAGE READJUSTMENT WITH CONSTANT RATE OF INFLATION



Graph 3

CHANGE TO THE RULE OF HALF-YEARLY WAGE READJUSTMENT WITH CONSTANT AVERAGE REAL WAGE



Assuming that the wage policy rule corrects the nominal wage according to full inflation of the previous 12 months, we have:

$$dw_t(t) = \sum_{j=1}^{12} dp_{t-j} \tag{21}$$

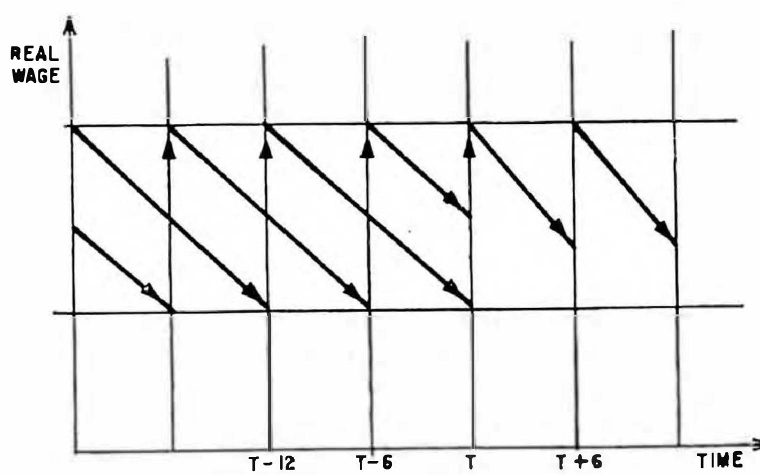
and therefore:

$$dp_t = \frac{1}{12} \sum_{j=1}^{12} dp_{t-j} \tag{22}$$

This last equation shows that, if $dp_{t-j} = z$ for $j = 1, 2, \ldots 12$, then $dp_t = z$, which characterizes inflationary equilibrium at the monthly rate of z.

Let us assume now that society decides to adopt half-yearly wage readjustments and the change-over from the existing rule of annual readjustments to the new system occurs during six months. In the first month T in the six-month transition period there will occur, simultaneously, a readjustment on the annual basis of workers' wages, the previous readjustment having been made 12 months before (in $T - 12$), and a readjustment on a half-yearly basis of workers who had their last readjustment six months before (in $T - 6$). As Graph 4 shows, after T these two

Graph 4



groups of workers start to have wage readjustment on a half-yearly basis. Repeating the same process in each month of the transition period, one will have, at the end, all the workers of the economy receiving readjustments on a half-yearly basis. [5]

What happens to the rate of inflation as a consequence of this transition from annual to half-yearly wage readjustments? Note that our equation (18) now has to be replaced by:

$$dp_t = \frac{1}{12} \left[dp_t(t) + dp_t(t - 6) \right] \tag{23}$$

If we accept, as a simplification, that the increases in the price of intermediate goods continue to be passed on to the prices of products on an annual basis, as in equation (20), we have:

$$dp_t(t - 6) = \gamma \sum_{j=1}^{6} dp_{t-j} \tag{24}$$

and therefore:

$$dp_t = \frac{1}{12} \sum_{j=1}^{12} dp_{t-j} + \frac{\gamma}{2} \left(\frac{1}{6} \sum_{j=1}^{6} dp_{t-j} \right) \tag{25}$$

Assuming an inflationary equilibrium up to month T with a monthly rate of inflation of z, so that $dp_{T-j} = z$ for $j = 1, 2, \ldots 12$, we have:

$$dp_t = z + \frac{\gamma z}{2} \tag{26}$$

which shows an inflationary acceleration in the first month of the transition phase.

[5] The shategy related to the change in Brazilian wage policy was slighth different from the one we are simulating here as there was a readjustment in November 1979, our month T, corresponding to the accumulated infation of the previous six months (22%) for all the workers who had had their last wage readjustment between the months $T - 12$ and $T - 6$, and from then on all wages started to be readjusted on a half-yearly basis. It is not reasonable, however, to assume that the whole wage readjustment of November had been simultaneously passed on to the prices, which justifies the simplification we are adopting.

To calculate the rate of inflation in the second month of the transition phase, we note that:

$$dp_{T+1} = dp_T + \frac{1}{12}\,(dp_{T-12}) + \frac{\gamma}{12}\,(dp_T - dp_{T-6})$$

$$= z + \left(\frac{13+\gamma}{12}\right)\frac{\gamma z}{2} \tag{27}$$

Repeating the calculation, we obtain for the sixth month of the transition period:

$$dp_{T+5} = Z\;\frac{13+\gamma^{5}}{12}\;\frac{\gamma z}{2} \tag{28}$$

If we accept, as a simplification, that a new inflationary equilibrium is eventually reached at this final monthly rate of inflation, which is represented by z', and introducing the value $\bar{\gamma} = 0.42$ obtained in Section 1, we have:

$$z' = 1.369\ z \tag{29}$$

as a measure of the inflationary impact of the change from annual to half-yearly wage readjustments. Taking the rate of inflation of 55% in 1979 (see Table 2) and assuming $z = (55\%)/12 = 4.583\%$, we obtain $z' = 6.27\%$, which is equivalent in annual terms to 75.3% (always maintaining the linear approximation that is being used here). If we add this value to the estimate of 24.0% for the external shock component calculated in the previous section, we get an estimate of 99.3% = (75.3% + 24.0%) for the inflation rate in 1980, which is very close to the actual value of 103.8% (see Table 2).

Our theoretical model also provides an explanation for the discrepancy of 19.8 percentual points between the rate of inflation (of industrial prices) for 1980 and the corresponding value projected by equation (2) of Section 1. It is easy, though rather tedious, to see that:

$$dp_{T+5} = \frac{1}{12}\sum_{j=1}^{12} dp_{T+5-j} + \left\{12 + \gamma + \gamma\left[\frac{13+\gamma}{12}\right] + \right.$$

$$\left. + \dots \gamma\left[\frac{13+\gamma}{12}\right]^{4}\right\}\frac{\gamma'}{24} \tag{30}$$

As the equation of regression of Section 1 only incorporates the first term of this equation, therefore, one can predict *a priori* a discrepancy between the econometrically estimated value equal to the second term. Substituting the values for z and γ, we obtain an estimate of 14.4 percentual points for this discrepancy of 19.8 percentual points found in the previous section.

4 — Conclusion

The econometric exercise of Section 1 suggests that, when wages and external shocks, caused by an increase in the domestic cost of imported inputs, are explicitly considered, the trade-off between the inflation — measured by industrial prices — and the product gap disappears.

The traditional inverse relationship between the rate of unemployment or the product gap — since these two variables are alternative measures of the degree of slack in the economy and form a stable relationship according to Okun's Law — and the rate of inflation of industrial prices apparently cannot be found in the data for the Brazilian economy in the last two decades. In this case, the possibility of exercising control over aggregate demand to combat inflation disappears.

Such a result is to say the least surprising for those that are accustomed to extrapolate the dynamics of prices in competitive microeconomic markets to the economy as a whole and to the labour market in particular. If a reduction of aggregate demand raises the unemployment rate to higher levels than the natural or equilibrium rate, in principle the growth rate of nominal wages should be reduced. Even if the mark ups are insensitive to demand, the reduction in the rate of growth of wages will reduce the rate of inflation.

This work questions the possibility of such a market mechanism working in an economy like the Brazilian one, where wages are corrected by law on the basis of past inflation in which case it only will not be passed on in full to costs if the expedient of increasing the labour turnover is used. Such an expedient is, however, extremely costly for the companies that invest in personnel selection and training and imposes an inevitable strain on company/worker relations, with negative impact on productivity.

The examination of the effects of a change from annual to half-yearly wage readjustments, presented in Section 3, shows also that the practice of wage indexation has important effects on real wages and the rate of inflation. A correct understanding of this relationship is vital if a policy for fighting inflation is to be designed without having perverse side-effects on real wages and the level of employment.

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On the causes of the recent inflationary acceleration: a comment *

Claudio R. Contador **

1 — Introduction

In a recent study, Resende and Lopes (1981) examined the causes of the acceleration in inflation in 1980 through an equation in which prices are explained by the wage policy and by external shocks. On the basis of empirical results, they conclude that there is no significant trade-off between inflation and product gap.

This paper questions these empirical results and shows that the employment of a longer period would lead to opposite conclusions; the inclusion of shocks raises the significance level of the trade-off.

2 — General observations

The empirical analysis of Resende and Lopes covers the 1960/78 period with data relating to industry. The dependent variable is the rate of growth of the Industrial Product, Wholesale Prices Index, Total Supply (column 26 of *Conjuntura Econômica*).

It is important to mention that the empirical models of the Phillips curve, in its varied versions, are not concerned with an isolated sector. Other authors, like Contador (1977 and 1980) and Lemgruber (1973 and 1974) emphasized the trade-off between aggregate variables, like the gap in GDP, or its deviation in relation to the trend, and a general price index. This is the first divergence in relation to Resende and Lopes' model.

Editor's note: Translation not revised by the author.

* Originally published in *Pesquisa e Planejamento Econômico, 12* (2) :607-14, August 1982.

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In second place, the authors seem to be more concerned with forecast of the growth of industrial prices in 1980 than with an understanding of a more ample phenomenon. For a forecast, there exist other simpler methodologies, like that of Box and Jenkins (1970), which do not require exact theoretical models.

Thirdly, the formation of inflation expectations is very simple in Resende and Lopes' empirical model, which presupposes the price variation in the previous period as the best expectation for the present. The formation of expectations is one of the most critical variables in these models, and the conclusions against the significant trade-off may result from Resende and Lopes' specification. However, this is not a crucial aspect in our comments.

Resende and Lopes' analysis is, nevertheless, laudable in drawing attention to the effects of real external and domestic (wages) shocks on inflation. The criticism that follows is partial, as it does not include the effects of the new system of wage readjustments (Law no. 6.708, 1979), almost outside the period used for the empirical analysis.

3 — The empirical results

The specification in question is more conventional than that of Resende and Lopes and takes the form of:

$$\frac{P}{P_t} = a_0 + a_1 \frac{P}{P_{t-1}} + a_2 H_t + a_3 \frac{S}{S_t} + u_t \qquad (1)$$

where P corresponds to the price index, H is the product gap or the deviation in relation to trend, S is the external "shock" and u the aleatory residues. It is expected that: $a_1 > 0$; $a_2 < 0$ if H corresponds to the gap, or $a_2 > 0$ if related to the deviation in the real product in relation to tendency; and $a_3 > 0$. The external shock includes variations in the price of oil in dollars and of intermediate goods, besides variations in the rate of exchange.

Therefore there exist two nominal variables on the right-hand side of the reduced form (1): past inflation and the nominal shocks. Consequently, it is not possible to impose *a priori* the condition where the parameter a_1 will be close to unity, as is normally done in more orthodox models that postulate a zero trade-off in the long term. It should be noted that despite its orthodoxy, the reduced form (1) is similar to the final specification adopted by Resende and Lopes.

Using the wholesale prices index, industrial products (the same as used by Resende and Lopes), Table 1 shows some empirical results. The first three regressions are based on December to December variations (inexplicably not tested by Resende and Lopes) and the following three on variations of annual averages. The values in brackets below the parameters correspond to the statistic t and those below the multiple determination coefficient are the same parameter adjusted for degrees of freedom.

To represent the oil shock the growth rate in the dollar price of imported oil was used according to column 46 of the foreign trade price index, from *Conjuntura Econômica* and, for the price of imported intermediate goods, column 41, figures which are only available after 1962, price constancy (zero growth) being assumed for the previous period.

All the parameters show the expected sign and satisfactory level of significance, with the exception of the one relating to the gap in regression 1. There is no evidence of serial correlation in the residues.

Contrary to Resende and Lopes' conclusions, the inclusion of the "shock" variable raises the level of significance of the gap parameter. The strong colinearity between the variations in oil prices and intermediate goods does not allow the two variables to be included in the same regression. However, the level of significance of a_2 is still modest (10%) and the technique of constructing the gap is somewhat controversial.

Repeating the experiment with the deviation in the real industrial product in relation to the exponential trend, where parameter a_2 should now be positive, the empirical results are even better, as is shown in Table 2. The inclusion of external shocks definitely improves the explanatory power of the model of the Phillips curve in a less conventional version, resticted to the industrial sector.

It still remains to be seen if similar conclusions would be obtained for the real aggregate product. After all, the trade-off has been discussed more for aggregate variables like the GDP. So as not to make these comments too long, Table 2, which follows, shows only the empirical results of the conventional model without supply shocks and the modified model with oil price shocks (*PP*). The period is the same (1950/79) in Tables 1 and 2, and the price indices correspond to annual averages. The inflation is

Table 1

Explanation of Industrial Product, Wholesale Prices: Trade-Off Measured by the Gap in the Industrial Product — in the 1950/78 Period

Regression	Indices	Constant	Past Inflation	Gap*	Shocks		R^2 ($\bar{R}^2$)	DW
					Oil	Intermediate Consumption		
1	Dec./Dec.	0,1881 (2,90)	0,6629 (3,89)	– 0,7901** (–1,58)	—	—	0,361 (0,314)	2,13
2	Dec./Dec.	0,1718 (3,09)	0,4521 (2,76)	– 0,7949 (– 1,82)	0,3303 (3,07)	—	0,532 (0,478)	1,88
3	Dec./Dec.	0,1743 (3,29)	0,3993 (2,50)	– 0,7570 (– 1,82)	—	0,3972 (3,60)	0,574 (0,525)	1,77
4	Average	0,1215 (2,06)	0,8276 (5,99)	– 0,0999 (–1,86)	—	—	0,574 (0,525)	2,39
5	Average	0,1131 (2,66)	0,6856 (4,82)	– 0,6542 (–1,87)	0,1913 (2,41)	—	0,647 (0,621)	2,18
6	Average	0,1156 (2,80)	0,6418 (4,50)	– 0,6133 (–1,80)	—	0,2347 (2,68)	0,666 (0,627)	2,13

* Gap obtained by the difference (in logs) between the actual and potential GDP. The potential GDP was estimated by the exponential tangency to the actual GDP in 1949 and 1974.

** Non-significant at the 5% level in the one-tailed test ($a_2 < 0$).

Table 2

Explanation of Industrial Product, Wholesale Prices: Trade-Off Meassured by Deviation in Relation to the Trend — 1950/78 Period

Regression	Indices	Constant	Past Inflation	Deviation	Shocks		R^2 ($\bar{R}^2$)	DW
					Oil	Intermediary Consumption		
1	Dec./Dec.	0,1304 (2,21)	0,6396 (3,97)	0,9021 (1,98)	—	—	0,390 (0,345)	2,13
2	Dec./Dec.	0,1149 (2,33)	0,4302 (2,76)	0,8745 (2,20)	0,3245 (3,10)	—	0,555 (0,504)	1,87
3	Dec./Dec.	0,1199 (2,45)	0,3796 (2,50)	0,8360 (2,20)	—	0,3903 (3,62)	0,595 (0,548)	1,76
4	Average	0,0777** (1,58)	0,8079 (6,33)	0,7761 (2,32)	—	—	0,599 (0,569)	2,43
5	Average	0,0654** (1,59)	0,6706 (5,06)	0,7189 (2,30)	0,1852 (2,30)	—	0,667 (0,628)	2,21
6	Average	0,0705 (1,76)	0,6299 (4,74)	0,6821 (2,24)	—	0,2376 (2,67)	0,686 (0,650)	2,17

* Devation in actual GDP (in logs) in relation to exponential in the period.

** Non-significant at the 10% level.

measured by the general price index, domestic supply availability (column 2 of *Conjuntura Econômica*):

$$\frac{\Delta P}{P_t} = 0.0855 + 0.9023 \frac{\Delta P}{P_{t-1}} - 0.6556\, H_t \qquad (2)$$

$$(2.19) \qquad (7.23) \qquad (-1.78)$$

$R^2 = 0.674$; $\bar{R}^2 = 0.650$; DW + 1.92

$$\frac{\Delta P}{P_t} = 0.0841 + 0.7702 \frac{P}{P_{t-1}} - 0.6588\, H_t + 0.1690 \frac{PP}{PP_t} \qquad (3)$$

$$(2.33) \qquad (6.00) \qquad (-1.93) \qquad (2.35)$$

$R^2 = 0.731$; $\bar{R}^2 = 0.700$; DW = 1.73

Once again, the inclusion of the oil price shock improves the significance level of the trade off between the inflation and the gap in the real GDP. Imposing the condition of non-existence of the trade-off in the long term (that is to say, $a_1 = 1$), the natural rate of idle capacity is estimated at around 15%. * Furthermore, each 10% increase in the cruzeiro cost of imported oil has the effect of raising the idle capacity by 2.5 percentual points, keeping the rate of inflation constant.

4 — Conclusions

Thus, the empirical results do not reject the existence of a significant trade-off between the rise in prices and idle capacity, nor the upward shift in the curve in the presence of supply shocks. The important implication of these results for economic policy is that, for a given expected rate of inflation, a supply shock, like that of oil, has the effect of raising the natural rate of unemployment.

* If $a_1 = 1$ and assuming that the other parameters do not change, equation (1) from the test becomes:

$$\frac{\Delta P}{P_t} - \frac{\Delta P}{P_{t-1}} = a_0 + a_1 H_t + a_2 \frac{\Delta S}{S_t} + u_t$$

In the long term, ignoring the residue, we obtain:

$$H_t = \frac{a_0}{a_1} \div \frac{\Delta S}{S_t}$$

According to the estimated values in the last regression, we find that, when $\frac{\Delta S}{S_t} = 0$, the natural idle capacity is equal to 0.127 (= 0.084 ÷ 0.6588) and, for each 10% increase in the price of oil, the idle capacity increases 0.025 (= 0.169 ÷ 0.6588 × 0.1).

This means that to return to or maintain the equivalent economic growth is more inflationary than in the absence of shocks. Any similarity with what occurred in 1980 in Brazil, when maintaining an 8% growth in the GDP resulted in an inflation of 110%, or with 1981, when the inflation obstinately refused to decline, even at the cost of a severe cooling-off, cannot be considered as a mere coincidence.

A more elaborate model can and should incorporate a series of other shock variables. In this respect, Resende and Lopes' work is highly laudable in the sense of opening up the way. However, unemployment (measured by the gap, by the trend deviation, or any other variable) cannot be ignored. One of the criticisms of economic policy in 1981 is exactly the attempt to revive a model which was completely wrong as regards the political and economic conditioners. A previous work by Contador (1980) warned of this danger, showing that the cyclical phase experienced by the Brazilian economy in 1979 was the complete opposite of that of the beginning of the "miracle" in 1968. Unfortunately, Resende and Lopes' article can unintentionally give the idea that fighting inflation does not have serious short term repercussions on economic growth, independent of the cyclical phase that the economy finds itself in.

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On the causes of the recent inflationary acceleration: a reply *

Francisco Lopes **
André Lara Resende **

Written at the end of 1980, our work tried to exámine the causes of the acceleration in inflation which occurred in the second half of 1978. The analysis concentrated on the impact of three factors on the inflationary process: the increase in the international price of oil, the maxidevaluation of the cruzeiro and the wage policy. One result of the study was that, when these shock elements are incorporated into the model of the Phillips curve, similar to that used previously by Lemgruber (1974) and Contador (1977), it becomes impossible of obtain a significant estimate for the traditional inverse relation between the rate of inflation (measured by the wolesale price index, total supply, industrial goods and the product gap. In our opinion, the absence of a perceptible trade-off between these variables can be easily explained in an economy with compulsory and generalized wage indexation. In any case, however, it seems to us that it is obvious that the estimates then existing for the trade-off — which normally indicate relatively modest losses in the level of activities associated with substantial falls in the rate of inflation — have to be viewed with great suspicion.

In his comments on our article, Contador seeks to question this empirical result stating that "... the employment of a longer period would lead to opposite conclusions: the inclusion of shocks raises the significance level of the trade-off." It is clear, however, that the new econometric exercise which he presents is a long way

Editor's note: Translation not revised by the author.

* Originally published in *Pesquisa e Planejamento Econômico*, *12* (2) :615-22, August 1982.

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from being a mere extension of the sample period of our model, as stated.

Contador is guilty of two shortcomings that completely misrepresent the question analysed in our work. One of them concerns the exchange policy: as the measure of external shock, he uses the rate of growth of the oil price *in dollars*, arbitrarily assuming zero growth before 1962. [1] The use of this variable may be acceptable if the exchange devaluations had been determined throughout the whole period analysed by a stable rule of indexation, keeping the parity of purchasing power in relation to the dollar. It is known however, that this did not happen (it should be noted that the period analysed is 1950/79). Instead of introducing an unnecessary problem of variable error, Contador would have done better to follow our own methodology, that uses the exchange rate to build an index in cruzeiros of imported inputs and estimates its regression through instrumental variables. Even better would have been to use the price in cruzeiros of an important oil derivative (like fuel oil), which. however, none of us had the idea of doing. The policy of subsidizing the internal price of oil, carried out up to 1979, produced a very different evolution of this price in relation to the international price (we have more to say about this later).

The second shortcoming in Contador's comments is, in our opinion, the most important. As is known, after 1965 a wage policy came into existence in the Brazilian economy, which determines in a compulsory and generalized way the minimum wage readjustments for all workers legally employed. In the first few years, at least until 1968, the wage indexation was imperfect and did not accurately accompany the evolution in the cost of living. [2] In this period, however, the wage policy introduced a deflationary shock in the economy, and it was this consideration that led to the inclusion of the minimum wage variation in our regression.

An alternative technique, that tries to capture the wage shock through a *dummy* variable, produced the results in Table 1 (it should be noted that the *dummy* allows the inflation lag coefficient to be less in the 1965/67 period). Both in the case of the wholesale price index, domestic supply, general (WPI — DS — General), and in the cost of living index, Rio de Janeiro (CLI-RJ), the regressions with the *dummy* show estimates of lower statistical significance for the gap coefficient (the values in brackets are the t — statistics

1 The justification for this hypothesis of zero growth was the inexistence of data for the period prior to 1962 in *Conjuntura Econômica*. As one is dealing with price *in dollars*, it is curious that Contador was not capable of producing a series of prices before this year.

2 See Simonsen (1974) in this respect.

Table 1

Estimates With Wage Policy Dummy (Dummy for 1965/67)

1 — WPI — DS
1952/81 Period
OLSQ

a — Without dummy

$$\dot{q}_d = \underset{(1.90)}{0.103} - \underset{(-2.44)}{1.025}\, H + \underset{(8.74)}{1.065}\, \dot{q}_d\,(-1) \qquad R^2 = 0.74,\ DW = 1.96,\ SER = 0.140$$

b — With dummy

$$\dot{q}_d = \underset{(0.89)}{0.049} - \underset{(-1.15)}{0.522}\, H + \underset{(9.58)}{1.145}\, \dot{q}_d\,(-1) - \underset{(-2.21)}{0.364}\,[\delta \cdot \dot{q}_d\,(-1)] \qquad R^2 = 0.78,\ DW = 2.32,\ SER = 0.130$$

2 — CLI—RJ
1953/79 Period
OLSQ

a — Without dummy

$$\Delta \dot{q}_c = \underset{(2.16)}{0.087} - \underset{(-2.36)}{0.807}\, H \qquad R^2 = 0.19,\ DW = 1.79,\ SER = 0.114$$

b — With dummy

$$\Delta \dot{q}_c = \underset{(1.15)}{0.042} - \underset{(-0.04)}{0.015}\, H - \underset{(-3.38)}{0.361}\,[\delta \cdot \dot{q}_c\,(-1)] \qquad R^2 = 0.46,\ DW = 1.76,\ SER = 0.095$$

Definition of symbols:
δ = 1 in the years 1965/67;
$\dot{q}_d$ = rate of variation of WPI-IA, annual average;
$\dot{q}_c$ = rate of variation of CLI-RJ, annual average;
H = GDP gap, potencial product growing at 7% a year, zero gap in 1970.

of the parameters). Despite the correct coefficient sign, it is not possible to reject the hypothesis of a zero coefficient, at a significance level of 5%.

The deflationary shock during the 1965/67 (1968?) period is not, however, the only difficulty that the wage policy introduces in the econometric analysis of the Brazilian inflationary process. It seems reasonable to accept that wage indexation was almost perfect from 1969 on (taking into consideration the wage bonus in 1974). Why, then, have we not eliminated from our sample the "difficult" years of 1965/68 and carried out the analysis with the data that remain, without considering the wage policy?

The answer is that, even if the wage indexation had been 100% perfect from 1969 on, one could not guarantee *a priori* that the gap coefficient would be the same both in the period prior to 1965 (when there was no formal indexation mechanism) and in the period of full indexation after 1969. There exists a comprehensive foreign literature on the practical difficulties of evaluating the inflationary impact of wage and price control policies, which emphasizes precisely the inadequacy of regressions that mix observations for periods with control, with observations for periods without control.[3] The estimates 1 and 2 of Table 2, for the same equation of the Phillips curve in the 1952/64 and 1969/81 periods, seem to confirm this objection. In any case, to estimate a regression with a 29 year sample (1950/79), for an economy that underwent a series of structural and institutional changes in the period, in our opinion, shows an unjustifiable confidence in the econometric method.

Probably, the safest strategy to avoid these difficulties, associated with the wage policy, is to work with a sample for the 1969/81 period, as we did in estimates 2 and 4 Table 2. It should be noted that, besides the gap coefficient losing significance (and showing the wrong sign) when we introduce external shock variables in estimates 3 and 4, a better statistical adjustment is also obtained when we utilize the rate of variation in the price in cruzeiros of fuel oil, as we have suggested previously. It is also worth noting that in estimate 3, when we utilize the rate of variation in the price in dollars of imports, the inflation lag coefficient appears with an unrealistically high value, while in estimate 4 the sum of the coefficients of inflation lag and the rate of variation in the price of fuel oil is approximately unitary, as one would expect.

As is known, small modifications in the hypotheses of a model or in the set of data used to estimate it, or even in the estimation

[3] For example, Lipsey and Parkin (1970) and Oi (1976).

Table 2

Estimates With External Shock
(Dependent Variable: Rate of Variation of WPI-IA, Annual Average)

Estimate (Period) /Method of Estimate	Independent Variable							
	Constant	GDP Gap	Rate of Variation-Dependent Vareble Lag	Rate of Variation-Prince in Dollars of Imports (General Index	Rate of Variation-Price in Cruzeiros of Fuel oil	R^2	DW	SER
1 — (1952/64) OLSQ	−0,704 (−0,69)	1,837 (1,21)	0,889 (3,68)			0,75	2,25	0,124
2 — (1969/81) OLSQ	−0,137 (−1,12)	−0,834 (−0,69)	1,792 (5,28)			0,77	1,58	0,131
3 — (1969/81) OLSQ	0,302 (−2,78)	0.223 (0,47)	1,940 (7,43)	0,645 (2,92)		0,88	1,71	0,098
4 — (1969/81) OLSQ	−0,380 (−0,91)	0,325 (1,63)	0,305 (2,34)		0,738 (9,17)	0,97	2,78	0,043

SOURCE: *Conjuntura Econômica* and CNP.

technique, allow an infinite number of combinations and even the possibility that radically different results may be obtained for the same phenomenon. In the last few decades, the more widespread use of econometrics and the countless unresolved debates have taught us that the adoption of a positivist stance does not transform economics into a natural science. More than ever before, it is necessary now to understand and critically evaluate hypotheses which form the basis for interpreting the facts. Contador, judging from his comments, does not understand the model presented in our work and the complex questions that motivated it. His comments are simply one more reestimation of the classical model of the Phillips curve for the Brazilian economy, the results of which are well-known. Our point is that the strong negative correlation between inflation and product gap in the sample period may be a consequence of the exclusion from this model of two key variables in the Brazilian inflationary process: wage policy and external shocks. The first was totally omitted and the second incorrectly measured by Contador. In this case, it is not surprising that the correlation reappears, which in our opinion may be spurious, between inflation and gap. Before debating the level of significance of the statistical coefficients, it is necessary to understand what is actually being discussed. If not, then the debate runs the risk of instead of contributing to knowledge, simply producing a lot of noise.

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Inflation and the level of activity in Brazil: an econometric study *

Francisco Lopes **

This work examines the empirical relation between inflation and the level of activity in Brazil. It starts with a critical review of some existing econometric studies, including our own previous work on the subject, with André Lara Resende. Then, a new form is put forward for modelling the inflationary process in an economy with unsynchronized wage indexation. Estimates given by this model with annual data for the period 1969/81 suggest that: a) both the industrial product gap and its variation contribute statistically to the explanation of the inflationary process; but b) the impact of variations of the level of activity on the rate of inflation is very small.

1 — Introduction

The relationship between rate of inflation and level of economic activity is a key factor in the design of a stabilization policy. The problem of stabilization always involves a choice between: a) a conventional strategy, based exclusively on the active control of nominal demand; b) alternative strategies that incorporate wage and price controls; and c) simply to give up the struggle against inflation. The option necessarily depends on an assessment of the relative costs of the three alternatives, which can only be done

Editor's note: Translation revised by the author.

* This work is the result of research financed by the National Economic Research Program. The author would like to thank his colleagues at PUC/RJ and two anonymous readers for their helpful comments. The definition of the variables and the data used in this article are contained in the Appendix (not included in this version of the work) which can be consulted by application to the author or the publishers.

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on the basis of an estimate of the empirical relationship between inflation and the level of activity.

This paper is divided into two parts. The first, including Sections 2 and 4, undertakes a critical review of existing econometric studies, and a methodological critique of our previous work with André Lara Resende. The conclusion that none of these studies is satisfactory justifies the development of a new econometric estimate, in the three sections of the second part of the work. The major problem is dealt with in Section 5, where we have built a model of the inflationary process in an economy with unsynchronized indexation and of fixed periodicity for wages. Section 6 presents some estimates of the model and inflation forecasts up to 1984 under different scenarios. This way, we obtain a quantitative assessment of the influence of variations in the level of activity on inflation. Section 7, the last of this part, presents a theoretical justification for the hypothesis, used in the model of Section 5, that wage policy determines the level of inertial inflation in our economy. The article ends with a brief section of conclusions.

2 — Traditional estimates of the Phillips Curve for Brazil

Table 1 presents a summary of some existing econometric estimates for the relationship between inflation and the level of activity in Brazil. Despite some differences in the variables used and in the choice of the dependent variable, all estimates are based on a simple model of the accelerationist Phillips Curve, which assumes an inverse relationship between the acceleration of inflation and the rate of unemployment. In the absence of data on the rate of unemployment, the transposition of this model to the Brazilian situation demands the use of the output gap (calculated according to the percentage deviation of an index of the real aggregate product in relation to a tendency line) as a demand variable: the hypothesis of a stable relationship between the unemployment rate and the output gap, that is to say, Okun's Law [cf. Okun (1970)], is therefore implicit.

A superficial examination of the table seems to show that the application of this model to the Brazilian case is quite successful. The lagged inflation coefficient is very significantly different from 1, confirming the accelerationist nature o the Phillips Curve. On the other hand, all the estimate for the gap coefficient are significantly different from zero and have the

Table 1

Traditional Estimates of the Phillips Curve

Author	Dependent Variable	Period	Number of Observations	Method of Estimation	Estimated Equations (Statistics t in Brackets)	R^2	DW	Standard of Regression Error
1 — Lemgruber (1974)	$\dot{q}_a$	1953/73	21	MQO	$0{,}90 - 0{,}005H + 0{,}951\dot{q}\,(-1)$ $(1{,}64) \quad (8{,}74)$	0,72	1,65	0,115
2 — Contador (1977)	H	1974/75	29	MQO	$0{,}037 - 0{,}159\dot{q}_d + 0{,}330\dot{q}_e$ $(2{,}87) \; (-3{,}01)^* \; (5{,}01)$	0,50	0,94	0,037
3 — Lemgruber (1980)	h	1950/79	30	MQO	$0{,}183 - 0{,}198\,\Delta\dot{q}_a \; 0{,}747\,h(-1)$ $(-0{,}25) \; (1{,}06) \quad (8{,}89)$	0,70	1,48	0,032
4 — Lemgruber (1980)	$\dot{q}_a$	1950/79	30	MQO	$0{,}33 - 0{,}523h + 0{,}913\,\dot{q}_a\,(-1)$ $(1{,}18) \; (3{,}31) \quad (0{,}09)$	0,75	2,29	0,089
5 — Contador (1982)	$\dot{q}_i$	1950/79	30	MQO(?)	$0{,}121 - 0{,}699h + 0{,}827\dot{q}_i\,(-1)$ $(2{,}08) \; (-1{,}80) \; (5{,}90)$	0,57	2,30	—
6 — Author's Estimate	$\dot{q}_d$	1952/81	30	MQO	$0{,}103 - 1{,}025H + 1{,}005\dot{q}_d\,(-1)$ $(1{,}90) \; (2{,}44) \quad (8{,}74)^*$	0,74	1,90	0,140
7 — Author's Estimate	$\dot{q}_d$	1952/64	12	MQO	$-0{,}074 + 1{,}837H + 0{,}889\dot{q}_d\,(-1)$ $(-0{,}60) \quad (1{,}21) \quad (3{,}68)$	0,75	2,25	0,124
8 — Author's Estimate	$\dot{q}_d$	1965/81	10	MQO	$0{,}118 - 1{,}205H + 1{,}053\dot{q}_d\,(-1)$	0,77	1,48	0,148

Symbols used: $\dot{q}_a$ = rate of inflation, implicit GDP deflator;
H = output gap, measured by the GDP;
$\dot{q}_d$ = rate of inflation, based on the annual averages of the general WPI-IA;
$\dot{q}_e$ = expected rate of inflation, calculated by an auto-regressive model;
h = output gap in industry;
$\dot{q}_a$ = rate of inflation, based on averages of GPHA; and
$\dot{q}_i$ = rate of inflation, based on the annual averages of WPI-global supply-industry.

NOTE: When comparing these equations, one must bear in mind that the measures of the gap used by the various authors are slightly different:

a) Lembruber (1974): $H=0$ in 1961 and growth rate of potential product of 7% a year;
b) Contador (1977): $H=0$ in 1951 and 1974;
c) Lemgruber (1980): $H=0$ in 1958 and growth rate of potential product of 0.8% a year;
d) Contador (1982): potential output estimated by the exponential tangency of the effective output in 1910 or
e) author's estimate: $H=0$ in 1928 and growth rate of potential product of 7% a year.

correct sign (with the exception of estimate n.º 7, which will be discussed later on), confirming the existence of a significant negative relationship between the acceleration of inflation and the level of activity.

However, when we go on to a more detailed analysis of the econometric evidence, certain peculiarities must be noted. In the first place, we see that the gap coefficient appears to be much greater than would be expected from traditional estimates of the same empirical relationship in other contries. Gordon's study (1977) of the American economy in the 70s, for example, suggests a coefficient for the unemployment rate between 0.25 and 0.50, which means that one extra percentage point of unemployment produces a fall, over a year, of 1/4 to 1/2 percentage point in the rate of inflation. If we assume that, according to Okun's Law, a variation of one percentage point in the unemployment rate corresponds to a variation of 2.5 percentage points in the output gap, then the gap coefficient of the Phillips Curve should be, according to Gordon's evidence, from 0.1 to 0.2. This is much less than the estimates between 0.5 and 1.0 of Table 1.

As the gap coefficient increases, the impact of variations of the level of activity on the rate of inflation also increases. Indeed, a second peculiarity common to almost all the estimates of Table 1 is that substantial reductions in the rate of inflation may be obtained with relatively moderate output costs. Using, for example, equation 8 leads to the following projections of the rate of inflation (in terms of annual averages) for the years 1982/84, on the hypothesis that the output gap is kept throughout the whole period at the same level of 13% reached in 1981.

$\bar{q}_d$ (1980) = 109.2% (observed)

$\bar{q}_d$ (1981) = 113.0% (observed)

$\bar{q}_d$ (1982) = 97.1% (projected)

$\bar{q}_d$ (1983) = 83.8% (projected)

$\bar{q}_d$ (1984) = 72.5% (projected)

These projections indicate that, even if the GDP growth rate returns in 1982 to its tendential level of 7% a year, the demand shock engendered in 1981 will have been sufficient to reduce the 1984 rate of inflation to around 60% of the 1981 level. Undoubtedly there is a substantial contrast between the "deflationist optimism" suggested by the Brazilian estimates of the Phillips Curve and the opposite "deflationist pessimism" which seems to impregnate the American literature that is based on the same empirical model [see, for example, Tobin (1980)].

We should also consider certain elements that increase our doubts as to the forecasting reliability of the equations in Table 1. The standard regression errors are substantial (in the order of 10 percentage points for the rate of inflation) and the equations lose all their adherence in the recent episode of accelerated inflation (the forecasting error in equation 6 for the year 1980 is of the order of 40 percentage points). Beside this, the estimate of the gap coefficient is extremely sensitive to the definition of the sample. If we compare, for example, equations 7 and 8, corresponding to periods before and after 1964, the gap coefficient changes sign when moving from one period to another.

The observations plots given below permit a visual analysis of the instability of this coefficient. The axes are inflation acceleration and the gap; Figure 1 has observations for the period 1953/81, Figure 2 for the subperiod 1953/64. It is clear, from the second graph, that the positive sign of the gap coefficient in the pre-1964 period results from the inflationary spurts of 1959 and 1963 which cannot be explained by the model. It is also interesting to notice that, if the points corresponding to the PAEG (1965/67) and to the energy shock of 1980 are eliminated from Graph I, it becomes difficult to detect any systematic relationship between changes in the rate of inflation and the output gap.

3 — My model with Lara Resende

When we examine the observations in Figure 1 is clear that there are two important omissions in the traditional specifications of the Phillips Curve for Brazil. One results from not taking the effect of external inflationary shocks into account which is hard to justify after the oil crisis. They also disregard a key institutional determinant of the Brazilian economy since 1965: the compulsory indexing of wages established by the wage policy laws. Implicity, Equations in Table 1 imply the assumption that the market mechanism neutralizes the wage laws, which cannot be guaranteed a priori and should certainly be tested empirically.

The model of the inflationary process that developed with André Lara Resende sought precisely to eliminate these two defficiencies of conventional Phillips curve analysis. On the one hand, it allows for external shock by introducing the price of imported inputs as a cost element in the price equation. On the other hand, it assumes that the labour market is divided into two

Figure 1

CHANGE IN THE RATE OF INFLATION

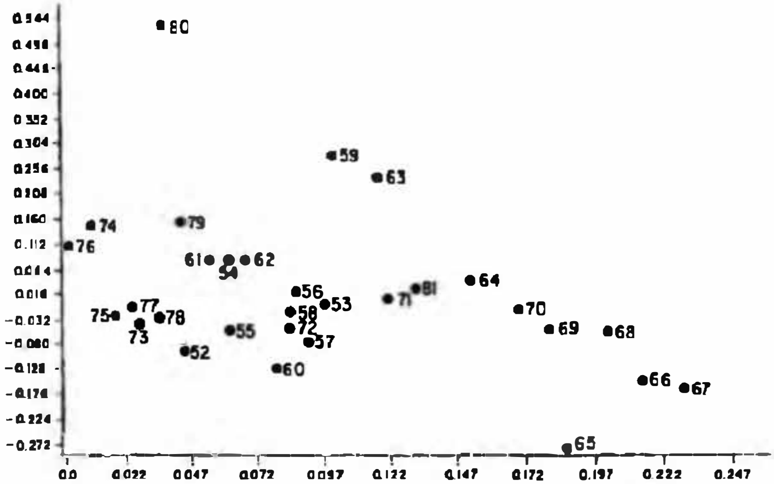
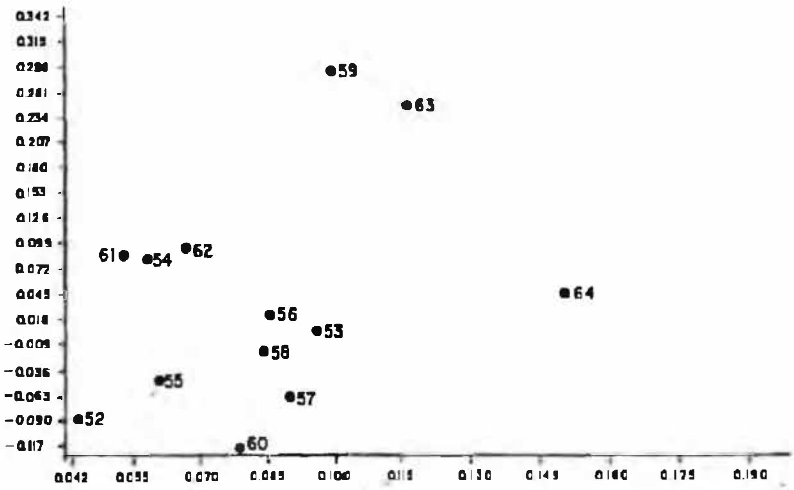


Figure 2

CHANGE IN THE RATE OF INFLATION



sectors: a market sector, in which the wage dynamics independent of the wage laws, and an institutional sector, in which wage behavior is defined by the wage laws. Since the relative size of the two sectors is not defined *a priori*, it is possible to test for the importance of wage policy as a determinant of the price dynamics.

One version of the model in structural form is the following:

$$\hat{w} = a\,\hat{w}_1 + (1-a)\,\hat{w}_2 \tag{1}$$

$$\hat{w}_1 = b - c\,H + \hat{q}_d(-1) \tag{2}$$

$$\hat{w}_2 = \hat{w}^* \tag{3}$$

$$\hat{q}_i = d_1\,\hat{q}_m + d_2\,(\hat{w} - g_p) + d_3\,\hat{q}_d\,(-1) \tag{4}$$

where: a, b, c, d_1, d_2 and d_3 are positive constants; $\hat{w}_1$ and $\hat{w}_2$ are the rates of change of the nominal wage in the market and institutional sectors; $\hat{w}$ is the average rate of change of the nominal wage; H is the aggregate output gap; $\hat{q}_d(-1)$ is the rate of inflation for annual averages of the WPI-IA, with one period lag; $\hat{w}^*$ is the wage readjustment defined by the wage policy; $\hat{q}$ is the rate of inflation for annual averages of the industry WPI-IA; $\hat{q}_m$ is the rate of change of the cruzeiro import price; and g_p is the growth rate of labour productivity.

The parameter a in equation (1) is the relative weight of the market sector in the labour force. If $a = 1$, the wage dynamics is determined by the conventional Phillips curve of equation (2); if $a = 0$, the wage dynamics determined by the wage policy in accordance with equation (3). Therefore, a measure of the relevance of the wage policy in determining the price dynamics of the economy can be obtained empirically by testing the hypothesis $a = 1$.

Equation (4) assumes a mark-up rule for fixing prices in the industrial sector, with $\hat{q}_m$ indicating the rate of change of the average cruzeiro price of imported inputs, $(\hat{w} - \hat{g}_p)$ indicating the rate of change of the labour cost per unit of product and $\hat{q}_d(-1)$ indicating the built-in element of inflationary inertia in the price of domestic inputs.

Substituting equations (1), (2) and (3) in the price equation (4), and defining an external shock variable $z = \hat{q}_m - q_d(-1)$, we get the following reduced form for the rate of inflation of industrial prices:

$$\begin{aligned}\hat{q}_i &= d_1 z + d_2 ab - d_2 acH + (d_1 + d_2 a + d_3)\,\hat{q}_d(-1) + \\ &\quad + d_2(1-a)\,(\hat{w}^* - g_p) - d_2 a\, g_p \\ &= f_1 z + f_0 - f_2\,H + f_3 q_d(-1) + f_4(\hat{w}^* - g_p) - f_5 g_p \end{aligned} \tag{5}$$

The estimation of this reduced form for the period 1960/78, using the rate of change of the minimum wage as a proxy for $\dot{w}^*$ and the growth rate of *per capita* output as a proxy for g_p, produced the results presented in Table 2. It should be noted that a dummy was introduced for 1963, as a result of the observation that, when the regression without a dummy is estimated (as in equation 5 of the table), the forecasting error in that year is systematically about three times larger than the standard regression error. Part of the explanation for this discrepancy may lie in the fact that the rate of change of the minimum wage is a particularly bad proxy for the rate of change of labour cost in 1963.

A surprising result that shows up in Table = : the coefficients f_2 and f_5 of the reduced form are not significantly different from zero and, besides, they have signs opposite to the expected ones. This means that these regressions do not reject the hypothesis $a = 0$, or that the nominal wage dynamics depends only on the wage policy. As the same results can be repeated by using the general WPI instead of the industry WPI (with some loss in the explanatory power of the regression), the conclusion is that there is no significant relationship between the rate of inflation and the level of activity in the Brazilian economy.

It is interesting to notice that, from the point of view of the model of this section, the traditional estimates of the Phillips Curve of the previous section are just the result of a statistical illusion. If the observations "contaminated" either by external shocks (1974, 1976, 1979 and 1980) or by imperfect indexing of the nominal wages (1965, 1966, 1967 and perhaps 1968) are eliminated from Figure 1, then the impression of a significant inverse relationship between the acceleration of inflation and the output gap disappears. The statistical illusion occurs because the inflationary external having occurred in periods of high activity level (small gap) and the deflationary wage shocks in periods of low activity level large gap).

4 — A problem of econometric identification?

Our work with André Lara Resende had the merit of questioning the traditional estimates of the Phillips Curve for Brazil. On the other hand, however, our extreme result surprised me. When the model of the previous section was conceived, my *a priori* expectation was that both, the wage and external shocks and the pressure of demand, would contribute to the explanation of the inflationary process, but that the inclusion of the shocks in the econometric analysis would substantially reduce the gap coefficient. The possibility of a zero coefficient, which cannot be reject from the evidence in Table 2 was not in our cogitations.

Table 2

Lara Resende/Lopes Model
(Dependent Variable: $\hat{q}_i$; Period 1960/78

	Independent Variables						
	Constant	s	$\hat{w}-g_p$	$\hat{q}_d(-1)$	g_B	H	Dummy[a]
Equation 1 $R^2 = 0.97$ SE = 0.04 DW = 2.30	0,0267 (0,45)	0,2770 (2,58)	0,6034 (2,68)	0,3513 (1,44)	0,1716 (0,33)	0,0736 (0,33)	0,2828 (5,32)
Equation 2 $R^2 = 0.97$ SE = 0.03 DW = 1.59	— —	0,3903 (4,99)	0,4219 (3,78)	0,5545 (5,35)	— —	— —	0,2927 (7,22)
Equation 3 $R^2 = 0.98$ SE = 0.03 DW = 2.08	0,023 (0,50)	0,388 (5,03)	0,465 (6,01)	0,454 (4,46)	0,073 (0,17)	0,036 (0,29)	0,254 (7,43)
Equation 4 $R^2 = 0.98$ SE = 0.05 DW = 1.95	—	0,431 (7,29)	0,426 (6,77)	0,552 (10,03)	— —	— —	0,290 (7,91)
Equation 5 $R^2 = 0.93$ SE = 0.07 DW = 1.91	0,143 (1,43)	0,433 (2,63)	0,569 (3,57)	0,224 (1,01)	—0,101 (—1,33)	0,121 (0,43)	— —

NOTES: Values in brackets are statistics t. Equations 1 and 2 were estimated by the method of instrumental variables, using as instruments the rate of change of the price in dollars of imports, $\hat{q}_d(-1)$, g_B, H, the constant, the dummy for 1963 and the growth rate of the quantum of imports. Equations 3, 4 and 5 were estimated by simple minimum squares. The external shock variable used in the regressions was defined as $s = (1+g_m)/(1+g_d(-1))$.

[a] Referring to the year 1963.

Let now examine this question carefully. It seems reasonable to assume that the wage policy determines the inertial inflation in our economy — what Eckstein (1981) called core inflation.[1] It is known, however, that firms can make the rate of change of their labour costs diverge from the wage policy standard: upwards, through promotions, extraordinary gratifications and other benefits; downwards; through labour turnover. It is clear that in certain areas of the labour market — corresponding to what Okun (1981) called "career labour markets" — there is almost no labour turnover, as it would break the implicit work

1 We get back to this issue on Section 7.

contract, characterizing the firm as a "bad employer" with negative repercussions on labour productivity and the average quality of the new workers that it might contract in the future. This would justify some downward inflexibility of the rate of change of labour costs in these sectors, relative to the inertial level defined by the wage policy.

It is not reasonable, however, to infer from this that the rate of inflation (even if measured in terms of industrial prices) must also be entirely insensitive to demand. There are sectors of the labour market (which Okun calls "casual labour markets") in which most of the jobs are of short duration and for which the notion of an implicit work contract has no relevance. This seems to be the case for important segments of agriculture, many services and even some industries (for example, civil construction and small industrial firms). In the American economy, approximately 10% of workers aged between 30 and 34 have jobs of less than one year's duration [cf. Hall (1980)]; for the Brazilian economy, this number is possibly many times greater. In these sectors, the wage policy cannot establish a very effective minimum for the rate of change of labour costs.

Besides we should also notice that the wage policy is irrelevant for self-employed workers or liberal professionals, and that it is common for firms to adopt the practice of giving discounts on the selling price when they are surprised by an abrupt demand cut and want to get rid of their surplus stocks rapidly. Even if the price index is calculated without taking these discounts into account, there may still be repercussions on the rate of inflation, if the phenomenon is occurring at the level of intermediary goods and the cost of inputs in other sectors is affected.

These reflections suggest that, despite the wage policy, there must be some connection between inflation and the level of activity, even if possibly tenuous. How then can we explain the results of Table 2? This is a complex question, for which we do not have a definitive answer. It is possible that the output gap variable is an inadequate measure of the level of activity. It is also possible that the model of equations (1) to (4) is incorrectly specified and that, as a result, a weak connection between inflation and the level of activity has not been captured. We must also consider the possibility of the mechanics of the wage readjustments precluding the econometric identification of the equation (5).

To understand this last point, let us consider an alternative version of the structural model of equations (1) to (4). In this model, the price equation is:

$$\hat{q} = j\,\hat{q}_m + (1 - j)\,\hat{q}_w \tag{6}$$

where $\hat{q}$ is the rate of change of the aggregate price index, $\hat{q}_m$ is, as before, the rate of change of the average price in cruzeiros of imports and $\hat{q}_w$ is the rate of change of the cost of labour. This last variable is determined by:

$$\hat{q}_w = k_0\, \hat{w}^* + (1 - k_0)\, \hat{w}^*(-1) + k_1 - k_2\, H \tag{7}$$

where the first two terms on the right-hand side of the equation are a consequence of the practice of unsynchronized wage readjustments with fixed periodicity (see Lopes and Bacha, 1983, for a detailed derivation of this relation) and the other two terms represent the impact of the aggregate demand on labour cost. The third equation of the model represents the wage indexation rule:

$$\hat{w}^* = (1 - \delta)\, q \tag{8}$$

which makes the rate of wage reajustment less than the current inflation rate if δ is positive. [2] This parameter represents the discretionary element of wage control, which seems to have been particularly important in the years 1965 to 1968.

Substituting (7) into (6) and using (8) to eliminate the term $\hat{w}^*(-1)$, we obtain:

$$\begin{aligned} \hat{q} &= j\, \hat{q}_m + (1 - j)\, k_0\, \hat{w}^* + (1 - j)\,(1 - k_0)\,(1 - \delta)\, p(-1) + \\ &\quad + (1 - j)\, k_1 - (1 - j)\,(1 - k_2)\, H \\ &= f_1'\, \hat{q}_m + f_4'\, \hat{w}^* + f_3'\, \hat{q}(-1) + f_0' - f_2'\, H \end{aligned} \tag{9}$$

[2] In our work with Edmar Bacha, equations (7) and (8) are collapsed into a single equation:

$$\hat{q}_w = \lambda \hat{p} + (1 - \lambda)\, \hat{p}(-1)$$

as we did not take into account the terms of demand of equation (7) and the discretionary policy parameter of equation (8). It would not, however, be correct to suppose that:

$$\hat{q}_w = \hat{w}^* = \lambda \hat{p} + (1 - \lambda)\, \hat{p}(-1)$$

because the variable $\hat{w}^*$ is not an adequate proxy for the rate of variation of the annual average labour cost $\hat{q}_w$. This seems to be confirmed by the following regression for the period 1960/73:

$$\underset{}{\hat{w}^*} = \underset{(5{,}13)}{0{,}88\, \hat{q}_c} + \underset{(0{,}13)}{0{,}02\, \hat{q}_c(-1)} \qquad R^2 = 0{,}76 \quad DW = 2{,}46$$

where q_c is the cost of living index (annual average).

which is not identical to equation (5), estimated in Table 2, only because we are ignoring here the productivity growth term g_y.

The identification problem becomes clear when we consider the model of two simultaneous equations, (8) and (9), with endogenous variables $\bar{q}$ and $\bar{w}^*$. It is obvious that this model does not allow us to identify equation (9); if we run a regression trying to estimate this equation, we are in fact estimating equation (8), which is the only one that is identified. In this regression, the coefficients f_0' and f_2' will tend to be non-significantly different from zero, but the coefficients f_1' and f_3' may be significant if the variables $\bar{q}_m$, $\bar{w}^*$ and $\bar{q}(-1)$ are strongly colinear, which is compatible with the results of Table 2.

Obviously, the judgement as to whether a particular equation is identified or not can only be made on the basis of the assumption that the structural model considered is in fact the correct one. As this is always impossible to establish *a priori,* we must take great care before rejecting a regression as a result of a non-identification argument: after all, any econometric model is only an approximation to reality.

It is interesting to notice that, from the point of view of our research objective, this identification problem (if it actually exists) is perfectly soluble. As our aim is to determine the magnitude of the partial derivate of the rate of inflation in relation to the gap, we only need an estimate of the reduced form of the model of equations (8) and (9), that is to say, an estimate of:

$$\bar{q} = f_1^r \, \bar{q}_m + f_3^r \, \bar{q}(-1) + f_0^r - f_2^r \, H \tag{10}$$

This, however, is not an easy equation to estimate, despite its apparent simplicity. The problem is that, in view of equation (8) all the coefficients f_r^r of the reduced form depend on the parameter of discretionary wage policy δ,[3] which probably took on different positive values in different years of the sample. This means that each term on the right-hand side of the equation would have to be split several times to capture these discretionary variations of the wage readjustment rule, destroying all the degrees of freedom of the regression.

The strategy that we will adopt here to avoid this problem of instability in the coefficients of equation (10) is to restrict our sample to the years after 1968, in which the wage policy

[3] The gap coefficient, for example, would be $f_2^r = f_2'/(U_4'(1-\delta))$.

produced readjustment rates that were approximately equal to the rate of inflation. For these years the parameter can be considered uniformly zero [cf. Simonsen (1974)].

Besides this, however, two other problems with the model of equations (6) and (8) must be fixed before trying to estimate it econometrically. One of them is in equation (8), which defines the rate of change of labour cost in terms of wage readjustments. This equation was derived in our work with Edmar Bacha on the basis of some simplifying hypotheses and mainly with a view to theoretical analysis. It is difficult to estimate the magnitude of the errors of approximation that result from using it in econometric analysis. The other problem has to do with the external shock variable in the price equation (6). There are two considerations to be made there: first, it is well known that the domestic price of petroleum derivatives behaved after 1974 quite differently from the international price of oil; second, the hypothesis of a unitary elasticity substitution is implicit in equation (6), while Brazilian and international evidence seems to indicate that the elasticity of substitution between domestic inputs and energy inputs is much less than 1, and probably not greater than 0.20. The next section develops a model of the inflation dynamics, in which all these problems are overcomes.

5 — A model of the inflationary process with unsynchronized wage indexation

Le us assume the following mark-up rule for the determination of industrial prices:

$$p_i = \{x_0 p_0 (-2) + x_m p_m (-2) + x_w c_w\} (1 + \alpha) \tag{11}$$

where p_i is the value of the industrial price index in a particular month, p_0 (—2) and P_m (—2) are the two month lagged values of the price of fuel-oil and the price in cruzeiros of non-petroleum imports, c_w is the unit labour cots which will be examined in detail below), the terms x_k, for $k = 0, m, w$ represent technical coefficients and α is the mark-up. The lag structure assumed here is arbitrary and presupposes — realistically in our opinion — that increases in labour cost affect prices more rapidly than increases in the costs of fuel-oil and other imported inputs. It would be better if the lag structure were determined empirically, but this was not attempted in the present survey.

Based on (11), and assuming a constant mark-up, we calculate the 12 months rate of inflation for industrial prices: [4]

$$(1 + \hat{p}_i) = \lambda_0 \{1 + \hat{p}_0(-2)\} + \lambda_m \{1 + \hat{p}_m(-2)\} + \\ + \lambda_w \{1 + \hat{c}_w\} \tag{12}$$

where: the rates of inflation in 12 months are calculated according to the formula $\hat{z} = (z/z(-12)) - 1$, so that, for example, $\hat{p}_i = (p_i/p_i(-12)) - 1$ e $\hat{p}_0(-2) = (p_0(-2)/p_0(-14)) - 1$; the elasticities are defined as $\lambda_0 = x_0\pi_0(-12)$, $\lambda_m = x_m\pi_m(12)$ and $\lambda_w = x_w\pi_w(-12)$, with the symbol π introduced to represent relative prices, $\pi_0 = p_0(-2)/p_i$, $\pi_m = p_m(-2)/p_i$ and $\pi_w = c_w/p_i$.

We will assume that the technical coefficients are determined by:

$$x_k = \bar{x}_k \{\pi_k(-12)\}^{-\sigma} \tag{13}$$

for $k = 0$, m or w, where the X_k are constants and σ may be interpreted as the elasticity of substitution between inputs. Notices that we have adopted the simplifying hypothesis that the technical coefficients respond to relative prices with a twelve month lag. The alternative hypothesis, that the technical coefficients respond to the current relative price, was also tested, but, as the value of the elasticity of substitution that we adopt is only 0.20, the two hypothesis produce practically identical estimates.

A small elasticity of substitution also allows us to assume that the sum of the elasticities λ_k is approximately unitary, [5] so we can rewrite (12) as:

$$\hat{p}_i = \lambda_0 \{\hat{p}_0(-2) - \hat{c}_w\} + \lambda_m \{\hat{p}_m(-2) - \hat{c}_w\} + \hat{c}_w$$

or, considering (13) and the definition of the λ_k:

$$\hat{p}_i = \bar{x}_0 [\pi_0(-12)]^{1-\sigma} [\hat{p}_0(-2) - c_w] + \\ + \bar{x}_m [\pi_m(-12)]^{1-\sigma} [\hat{p}_m(-2) - \hat{c}_w] + \hat{c}_w \tag{14}$$

[4] The reader will notice that our notation uses the symbol p to indicate the value of a price index in a particular month and the symbol q to indicate the average value of the same price index over 12 (or 6) months. In this way, $\hat{p}$ is a rate of inflation measured in terms of annual (or half-yearly) averages.

[5] It is easy to verify that the correct formula in this case is:

$$\sum_k \lambda_k \frac{x_k}{x_k(-12)} = 1$$

As the wage readjustments have a fixed periodicity of n periods, and are more or less uniformly distributed over time, we can assume that labour cost is defined by:

$$c_w = \frac{1}{n} \sum_{k=0}^{k=n} w(-k) \tag{15}$$

where $w(-k)$ is the nominal wage with a lag of k months. In this way, the cost of labour in a particular month is a simple average of the values of the nominal wage that were observed in the last n months.

If we consider that the wage policy determines a standard for the wage readjustments, but that the effective readjustments may differ from this standard as a result of the level of activity, it seems reasonable to construct the following approximate measure for the rate of change of labour cost.

$$\dot{c}_w = \dot{c}^*_w + a - bh \tag{16}$$

where c^*_w is a simple average for the last n months of a nominal wage index defined by the wage policy (w^*), that is to say: [6]

$$c^*_w = \frac{1}{n} \sum_{n=0}^{k=n-1} w^*(-k) \tag{17}$$

[6] It should be noted that the generic term n is being used here so that we can take into account the change in wage policy that occurred in 1979, when the value of n was reduced from 12 to 6. This change will be simulated here in the following way:

a) Before 1979:

$$c^*_w = \frac{1}{12} \sum_{k=0}^{k=11} w^*(-k)$$

b) For December 1979:

$$c^*_w = \frac{1}{12} \left\{ \sum_{k=0}^{k=5} w^*(-k) + (1.22) \sum_{k=6}^{k=11} w^*(k) \right\}$$

c) After 1979:

$$c^*_w = \frac{1}{6} \sum_{k=0}^{k=5} w^*(-k)$$

A measure or the potential inflationary impact of this change in periodicity can be obtained by comparing the values of $\dot{c}^*_w$ which would be produced

Note that $\bar{w}^*$ is the rate of readjustment determined by the wage policy and w^* is an index constructed on the basis of $\bar{w}^*$.

Substituting (16) into (14), we have:

$$\hat{p}_t = \bar{x}_0 f_0(-12)\ (\hat{p}_0(-2) - \hat{c}_w^*) + x_m f_m(-12)\cdot(\hat{p}_m(-2) - \hat{c}_w^*) + \\ + \hat{c}_w^* + a_t - b_t h \qquad (18)$$

$$= \bar{x}_0 z_0 + \bar{x}_m z_m + \hat{c}_w^* + a_t - b_t h$$

where $f_k(-12) = \P_k(-12)^{1-\sigma}$ and the variables $z_k = f_k(-12)$ $\hat{p}_k(-2) - \hat{c}_w^*$ indicate the inflationary shocks that may be attributed to the variations in the cruzeiro price of petroleum (for $k = 0$) and of other imported inputs (for $k = m$). Note that this equation defines a relation between the 12 months reate of inflation of the industrial prices and the rate of change measured by averages, annual or half-yearly, of the nominal wage defined by the wage policy, plus supply and demand shocks. In our opinion, this is the only specification of the inflation dynamics that is compatible with the mechanics of unsynchronized wage readjustments of fixed periodicity existing in our economy.

after 1979 according to the formula of item "a" with the "real'l values produced by the formulas of items "b" and "c". The result is the following:

(In %)

Year	(Old Formula)	("Real" Values)	Difference
1979	47	62	0,15
1980	80	88	0,08
1981	102	104	0,02

In periods of inflationary acceleration, it is natural that the rate of inflation measured by six months averages should be greater than the rate of inflation measured by twelve months averages, but the value of the difference in the year 1979 suggests the possibility of an additional inflationary impact as a result of the change in periodicity.

Note, however, that this is an inflationary impact on the cost of labour, which will overflow onto the prices if the mark-ups are fixed, but which can also be absorbed by mark-ups reductions without any inflationary effect on prices. Our experience in this research, in which we experimented with the two alternative measures of the rate of change of labour cost in 1979, favours the first hypothesis, but in our opinion this is an empirical question that remains open. In any case in the worst of all hypotheses, the inflationary impact that may be attributed to the change in the wage policy is of the order of 10 percentage points, which is quite small in view of the magnitude of the recent inflationary spurt.

6 — Empirical implementation of the model

Table 3 presents estimates of equation (18) and some possible variants. As can be seen, the statistical adjustment of the model to the data is quite satisfactory. It should also be observed that:

a) the coefficients of the supply shock variables, z_0 and z_m are significant in all the estimates;

b) changes in the value assumed for the elasticity of substitution (compare estimates 1 to 6) has little effect on the estimated values of the coefficients: the most important difference is in the coefficient of z_0, which doubles when the value of the elasticity is altered from 0.20 to 1;

c) the gap coefficient has a correct sign in all the estimates and is highly significant in the first four; in the last two, the significance is smaller, but the estimated value of the coefficient is practically the same as in estimate 1;

d) our equation does not give a very good explanation for the year 1973, when the forecast error is systematically two to three times greater than the average standard error for the other years, which may be the result of an error of measurement in the dependent variable (remember the controversy over the underestimation of inflation measures for 1973), justifying the use of the dummy in estimates 2, 3 and 4; and

e) the best estimates from the point of view of the standard regression error (SRE) are 3 and 4, in which the change in the gap appears significantly as a demand variable.

These estimates, 3 and 4, suggest an interesting notion about the way changes in the level of activity affect the rate of inflation. The implication is that a change in the gap, which is sustained over time, produces a much greater impact on the inflationary process in the first year after the change than in subsequent years. A justification for this effect may lie in the behaviour of mark-ups: if the rate of change of the average mark-up in industry is positively associated with the rate of change of the industrial product, it follows that it is also negatively associated with variations in the industrial output gap.[7] We can also recall

[7] Let y be the logarithm of the industrial product and $\bar{y}$ the logarithm of the potential product. We can define, without a large error of approximation, $h = \bar{y} - y$ which implies $\Delta h = \Delta\bar{y} - \Delta y$. In this last expression, the first term is (approximately) the growth rate of the potential product — which is assumed to be constant — and the secon the growth rate of the industrial product (QED).

Table 3

Estimate of Equation (18)
(Dependent Variable: p_t – WPI – Global Supply-Industry; Period: 1969/81 – Values for December of Each Year – 13 Observations) – Ordinary Least Squares

Estimate	Independent Variables							R^2	DW	SER
	z_0	z_m	Constant	h	Δh	$\dot{c}^e$	Dummy (1973)			
n.° 1 (σ = 0,20)	0,031 (2,36)	0,237 (4,73)	−0,013 (−0,87)	−0,316 (−2,20)	—	1*	—	0,90	1,72	0,035
n.° 2 (σ = 0,20)	0,021 (2,03)	0,274 (6,73)	− 0,02 (0,15)	− 0,308 (− 3,32)	—	1*	−0,081 (−2,43)	0,94	1,49	0,028
n.° 3 (σ = 0,20)	0,019 (2,03)	0,299 (8,12)	− 0,009 (−0,82)	− 0,298 (−2,78)	− 3,334 (−2,25)	1*	(−0,103) (3,58)	0,96	2,17	0,023
n.° 4 (σ = 0,20)	0,018 (1,97)	0,296 (8,26)	—	− 0,366 (−5,40)	− 0,299 (− 2,15)	1*	− 0,108 (−3.94)	0,96	2,07	0.026
n.° 5 (σ = 0,20)	0,035 (2,29)	0,246 (4,53)	− 0,005 (−0,26)	−0,268 (−1,58)	—	0,965 (16,60)	—	0,99	1,76	0,037
n.° 6 (σ = 1)	0.065 (3,68)	0,259 (4,38)	−0,023 (−1,44)	−0.200 (−1,39)	—	1*	—	0,90	1,86	0,035

*Exogenous restriction on the coefficient.

that, in the classic studies of Phillips (1958) and Lipsey (1960), both the unemployment rate and its rate of change used to explain the inflationary process, and that Lipsey has constructed a justification for this based on the spread of unemployment among different sectors of the economy.

Our prefered justification for this result is based on the notion that in the Brazilian economy the behaviour of the potential product of industry must be affected by the behaviour of the level of activity. Think of an economy with strong population pressure, in which there exist, side by side, a "modern" segment and a "traditional" segment of the labour market. What characterizes a worker as a member of the modern segment may be his greater level of qualification and the fact of living in an economically developed geographical area of the country. The supply of labour for industry depends on the labour force existing in this modern segment of the labour market. It is reasonable to suppose that in the short term there is an inelastic labour supply curve for industry, but that in the medium term the migration of workers from the traditional sector to the modern sector (or vice versa) in response to variations in industrial employment, makes the short-term supply curve shift horizontally. The result is that a change in the level of industrial activity has its effect on the excess demand (or supply) in the modern segment of the labour market and as a result, its effect on the nominal wage in this sector, partially neutralized in the medium term by these shifts in the labour supply. [8]

It is important to notice that the gap coefficient estimated in Table 3 cannot be compared with conventional estimates of this coefficient, like those in Table 1, as the variable $\hat{p}_t - c^*_t$

[8] To make our econometric result consistent with this argument, the measurement of the potential product has to be reformulated in the following way: let us consider our measure of gap $h = \bar{y} - y$, where y is the usual concept of potential output, constructed on the basis of the assumption of a constant growth rate, $\bar{g}$, and let us assume that the estimated price equation is:

$$\hat{p} = a(h + \Delta h) + b$$

in which, as in our estimate in Table 3, the coefficient of h and Δh is the same.

Defining a new measure of potential output as $y^* = 0.5\ (\bar{y} + y(1))$, we can rewrite this last equation as:

$$\hat{p} = -a(y^* - y) + b^*$$

where $b^* = h - a\bar{g}$. Note that the growth rate of y^* is approximately $g^* = 0.5\ (g + g(1))$, where $g(1)$ is the growth rate of the effective product in the previous year.

is different from the annual acceleration of inflation $\hat{p}_t - \hat{p}_t(-12)$. We can, however, calculate this annual acceleration if we use our equation to project $\hat{p}_t$ month by month, for the 12 following months. To do this, however, we need another equation that will allow us to determine the evolution over time of w^* and, consequently, of $\hat{c}^*_w$. The equation that we will use is for the rate of variation in 12 months of w^*, with a lead of two months reflecting the current systematic of using the National Consumer Price Index in fixing the wage readjustment:

$$\hat{w}^*(+2) = \underset{(5{,}65)}{0{,}32\,\hat{p}_t} + 0{,}68\,\hat{c}^*_w - \underset{(-1{,}87)}{0{,}21\,z_a} + \underset{(7{,}87)}{0{,}14}\ \text{dummy-1974}$$

$$R^2 = 0{,}95$$
$$DW = 2{,}08$$
$$SER = 0{,}016$$

This equation was estimated by ordinary minimum squares for the period 1969/80 with the restriction that the coefficients of $\hat{p}_t$ and $\hat{c}^*_w$ add up to 1. The variable z_a is a measure of agricultural shock, equal to the difference between the rate of agricultural production for internal supply and the trend value of this variable, of the order of 4% per year. A dummy variable was included to capture the 10% wage bonus of 1974.

The results of a simulation exercise for the period 1982/84 are presented in Table 4. In the three simulations presented, we assume that the internal price of fuel-oil changes proportionally to the WPI-IA-industry and that the rate of change of the internal price of imports maintains the same relation of proportionality observed in 1981 with the rate of wage readjustment (see the explanatory notes in the table). Of course, these hypotheses are arbitrary and have to be taken into account when reading the results of the simulations.

The three simulations differ in the following way:

a) reference simulation — industrial output gap increasing throughout 1982 to a figure of 27% (corresponding to an industrial output growth rate of 4%) and remaining constant thereafter; there is no agricultural shock ($z_a = 0$);

b) alternative 1 — constant gap at the level or 22.4% observed in 1981 (corresponding to an industrial output growth rate of 8.5% over the whole period); there is no agricultural shock; and

Table 4

Simulations of the 12 Months Rate of Inflation of Industrial Prices (WPI-IA)

Year/Month	Reference Simulation	Alternative 1	Alternative 2
1981/12	99,6	99,6	99,6
1982/06	93,2	96,3	92,5
1982/12	92,3	96,4	91,5
1983/06	96,7	100,2	96,1
1983/12	98,7	103,4	97,3
1984/06	99,6	105,6	97,5
1984/12	99,9	107,2	97,4

NOTES: a) Value observed in 1981/12.

b) In all the simulations it is assumed that: $\hat{p}_a = \hat{p}_i$ (internal price of fuel-oil grows together with the WPI-IA industry); and $\hat{p}_m = (1 + \hat{w}^*)\ (1.10) - 1$ (the relationship between the rates of change of the price in cruzeiros of imports and the wage readjustment index — that is to say, the National Consumer Price Index — observed in 1981 is maintained).

c) Reference simulation — assumes that the gap increases throughout 1982 to 27% by the end of the year (equivalent to an industrial output growth rate of 4% in 1982) and constant thereafter; agricultural shock $z_a = 0$.

d) Alternative 1 — assumes a constant gap at the level of 22.4% observed in 1981, and agricultural shock $z_a = 0$.

e) Alternative 2 — equal to the reference simulation as far as the evolution of the gap is concerned, but assumes a deflationary agricultural shock: $z_a = 5\%$ in 1982.

c) alternative 2 – the only difference in relation to the reference simulation is the hypothesis of a deflationary agricultural shock in 1982, with $z_a = 5\%$.

It is interesting to note that, despite the statistical significance of the gap and gap variation coefficient in our regression, the effect on the inflation dynamics of the strong demand shock suffered by the Brazilian economy in 1981 appears to be small. In the reference simulation, for example, though the demand shock is intensified throughout 1982 the rate of inflation falls only about 7 percentage points, and later increases once again, presenting at the end of 1984 practically the same value observed at the end of 1981. Comparing alternative 1 with the reference simulation, we see that if the gap does not increase by about 5 percentage points throughout 1982 (reimaining at 22.4%, instead of going up to 27%), the rate of inflation at the end of 1982 will be greater by about 4 percentage points and, at the end of 1984, by about 7 percentage points. The agricultural deflationary shock of alternative 2 produces even less significant differences in relation to the reference simulation.

It is clear that the simulation results depend on our hypotheses on the internal price of petroleum and other imports. Roughly speaking, however, they seem to corroborate the deflationist pessimism suggested by our previous work with Lara Resende. In spite of its high social cost, recession seems to be a rather inefficient mechanism for fighting inflation in Brazil.

7 — A note on the role of expectations

Our model of the inflationary process assumes that the wage policy determines the level of inertial inflation in the economy. In this section we are going to show that this hypothesis would be adequate even if the wage policy were applied only to a small segment of the labour market, with the salaries in the rest of the market being determined by the free play of market forces.

Let us consider a model similar to that in our work with Lara Resende, with the labour market divided into two sectors. In sector 1, the market operates freely, and the variation of the nominal wage is determined by a Phillips Curve:

$$\hat{w}_1 = -b\ h_1(-1) + \hat{p}^e \tag{21}$$

where h_1 is the rate of unemployment in the sector and $\hat{p}^e$ the expected rate of inflation. Sector 2 is the segment of the labour market in which the wage policy determines the inertial level of inflation, so that the variation of the nominal wage is:

$$\hat{w}_2 = -c\ h_2(-1) + \hat{w}_0 \tag{22}$$

where h_2 is the rate of unemployment in the sector and $\hat{w}_0$ the inertial level defined by the policy. As far as the present discussion is concerned, it is not necessary to specify this last variable in any detail: it is sufficient to assume that it is determined by a stable rule fully known to the public. We also assume that the rate of inflation is a weighted average of the rates of change of nominal wage in the two sectors:

$$\hat{p} = a\ \hat{w}_1 + (1 - a)\ \hat{w}_2 \tag{23}$$

Substituting (21) and (22) into (23), we obtain a reduce form of the model:

$$\hat{p} = -ab\ h_1(-1) + a\hat{p}^e - (1 - a)\ c\ h_2(-1) + \\ + (1 - a)\ \hat{w}_0 \tag{24}$$

which, with rational expectations, can be used to calculate the expected rate of inflation:

$$\hat{p}^e = \frac{-1}{(1 - a)} \{ab\ h_1(-1) + (1 - a)\ c\ h_2(-1)\} + \hat{w}_0 \quad (25)$$

Applying (25) into (24), we have:

$$\hat{p} = \frac{-ab}{(1 - a)}\ h_1(-1) - c\ h_2(-1) + \hat{w}_0 \quad (26)$$

which shows that the inertial level of inflation in the economy (that is to say, the rate of inflation that occurs when $h_1(-1) = h_2(-1) = 0$) is determined by the wage policy, despite there being a segment of the labour market in which wages are determined by the free play of market forces.

8 — Conclusion

This work has studied the empirical relationship between inflation and the level of activity in Brazil. The first three sections were devoted to a critical review of some econometric studies on the subject. Various estimates of the conventional Phillips Curve, made for Brazil by Lemgruber and Contador, are examined in Section 2. The output gap coefficients estimated by these authors are much greater than would be expected on the basis of similar estimates for the American economy. Besides this, the standard regression errors are substantial, the equations have very little predictive power in the recent episode of inflation acceleration and the estimate of the gap coefficient appears to be very sensitive to the definition of the sample period.

Section 3 reproduces the results of our previous work, which sought to correct two unjustifiable omissions in the conventional estimates of the Phillips Curve for Brazil: not taking into account the effect of external shocks and ignoring the compulsory indexing of wages imposed by the wage policy legislation. As we have seen, this work produced the unexpected result that there is no statistically significant relationship between inflation and the level of activity.

This surprising result is critically assessed in Section 4. Some theoretical arguments are presented, suggesting that this relationship can coexist with a Brazilian type wage policy. It follows a discussion of the possibility that the possibility that the equation estimated in our work with Lara Resende may not be econometrically identified. This type of argument, however, always involves the assumption that the structural model considered is in fact the correct one, which obviously cannot be established *a priori*. The section also discusses some ways of overcoming this potential non-identification problem.

At this point, the article should have demonstrated the need of a new model of the inflationary process in an economy with an unsynchronized indexation of wages. In this model, which is presented in Section 5, the 13 months rate of inflation is explained by a lagged inflation rate constructed on the basis of average values of the wage index. The model also takes external inflationary shocks into consideration and allows for the fact that, as a result of subsidies, the behaviour of the domestic price of petroleum derivatives does not necessarily follows the behaviour of the international price.

Estimates of the model with annual data for the period 1969/81 are reported in the following section, showing that both the industrial output gap and its variation contribute significantly to the econometric explanation of the rate of inflation of industrial a prices. However, a simulation exercise with this equation, also presented in Section 6, shows that *the impact of variations of the level of activity on the rate of inflation is very small.* Assuming, for example, that the output gap increases during 1982 to approximately 27% (corresponding to an industrial growth trade rate of 4% in the year) and then remains constant at this high level, we obtain a reduction in the rate of inflation of industrial prices from 99% in December 1981 to about 92% in December 1982 (in the most optimistic simulation), that is to say, a reduction of only 7 percentage points as a result of the greatest depression in the level of activity suffered by the Brazilian economy in recent history. In our simulations, the rate of inflation increases from 1983 onwards, but this result depends crucially on the hypothesis adopted for exchange policy.

The last section of the work takes a brief look at a theoretical assumption that is implicit in the model used in the two previous sections. The hypothesis is that the wage policy determines the inertial level of inflation, and this section demonstrates that it would be adequate even it the wage policy were applied only to a small segment of the labour market, with the wages in the rest of the market being determined by the free play of market forces. Basically, this theorem tells us that in an economy in which the labour market is segmented in two sectors, with the wages being determined in one of them by a wage policy rule and being determined in the other by a conventional Phillips Curve with rational expectations, then inertial inflation (which is the rate of inflation which occurs when the two markets are simultaneously in equilibrium (is determined by the wage policy rule, that is to say, in this economy the wage policy rule dominates the expectations.

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Impresso no Centro de Serviços Gráficos do IBGE — O. S.23 364

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