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"THE FORMATION OF INFLATION EXPECTATIONS IN BRAZIL: A STUDY OF THE FISHER EFFECT IN A SIGNAL EXTRACTION FRAMEWORK"

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Sumário

Este paper deriva um mecanismo de extração de sinal para examinar todas as implicações testáveis da equação de Fisher. O sinal é a taxa de juros nominal líquida de impostos, a qual, sob o modelo de Fisher, é igual à *soma* da expectativa de inflação com a taxa real de juros (que se supõe igual a uma constante ou a uma constante mais uma diferença de martingale). Todos os modelos lineares alternativos podem ser representados como ruído adicionado ao sinal. As literaturas internacional e brasileira são concisamente revistas e os testes empíricos reinterpretados como casos especiais do mecanismo de extração de sinal. O teste do modelo é feito com dados de taxa de juros de CDBs de grandes bancos no período 1973 / 1990. Pouco ruído é encontrado, ou seja, a equação de Fisher parece ser uma aproximação válida para o processo de formação das taxas de juros no Brasil. Este resultado implica na impossibilidade do governo tentar diminuir sistematicamente o custo do financiamento de seus déficits emitindo dívida não indexada. Dadas as amplas flutuações observadas nas taxas reais ex post, o sucesso da equação de Fisher revela a existência de substanciais erros de previsão da inflação, o que indica a necessidade de mais pesquisa na área de como agentes formam expectativas inflacionárias. Quando a inflação acelera rapidamente, como aconteceu nos últimos anos da década dos 80 no Brasil, os índices de preço captam tal aceleração como atraso. Isto gera um efeito Fisher até para títulos indexados. Faz-se também um teste para este efeito neste paper.

Abstract

This paper derives a signal extraction framework for examining all testable implications of the Fisher equation. The signal is the net of taxes nominal interest rate, which, under the Fisher model, equals inflation expectation plus the real rate (assumed to be constant or a constant plus a martingale difference). All alternative linear models can be represented as noise added to the signal. The international and Brazilian literature are briefly reviewed and the empirical tests reinterpreted in the signal extraction framework. The model is tested with Brazilian data for the period 1973 / 1990 using interest rate data on non-indexed certificates of deposit from a sample of major Brazilian banks. The framework detects little noise, i.e., the Fisher equation seems to reasonably fit the Brazilian evidence. This result carries the policy implication that the government cannot have the burden of financing its fiscal deficits ameliorated by issuing non-indexed debt in periods when inflation is escalating. Given the large fluctuations observed in ex post real rates in Brasil, the reasonable success of the Fisher equation also implies the existence of large inflation forecast errors, which suggest the need for further research on how agents form inflation expectations. When inflation escalates rapidly, as it did in the late '80s in Brazil, price indices lag behind true inflation. This generates a Fisher effect even for indexed securities. A test for this effect has been also included in the paper.

1. INTRODUCTION*

Irving Fisher suggested that nominal interest rates adjust one-for-one to changes in expected inflation.[1] Whether or not the Fisher effect holds is a topic of clear importance in finance. And it is even more important to macroeconomic policy in developing countries. If increases in inflation expectations do not get fully incorporated in nominal interest rates, governments may have an incentive to run debt-financed fiscal deficits. This is particularly relevant for highly inflationary economies like Brazil.

The Fisher effect literature is large. As frequently happens in these cases, different papers concentrate on different testable implications of the null model, usually generating mixed results. The signal extraction framework used in this paper provides a unifying basis for all previously suggested tests in the literature. Following the modern principles of macroeconomics, the models tested in this paper assume that expectations are rational.

The first empirical application studies the existence of the Fisher effect in Brazil using data on the private banks' certificates of deposit (CD) market from 1973 to 1990. The econometric framework used provides measures of the plausibility of the Fisher model. These measures are important because macroeconomic models are judged and,

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1 Fisher himself had a word of caution: "The money rate and the real rate are normally identical; that is, they will ... be the same when the purchasing power of the dollar in terms of the cost of living is constant or stable. When the cost of living is not stable, the rate of interest takes the appreciation and depreciation into account to some extent, but only slightly and, in general, indirectly. ... That is ... when prices are falling, the rate of interest tends to be low, but not so low as it should be to compensate for the fall "(Fisher, 1930).

Note also that under taxes, this effect should be greater than one-to-one. If the tax code is not indexed, a 1% increase in inflation will require a $(1 + t)\%$ increase in nominal rates (t being the relevant tax rate) in order to keep the real return unchanged.

For a review of the large literature on the Fisher effect, see, for example, Summers (1983). The Brazilian and international evidence is reviewed in Rocha (1988).

more importantly, used in economic policy on the basis of their plausibility, not on the basis of the correctness of their specification.

The results indicate that the Fisher model satisfactorily fits the Brazilian data. This conclusion carries the policy implication that the government cannot have the burden of financing its fiscal deficits ameliorated by issuing non-indexed debt in periods when inflation is escalating.

The Brazilian experience can also be used to test an important corollary of the Fisher relation. Given the methodology of price index calculation in Brazil, the inflation index for a given month is actually a proxy for the previous month's inflation (this point is explained in Section 4). When inflation escalates rapidly as it did in Brazil in the latter part of the '80s, the inflation index, which guides backward financial indexation, underestimates true inflation. With rational investors aware of this statistical problem, the *promised* real rates of indexed securities[2] should rise to account for this fact, although the actual ex ante real rate may not vary as much. This is because the measured inflation, which is lagged by one month, underestimates actual inflation when inflation is accelerating. Tests for this Fisher effect on indexed securities are also undertaken.

After this brief introduction, Section 2 outlines the basic features of the signal extraction approach used in the empirical part of the paper. All previous tests of the Fisher equation are subsumed under the signal extraction framework considered here. Appendix 1 reviews the Fisher effect literature and shows that all previously suggested tests of the Fisher relation are special cases of the one considered in this paper. Section 3 presents the estimation using Brazilian data and results. Section 4 analyzes the existence of the Fisher effect for indexed securities. Section 5 concludes.

2 In Brazil, when one buys a CD indexed to inflation, the nominal yield is given by the growth in the official index *plus* the *promised* real rate. Since actual inflation may differ from the growth rate in the index, the promised real rate may also differ from the ex post real rate.

2. THE SIGNAL EXTRACTION FRAMEWORK

In this section, I briefly outline the econometric framework used in this paper, as applied to the Fisher effect. The general framework was developed in Durlauf and Hall (1989a) — hereafter DH — where proofs of the theorems used here can be found.

The null model is the Fisher equation.

$$H_0: R_t = \rho + \pi^e_{t+1} \tag{1}$$

where R_t = nominal interest rate net of taxes from t to t+1 (%);

ρ = expected real interest rate net of taxes (assumed constant) (%);

π^e_{t+1} = expected inflation rate from t to t+1 (%).

Equation (1) may be thought of as generated by a model with risk neutral agents and one asset which pays a constant (expected) return, ρ_t.[3] Due to non-linearities intrinsic to the calculations of the real rate and the use of the (linear) expectation operator, the Fisher effect (equation (1)) under uncertainty holds only as an approximation.[4]

In the signal extraction framework, the signal is defined to be $R_t = \rho + \pi^e_{t+1}$. All alternative linear models may be expressed as a noise term (S_t) added to the signal (equation (2)).

$$H_1: R_t = \rho + \pi^e_{t+1} + S_t \tag{2}$$

where S_t = model noise.

By rational expectations, the current inflation rate, π_{t+1} (the inflation that occurs during the maturity of the security that pays R_t) is given by equation (3).

3 $\rho_t = \rho$ or $\rho_t = \rho + \varepsilon_t$, ε_t being a martingale difference. The two alternatives are observationally equivalent.

4 For the Brazilian case, where the cross effect is relevant, equation (1) should read:

$$R_t = \rho + \pi^e_{t+1} + \rho\, \pi^e_{t+1} \qquad (1').$$

Under uncertainty, even equation (1') holds only as an approximation.

$$\pi_{t+1} = \pi^{e}_{t+1} + \nu_t \tag{3}$$

where π_{t+1} = inflation rate from t to t+1 (%);

ν_t = forecast error.

By subtracting (3) from (2), (4) is obtained.

$$R_t - \pi_{t+1} = \rho + S_t - \nu_t \tag{4}$$

Equation (4) has two unobservable variables (besides the constant, ρ), S_t and ν_t. Rational expectations imply that is white noise[5], i. e., no systematic errors are made when forecasting inflation. To be sure, ν_t is the forecast error, which is part of this model of rational expectations and the Fisher effect. S_t, on the other hand, represents everything that the null model does **not** account for. S_t could be, for example, the effect of expected monetary policy on the ex ante real interest rate.

The DH framework uses all available information (past and future) to separate the two unobservable variables, S_t and ν_t. Given the characteristics of this model, corollary 1.1 in DH[6] implies that a regression of $[R_t - (\pi_{t+1} + \rho)]$ on a set of variables known at time t that includes lagged inflation and current and lagged nominal interest rates captures **all** testable implications of the model of rational expectations and the Fisher effect. Intuitively, future information cannot be used to separate S_t from ν_t because, by rational expectations, ν_t is orthogonal to everything known at t. Therefore, a regression of $[R_t - \pi_{t+1}]$ on a constant and variables known at time t can only detect noise. Under the null hypothesis, no variable should show up significant in such a regression. Future variables (indexed $t+1$, $t+2$, ...), however, are not necessarily

5 I use the term white noise because it is better known. In general, ν_t could be a martingale difference. The same caveat applies for the use of the term random walk instead of martingale in this paper.

6 Corollary 1.1 (Durlauf and Hall, 1989a, p.10/1) states the equivalence of predictor and smoother tests of model noise when the information innovations are modelled as a subset of available information to the econometrician.

orthogonal to ν_t or S_t. Therefore, those variables *cannot* be used to separate the unobservables.

Thus, the DH framework uses all information available to the econometrician (past and future) to estimate the tightest lower bound for the model noise. This estimate gives an assessment of the plausibility of the original model. If much noise is detected, this indicates the existence of other factors not accounted for by the null model. Besides the graphical presentation, two different metrics, presented in Section 3, are used to assess how reasonable the model is. In other words, if the null hypothesis is a good approximation to reality, then information available at time t should explain very little of the variation of $[R_t - (\pi_{t+1} + \rho)]$.

3. ESTIMATION AND RESULTS FOR BRAZIL

a. Restatement of the DH Test for High Inflationary Economies

As mentioned before, when inflation rate is high the cross term in equation (1') cannot be neglected. For this case, equation (4) should read [6]

$$\ln\left(\frac{1 + R_t}{1 + \pi_{t+1}}\right) = \ln(1 + \rho) + S_t - \nu_t \qquad (5).$$

Therefore, the DH regressions have the natural logarithm of (1 + ex post real interest rate) as the dependent variable, and different sets of explanatory variables ($L_x(t)$). Unlike the regressions in levels, these regressions in logs are invariant to the unit in which the rates are expressed (% per year, % per month, etc.).[7]

All regressions use OLS with the standard errors corrected by White's (1986) heteroskedasticity-autocorrelation consistent covariance matrix estimator. The need for this correction is spelled out next, when the estimations details are explained. Once an estimate of the lower bound of the model noise is obtained, two measures, in addition to the graphical presentation, provide assessments of the model plausibility. The first one is the ratio $\sigma^2_S / \sigma^2_{R-\pi}$, where $\sigma^2_{R-\pi}$ is the sample variance of the series ln $\left(\frac{1 + R_t}{1 + \pi_{t+1}}\right)$. Noise and expectation errors are the two uncorrelated components of the ex post real rate variance; therefore, this ratio is a measure of how much of the movement in the ex post *real* interest rate is due to model noise, i.e., to causes extraneous to the model. This ratio, however, does not answer the question of what

6 Note that for small values of R_t, π_{t+1} and ρ, equation (5) is well approximated by equation (4). The forecast error in equation (5), ν_t, is approximately the % error in the price level forecast, i.e.:

$$\nu_t = \ln\left(\frac{1 + \pi_{t+1}}{1 + \pi^e_{t+1}}\right).$$

7 For this invariance proposition to hold, the regressors must also be of the form ln(1 + rate). The invariance does not apply to the constant coefficient.

fraction of the movements in the *nominal* interest rate is explained by noise. For this purpose the ratio σ_S^2/σ_R^2 is computed, where σ_R^2 is the sample variance of $\ln(1+R_t)$. This measure is not bounded between 0 and 1, because noise and expected inflation may be correlated. If the two are positively (negatively) correlated, the ratio above will underestimate (overestimate) the contribution of model noise.[8] Despite this unavoidable flaw, this normalization of the noise variance provides a good assessment of the plausibility of the model. In the context of the dividend stock price model, Durlauf and Hall's (1989b) analogous normalizations of noise variance are all near or above 100%, showing the complete failure of that model.

b. The Data

Gathering data on interest rates in Brazil is not an easy task. Three main reasons account for this difficulty. First, few organizations do a consistent job of compiling interest rates. The Brazilian Central Bank (BACEN) gathers these data for auditing purposes, but does not publish them in its monthly bulletin. Second, Brazilian economic history is full of periods in which interest rates had a cap, or other controls were imposed on banks' activities. Banks then circumvented the legal restrictions by resorting to non-standard ways of charging customers more for borrowing. This makes standardly computed interest rates a poor indicator of the cost of borrowing. This should not greatly affect this work, because I use rates offered on banks liabilities (CDs), not rates charged on assets (bank loans). Third, taxation of interest income varied a lot (as much as 7 times in just one year).

The data set used in this paper is composed of deposit rates from a sample of Brazilian banks. These rates are an average of the rates quoted the first week of each month by the banks on their certificates of deposit (CDB - certificado de depósito

8 Suppose the economy follows some kind of Tobin-Mundell effect: ex ante real rates fall when inflation rises and vice-versa. This would be captured by the model noise, which under the above assumption is negatively correlated with inflation expectation. Suppose further that this Tobin-Mundell effect is very powerful in determining the *nominal* (not only the *real*) interest rate. It is then conceivable that the ratio σ_S^2/σ_R^2 could exceed 1.

bancário).[9] To address the tax problem, all the tests will use rates net of withholding taxes. Given the Brazilian tax system, this is a good proxy for the net of tax rates relevant to investors.[10] These interest rates were published in *Taxa de Juros no Brasil* (1990) — Interest Rates in Brazil.

Figure 1 shows monthly inflation rates in Brazil measured by the general price index (IGP-FGV), the wholesale price index (IPA-FGV), and the consumer price index (IPC-FGV) – all three price indices compiled by the statistical bureau of the Fundação Getúlio Vargas. Inflation has a clear upward trend after 1979, accelerating towards hyperinflation in the later years of the sample.

Some periods were excluded from the sample for reasons I now explain. Figure 2 shows the evolution of the nominal and real rates (both net of withholding taxes), as well as inflation. Paradoxically, extended periods show negative real rates. Unfortunately, despite much effort, I was unable to find a measure of the volume of trading in this market. For one of the periods that display negative real rates, 1979 to 1980, the explanation is that controls were imposed on the banks' assets. The banks could not lend unless they borrowed from abroad. Therefore, banks were unwilling to pay competitive rates. Chow tests, not reported here, confirmed the structural difference of this period. For this reason, I excluded those years from the regressions.

On February 28, 1986, the Brazilian government launched the Cruzado plan, starting a series of five attempts to reduce and stabilize the inflation rate. As Figure 1 shows, the first three attempts failed, with inflation resuming its explosive path after a briefer recess each time. However, when stabilization plans are launched, their rules are designed assuming success (and they are not at all robust to failure). Assuming that inflation would fall from 14 percent a month to around zero (as happened in the first

9 For the August/82 to January/84 period, the CDBs were indexed to inflation by law. During this period, the series of nominal rates on non-indexed securities is composed of rates on exchange-bills (LC - letra de câmbio), which are fairly good substitutes for the CDBs.

10 Evidence for this claim is provided in Rocha (1988, Table 2.1, pp. 17). Until July 23, 1974 the Brazilian Treasury-Bills (LTNs) were not taxed. For a few months after taxation began, both taxable LTNs (issued after July 23) and non-taxable LTNs (issued before July 23) were traded in the market. The difference on the yield of those securities remained very close to the tax rate (30%).

month of the Cruzado plan), all credit contracts denominated in Cr$ (Cruzeiros) would have originated a gigantic wealth transfer towards creditors. This is because, when inflation disappeared, high *nominal* rates would have been transformed into unbearably high *real* rates. To prevent this problem, the government carried out a monetary reform (the Cruzado - Cz$ - was created) and announced a table of daily conversion factors from Cr$ to Cz$.[11] That table assumed a given inflation expectation in the old money - Cr$. Thus, the Cruzeiro depreciated everyday compared to the Cruzado.

11 Similar schemes have been used in the past. The first such scheme that I am aware of was implemented during the French Revolution (Velde and Sargent, 1990).

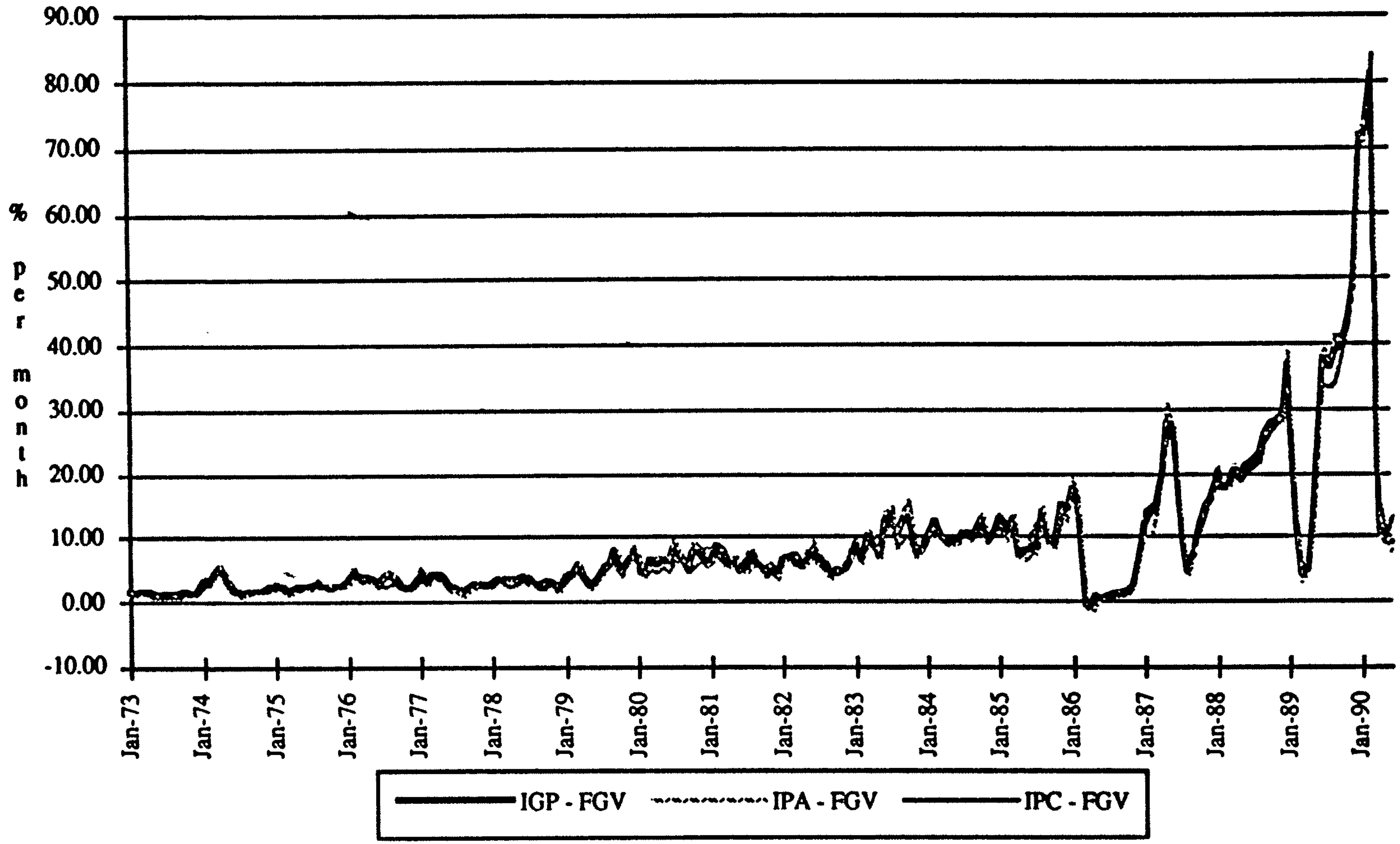

Figure 1: Brazil: Monthly Inflation Rates Measured by Different Price Indices

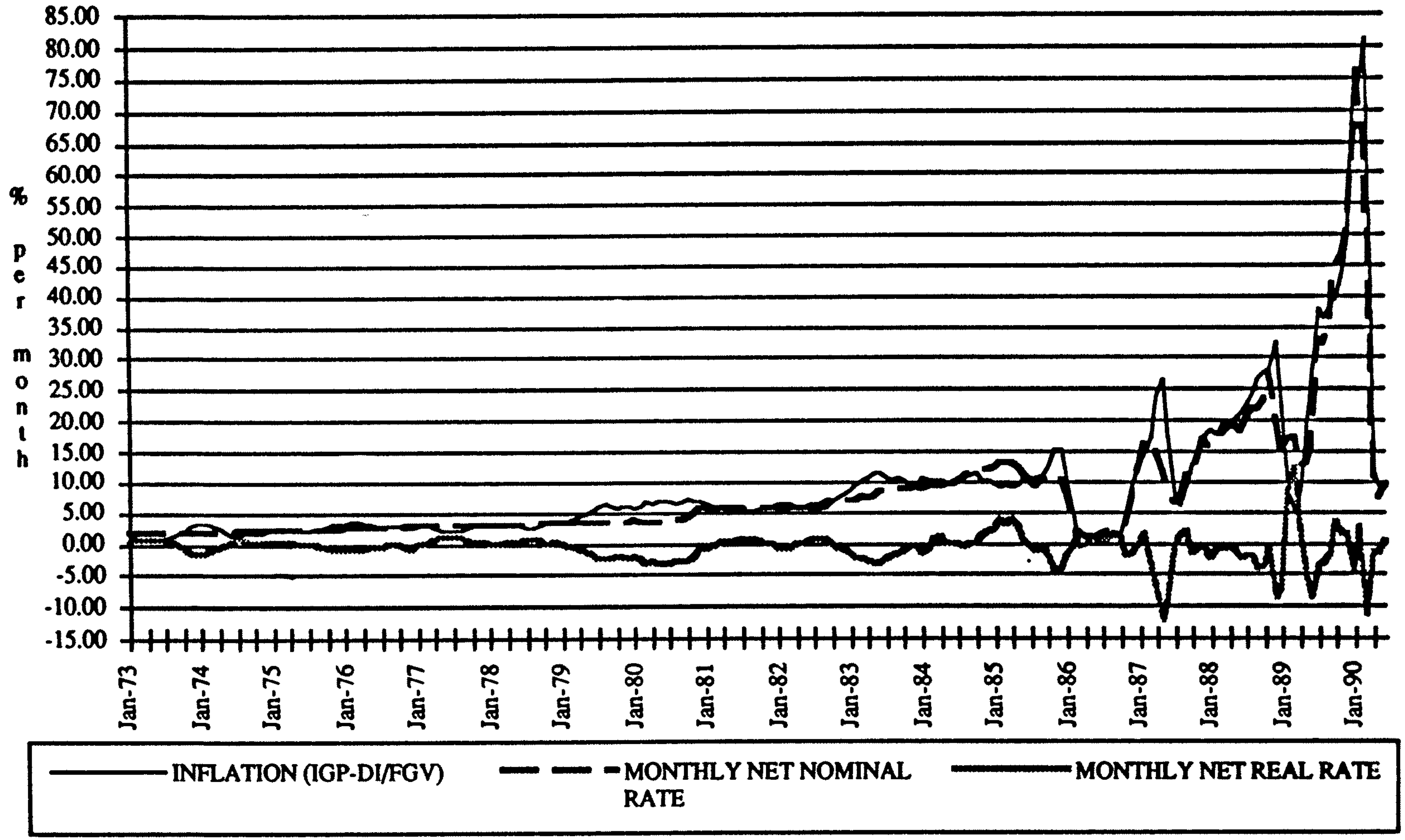

Figure 2: Non-Indexed CDBs: Net Nominal and Real Rates and Inflation During Maturity

Since inflation was typically escalating before the launching of the plans (after all, that was the main reason why they were adopted), the inflation expectation just before the plans tended to be higher than in the days or months before. Therefore, the table of conversion factors from Cr$ to Cz$ tended to lower ex post real interest rates, by stipulating too high a figure for what should have been the expected inflation implicit in the financial contracts when they were written. It could be argued that had the government done anything, the same effect or even worse would have come about, for inflation was rising. Nevertheless, to guard against the criticism that these artificial conversion mechanisms invalidate the conclusions, separate regressions were carried out for the period pre and post the Cruzado plan.

One important question related to the Fisher effect topic is the link between the Brazilian and world financial markets. If real interest parity held, the success of the Fisher model in explaining the Brazilian data would mean that the Fisher relation also held for foreign markets. For real interest parity to hold, two conditions are required. (See Frankel, 1991): i) uncovered interest parity must hold, and ii) the expected *real* depreciation must be zero. These conditions seem very strong for Brazil, where investments abroad were basically forbidden. Furthermore, the real exchange rate varied considerably in recent history. Nevertheless, it is probably warranted to assume that some link between the Brazilian and world markets existed through the US$ black market, or through other means (underinvoicing of Brazilian exports or overinvoicing of Brazilian imports, for example). Figure 3 shows the ex post real rates of the following thought experiment: invest in Brazilian CDs or buy US$ in the black market, invest in US CDs (6 months maturity), and convert back into Brazilian currency at the term date at the black market exchange rate. One must bear in mind that the latter investment option carried significant transaction costs not accounted for in these calculations, not to mention that it was illegal. Nevertheless, Figure 3 shows that for many periods the Brazilian CD performed better than its american counterpart, especially for the period before 1979, when non-indexed securities certainly constituted the bulk of the private CD market in Brazil.

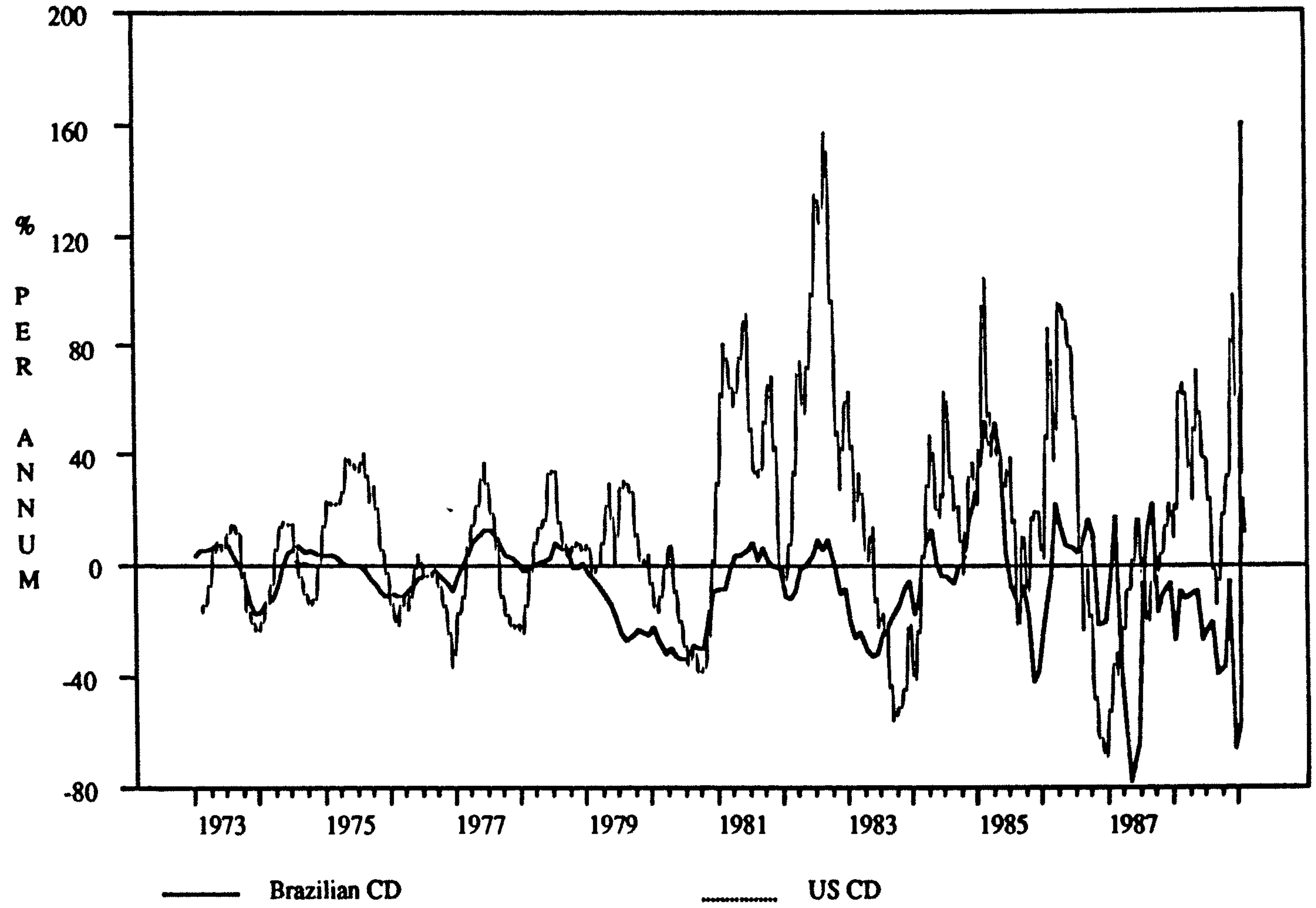

Figure 3: Ex Post Real Returns: US and Brazilian CDs

The next section also contains an explanation of how the link between the Brazilian and US capital markets is treated in this paper.

c. Estimation

The DH orthogonality test is carried out by using OLS in equation (6).

$$\ln\left(\frac{1+R_t}{1+\pi_{t+1}}\right) = c + \Gamma(L)\,\pi_t + \Psi(L)\,R_t + \varepsilon_t \tag{6}$$

where $\Gamma(L)$ and $\Psi(L)$ are lag polynomial operators $(\Gamma_0 + \Gamma_1 L + \Gamma_2 L^2 + \ldots)$. Therefore, the information set, $L_x(t)$, is composed of lagged inflation rates and current and lagged net nominal interest rates. Estimation is conducted first for lagged inflation rates only, then for interest rates only, and finally for the two sets combined.

Different measures of money and government debt (in real terms) are added to the explanatory variables. The alternative hypotheses contemplated by including monetary aggregates as regressors are the existence of a short term liquidity effect (the downwards short-run impact of money growth on interest rates) and, in the long run, the existence of the Tobin effect (Tobin, 1965). Previously suggested tests (see Appendix I for a quick review) have used the output gap as an explanatory variable. The most commonly invoked justification for this is the existence of a short run Phillips curve. In a boom, actual inflation systematically exceeds its expected value, depressing ex post real rates. GDP is not available on a monthly basis in Brazil. To proxy for the GDP gap, a measure of the industrial product gap was constructed by extracting its stochastic and deterministic trends, as well as its seasonal components. This two-step procedure is efficient and should not generate inconsistent standard errors, as pointed out in Pagan (1984, p. 233). Nevertheless, alternative results were generated by a one-step procedure which consisted of a regression of the ex post real rate on a constant, first-differenced industrial product and seasonal dummies.

The link between Brazilian and world financial markets was mentioned in the previous section. Since the purpose of this paper is not to test interest parity conditions in Brazil, the empirical approach will be expedient. The ex post real rate in Brazil is regressed against a constant, the nominal rate for 6 month CDs in the US, and the spread between the black market and official exchange rate (and seasonal dummies). The idea is that the spread in the black market exhibits some sort of mean reversion: High spreads forecast lower gains in investments abroad, and vice-versa.

The first technical point is the choice of the number of lags to include in the regression. Suppose the true model is composed of many lags. By truncating the lag length, one would erroneously reject the model. To prevent this problem, three criteria for choosing the optimal lag length are reported: Akaike, Schwarz (SC), and BIC. These criteria measure the trade-off between the decrease in the residuals' variance and the increase in the number of lags. The optimal distributed lag length is chosen by minimizing the functions below with respect to the lag length d.

$$\text{AKAIKE(d)} = \ln\left(T^{-1}\sum_{t=1}^{T}\hat{e}_t^{\,2}\right) + \frac{2\,d}{T} \tag{7}$$

$$\text{SC(d)} = \ln\left(T^{-1}\sum_{t=1}^{T}\hat{e}_t^{\,2}\right) + \frac{\ln(T)\,d}{T} \tag{8}$$

$$\text{BIC(d)} = \ln\left(T^{-1}\sum_{t=1}^{T}\hat{e}_t^{\,2}\right) + \frac{2\,\ln(T)\,d}{T} \tag{9}$$

where $\hat{e}_t$ is a sample projection residual. The Akaike criterion is known to overestimate the lag length (Judge et al, 1985, p.245). See Hannan and Deistler (1988) for an analysis of the BIC criterion.

The second technical point regards the consistency of the standard errors. The CDBs whose rates are used here had varying maturities. With rising inflation, contract lengths decrease. The longest maturity observed was 6 months. It is well known that a 6 step ahead forecast generates a MA(5) forecast error (Hansen and Hodrick, 1980). To obtain consistent estimates of the standard errors, White's (1986) heteroskedasticity-autocorrelation consistent covariance matrix estimator is employed. This non-parametric correction is the best possible correction in the absence of further information about the stochastic structure of the expected inflation process.

d. Results

Table 1 reports the results for the information sets composed of lagged inflation, current and lagged interest rates, and a constant. Each column reports the result of the regression of the ex post real interest rate on a constant and n lags of inflation, n being the number of lags. The SC and BIC criteria chose $n = 1$. The Akaike criterion chose $n = 23$ out of a possible 24. Given the Akaike's tendency to overestimate the lag length, the conclusions are based on the other two criteria. The fact that the lag length chosen by the SC and BIC criteria is small is a reassurance that adequate lag lengths are being considered here. The estimation is conducted for the whole sample (1973:1 to 1990:6), excluding the period 1979:1 to 1981:4 because of the controls imposed on banks at that time.

The first row is $\sigma_S^2 / \sigma_{R-\pi}^2$, i.e., a measure of how much movement of the ex post real interest rate is captured by the tightest lower bound estimate of noise. Noise and expectation error are the two orthogonal components of the ex post real rate movements. Therefore, the first row indicates that more than 80 percent of the movements in ex post real interest rates may be attributed to forecast errors in inflation. Therefore, despite the wild swings of the *ex post* real interest rates shown in Figure 2, the null model — equation (1) — seems to reasonably fit the empirical evidence.

The second row of Table 1 is σ_S^2 / σ_R^2, i.e., a measure of how much movement of the net nominal interest rate is explained by model noise. It is around 1 percent. Therefore, the null model does very well in explaining the Brazilian data, as far as lagged inflation, and current and lagged interest rates capture agents' inflation forecasts.

TABLE 1

ORTHOGONALITY MEASURES AND TESTS BASED ON CURRENT AND LAGGED INTEREST RATES, AND LAGGED INFLATION

	NUMBER OF LAGS INCLUDED					
	1	2	3	4	5	6
NOISE BOUNDS	BIC *					
	SC *					
% of variance of ex post real rate	8.5	13.7	17.2	17.6	17.7	17.9
% of variance of nominal rate	0.7	1.1	1.4	1.4	1.4	1.5
NOISE TESTS						
χ^2 statistic for all Inflation lags = 0	0.6	1.8	9.5	10.9	10.9	11.1
Degrees of freedom	1	2	3	4	5	6
Significance Level (%)	45.4	40.6	2.3	2.8	5.3	8.5
χ^2 statistic for all Nominal Rate lags = 0	1.4	1.9	6.1	6.3	6.9	6.6
Degrees of freedom	2	3	4	5	6	7
Significance Level (%)	48.7	58.7	18.9	27.5	32.7	46.7
χ^2 statistic for all Nominal Rate and Inflation lags = 0	9.0	9.9	37.8	39.4	44.4	51.0
Degrees of freedom	3	5	7	9	11	13
Significance Level (%)	2.9	7.8	0.0	0.0	0.0	0.0

Sample: 1973:1 to 1978:12 and 1981:4 to 1990:6 (the starting date varies with the number of lags).
Method of estimation: OLS with White's (1986) correction for the standard errors' estimates.
* BIC and SC criteria chose LAGS = 1. Akaike criterion chose LAGS = 23 out of a possible 24.

The noise tests in the last 3 rows indicate the presence of noise at the 1 percent significance level when three or more lags of inflation and nominal net interest rates are included in the regressions. The noise tests for inflation and nominal net interest rate separately indicate that inflation is more efficient in capturing noise. These results were corroborated by comparing the results of regressions (not reported here) of the ex post real rate on lagged inflation to those of regressions of the ex post real rate on current and lagged nominal net interest rate. The noise tests also indicate the need to include at least three lags to adequately capture noise. I performed regressions with only the constant and either the third or fourth lags, to investigate whether these lags had any special explanatory power. This is **not** the case. The increase in the χ^2 statistic is probably due to the fact that more regressors provide a better fit, especially in the latter part of the sample, when the ex post real interest rate swings wildly (see Figure 2).

Figure 4 shows the net nominal rate (% per month), the ex post net real rate (% per month), and the noise estimate (% per month) based on the projection on one to five times lagged inflation, zero to five times lagged interest rates, and a constant. The period 1979:1 to 1981:3 was excluded from the analysis due to government controls on banks. The overall conclusion is that elements not accounted for in the Fisher relation (noise) have negligible explanatory power on both nominal and ex ante real interest rates for the sample period studied.

When working on financial noise, one has to guard against spurious inferences and non-standard asymptotics caused by the presence of unit roots in the series. The low R^2s observed in the regressions do not suggest the presence of the spurious inference problem that arises when independent random walks are regressed against one another. The non-standard asymptotics problem occurs when the independent variables are integrated and the dependent variable is stationary. This problem generates consistent estimates with non-standard asymptotic distributions. In the present setting, this is a more likely event. To guard against it, the hypothesis tests were computed for all but the last term in each distributed lag. The asymptotic distribution of this subset of parameters is standard. These tests, not reported here, did not change significantly the basic result. The rule was that the χ^2 statistics for the presence of noise decreased a little bit.

A set of regressions analogous to the one reported in Table 1 was carried out for the period prior to the stabilization plans (before 1986). These results are not reported here. The overall impression is that the same picture emerges as for the complete sample, despite localized differences. There is little noise. The regressions for which several lags of both inflation and the nominal rate were included as regressors were able to detect more noise. This may be attributed, however, to the effect of a large number of regressors in the smaller sample (128 observations) resulting from the exclusion of the period of stabilization plans.

The same econometric analysis was also performed for the period of the stabilization plans (1986 onwards). These results are not reported here. The fundamental conclusion that there is little noise is once again corroborated. With this small sample (54 observations), noise is highly exaggerated for the longer lags.

Chow tests for different information sets were also carried out. The results, not reported here, show that there is very weak evidence of structural change during the period of the stabilization plans. The overall conclusion is that the observed empirical corroboration of the Fisherian model does **not** hinge on the massive government intervention through the recent stabilization attempts.

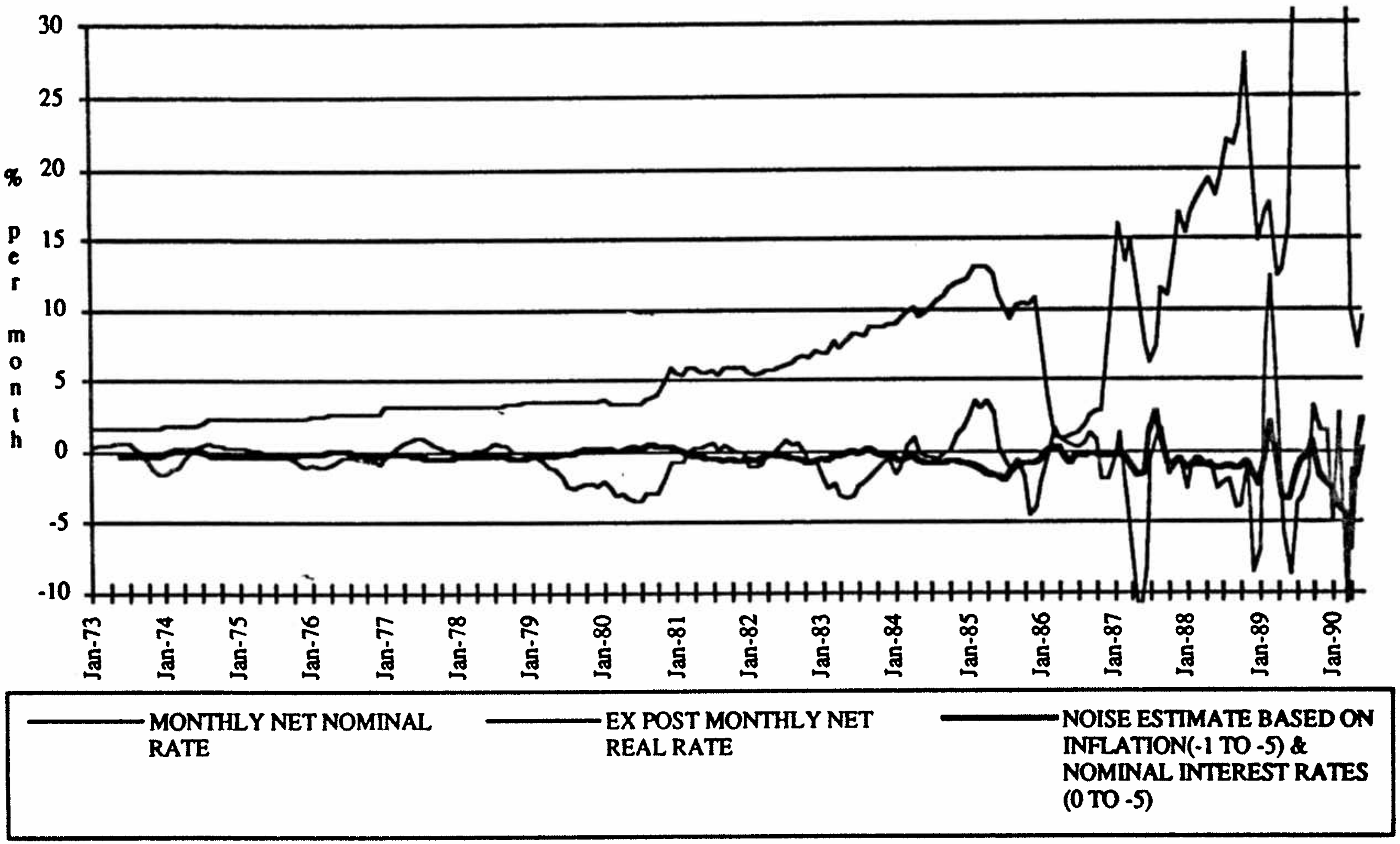

Figure 4: Noise Estimates Based on Lagged Inflation and on Current and Lagged Nominal Interest Rate

Several other interesting alternative hypotheses were contemplated by the inclusion of different regressors, as mentioned in the previous section. None of the regressors included, or combinations of them, was able to detect substantially more noise than the regressions in Table 1. The results for the whole sample are not reported here. I report in this paper the results of the inclusion of alternative sets of regressors for part of the sample, for reasons explained below.

An important objection to the above results regards the importance of the non-indexed CD market in Brazil in the '80s. Indexed CDs were first introduced in the early '80s (see Section 4). However, I was unable to find data on the ratio of the amounts transacted in both markets (indexed and non-indexed), except for the last 5 years. Indirect evidence from the market of government securities suggests that the volume of non-indexed securities transacted became less and less important vis-à-vis the volume of indexed securities transacted as inflation escalated, although this movement has not been monotonic. To guard against the criticism that the data used in the tests is from a thin market that does not reflect the financial transactions being made in the economy, a similar econometric analysis was performed for the period pre bank controls, 1973:1 to 1978:12. For this period it is reasonable to guarantee that the data is representative of the Brazilian CD market. For the period 1973 – 1978, the ex post real rate had a mean of –0.6 percent per annum with a standard deviation of 8.9 percent per annum. For the whole sample, those figures were -4.8 percent and 31.3 percent.

Tables 2 to 5 report the results for this sample period (1973:1 to 1978:12). The comparison of Tables 1 and 2 reveals that the noise ratios were higher for all lag lengths in the sub-sample, but not dramatically so. This effect is stronger for the longer lags. The fact that the noise ratios were higher may be attributed to the fact that inflation in the '70s was much more *well-behaved* than in the '80s, as Figure 1 shows. Therefore, one would presume that smaller inflation forecast errors would be incurred. To see this, remember that the two uncorrelated components of the ex-post real rate are the estimated tightest lower bound for model noise and the forecast error. Therefore, the ratio of noise variance to ex post real rate variance increases when the forecast error variance declines, ceteris paribus. However, the small sample size (72 observations) precludes any definitive conclusion. LAGS = 6 in Tables 1 and 2 correspond to 14 regressors (a constant, 6 inflation lags, and current and 6 lags of interest rate). It is unclear how much of the increase in noise detection is due solely to the increased number of regressors. To provide an idea of this effect, take the % of variance of ex post real rate for LAGS = 6 in Table 2, viz, 43 percent. This value is the R^2 of that regression. The $\bar{R}^2$, which corrects for the number of regressors, is significantly lower, viz, 28 percent.

The other marked difference between Tables 1 and 2 is the second row (% of variance of nominal rate) which is much higher in Table 2. This is due to the much smaller variance of the nominal rate during the '70s. Indeed, the variance of the nominal rate during 1973:1 to 1978:12 was smaller than the variance of the ex post real rate.

For that sub-period, ex post real interest rates exhibit marked seasonality. A regression of the ex post real rate against a constant and 11 monthly dummies provides an R^2 of 23 percent, as opposed to 6 percent for the whole sample (the $\bar{R}^2$s are 8.8 percent and 0.2 percent, respectively). Therefore, the measures of noise detection presented above are not very impressive.

TABLE 2

ORTHOGONALITY MEASURES AND TESTS BASED ON CURRENT AND LAGGED INTEREST RATES, AND LAGGED INFLATION: JANUARY / 1973 TO DECEMBER / 1978

	NUMBER OF LAGS INCLUDED					
	1	2	3	4	5	6
NOISE BOUNDS	BIC *					
	SC *					
% of variance of ex post real rate	12.5	17.6	20.4	25.9	35.5	43.0
% of variance of nominal rate	15.3	21.5	24.6	31.0	41.6	49.4
NOISE TESTS						
χ^2 statistic for all Inflation lags = 0	2.8	10.9	10.6	16.1	25.8	23.8
Degrees of freedom	1	2	3	4	5	6
Significance Level (%)	9.3	0.4	1.4	0.3	0.0	0.0
χ^2 statistic for all Nominal Rate lags = 0	5.9	9.5	15.3	26.4	28.4	58.5
Degrees of freedom	2	3	4	5	6	7
Significance Level (%)	5.2	2.3	0.4	0.0	0.0	0.0
χ^2 statistic for all Nominal Rate and Inflation lags = 0	14.2	27.3	36.3	82.8	99.2	140.1
Degrees of freedom	3	5	7	9	11	13
Significance Level (%)	0.3	0.0	0.0	0.0	0.0	0.0

Sample: 1973:1 to 1978:12 (the starting date varies with the number of lags).
Method of estimation: OLS with White's (1986) correction for the standard errors' estimates.
* BIC and SC criteria chose LAGS = 1.

Table 3 presents the results when a proxy for the output gap is included as a regressor. The proxy used is the residual of the regression of the (log) industrial production index on a constant, a time trend, and 11 seasonal dummies. I tried the same regressions for a measure of the gap which contemplated the existence of a unit root in the industrial production index.[12] The third alternative set of regressions undertaken was the one-step procedure previously described. The ex post real rate was regressed on a constant, the first-difference of (log) industrial production index, and 11 seasonal dummies. From all these alternatives, Table 3 presents the case in which most noise was detected. Regressions combining (several lags of) the proxy for the output gap, the inflation rate, and the nominal interest rate were also undertaken. The conclusion remains that noise, although present, is not extreme.

Table 4 presents the results when (log) monetary aggregates (divided by the price level) are used as regressors. The one measure that performed better in detecting noise was the M4 *gap*, which was constructed similar to the above output gap (see table 4). The Federal Government Debt was also tried as a regressor, but detected less noise than M1 or M4. Despite the increase in noise detection by the M4 *gap*, the results still indicate that the Fisher relation accounts for a significant part of the movement of interest rates. To provide some basis of comparison, the reader should refer to the study by Durlauf and Hall (1989b) of a similar model – the dividend-stock price model – which assumed risk-neutrality to price assets. In that context, the same noise ratios presented here are all near or above 100 percent, showing the complete failure of that model.

12 The Dickey-Fuller test rejected the existence of a unit root when a time trend was included in the test, and did not reject otherwise. The augmented Dickey-Fuller test (with p=12) did not reject the existence of a unit root for any case.

TABLE 3

ORTHOGONALITY MEASURES AND TESTS BASED ON LAGGED INDUSTRIAL PRODUCT GAP *

	NUMBER OF LAGS INCLUDED					
	1	2	3	4	5	6
NOISE BOUNDS	BIC ** SC **					
% of variance of ex post real rate	21.3	24.1	23.5	24.3	24.0	25.3
% of variance of nominal rate	26.2	29.4	28.4	29.0	28.1	29.1
NOISE TESTS						
χ^2 statistic for LAGS = 0	7.6	9.8	10.5	15.8	16.3	16.3
Degrees of freedom	1	2	3	4	5	6
Significance Level (%)	0.6	0.7	1.5	0.3	0.6	1.2
χ^2 statistic for LAGS & constant = 0	7.9	10.5	11.7	18.4	17.4	17.3
Degrees of freedom	2	3	4	5	6	7
Significance Level (%)	1.9	1.4	1.9	0.2	0.8	1.6

Sample: 1973:1 to 1978:12 (the starting date varies with the lags).
Method of estimation: OLS with White's (1986) correction for the standard errors' estimates.
* This measure of the industrial product gap was constructed by removing the deterministic trend and the seasonal components. This measure proved to be the most efficient one in detecting noise.
** BIC criterion chose LAGS = 0; SC, LAGS = 1; Akaike, LAGS = 24 out of a possible 24.

TABLE 4

ORTHOGONALITY MEASURES AND TESTS: M1 AND M4

	NUMBER OF LAGS INCLUDED					
	1	2	3	4	5	6
Variable: **M1 (REAL)**	BIC		SC			
% of variance of ex post real rate	21.1	22.1	22.6	23.7	25.5	32.1
% of variance of nominal rate	26.1	27.1	27.4	28.3	29.9	36.9
χ^2 statistic for LAGS = 0	14.6	17.2	27.3	42.5	54.2	67.5
Degrees of freedom	1	2	3	4	5	6
Significance Level (%)	0.0	0.0	0.0	0.0	0.0	0.0
Variable: **M4 (REAL)**	BIC SC					
% of variance of ex post real rate	0.0	4.6	5.0	5.8	8.9	18.5
% of variance of nominal rate	0.0	5.6	6.1	7.0	10.4	21.3
χ^2 statistic for LAGS = 0	0.0	3.5	4.5	4.3	4.6	28.2
Degrees of freedom	1	2	3	4	5	6
Significance Level (%)	98.1	17.4	21.3	36.7	46.3	0.0
Variable: **ΔM4 (REAL)**	BIC SC					
% of variance of ex post real rate	4.3	4.3	4.5	7.2	17.0	34.5
% of variance of nominal rate	5.3	5.2	5.4	8.4	19.5	38.9
χ^2 statistic for LAGS = 0	3.6	4.4	4.3	4.0	24.0	72.2
Degrees of freedom	1	2	3	4	5	6
Significance Level (%)	5.7	11.0	23.3	40.3	0.0	0.0
Variable: **M4 *GAP** (REAL)**	BIC SC					
% of variance of ex post real rate	45.8	48.6	49.6	49.8	49.3	48.4
% of variance of nominal rate	56.5	59.4	60.1	59.4	57.7	55.6
χ^2 statistic for LAGS = 0	43.5	60.5	61.0	67.4	77.7	85.5
Degrees of freedom	1	2	3	4	5	6
Significance Level (%)	0.0	0.0	0.0	0.0	0.0	0.0

Sample: 1973:1 to 1978:12 (the starting date varies with the lags).
Method of estimation: OLS with White's (1986) correction for the standard errors' estimates.
* M4 *GAP* stands for the residual of the regression of M4 (REAL) on a constant, a time trend, and 11 seasonal dummies. The constant estimate is added later to the residual.

Table 5 presents the attempt to infer a link between the Brazilian and US capital markets. The (admittedly expedient) rationale for such a regression was presented in the previous section. The results show that the spread between the black market exchange rate and the official exchange rate has some explanatory power, although the nominal rate of US CDs does not. The extension of this analysis to the whole sample period shows that the link was much weaker (the noise ratios were lower) and that all the explanatory power continues to come from the black market spread. However, as noted previously, after 1979, this market may have been too thin to be representative.

Finally, it is worthwhile to mention that lagged ex post real rates were also used as explanatory variables. Since the ex post real rate for a 6 month CD bought at time t is not known until $t + 6$, the most recent lag that can be incorporated in the regressions as an explanatory variable is $t - 6$. Lags of the ex post real rate were not very useful in detecting noise.

In summary, the noise ratios were higher for the sub-period 1973 / 1978, for which it can be guaranteed that non-indexed CDs constituted the bulk of the CD market, than for the whole sample (1973 / 1990). This may be attributed to the stochastic process followed by inflation, which became erratic and explosive in the '80s inducing larger forecast errors and, consequently, more volatile ex post real rates. From all the regressors that were suggested by alternative theoretical models to the Fisher relation, the monetary aggregate M4 *gap* (see explanation of its construction above) was the most effective in detecting noise. The increase in noise detection has to be interpreted with caution because of the effect of a large number of regressors in a small sample. When compared to a similar asset pricing model that assumes risk-neutrality – the dividend-stock price model – the Fisher model performs substantially better. Therefore, the empirical evidence cannot reject the existence of a Fisher effect in Brazil, although other factors also influenced interest rates.

TABLE 5

ORTHOGONALITY MEASURES AND TESTS BASED ON CURRENT AND LAGGED INTEREST RATES ON US CDs, AND LAGGED SPREAD BETWEEN THE BLACK AND OFFICIAL US$ MARKETS IN BRAZIL

	NUMBER OF LAGS INCLUDED					
	1	2	3	4	5	6
NOISE BOUNDS	BIC * SC *					
% of variance of ex post real rate	16.6	21.8	26.3	28.9	31.0	34.2
% of variance of nominal rate	20.5	26.7	31.9	34.5	36.3	39.3
NOISE TESTS						
χ^2 statistic for all Black / Official Market spread lags = 0	7.0	10.7	15.7	41.3	45.6	41.7
Degrees of freedom	1	2	3	4	5	6
Significance Level (%)	0.8	0.5	0.1	0.0	0.0	0.0
χ^2 statistic for all US Nominal Rate lags = 0	1.4	4.2	4.6	4.8	5.8	5.3
Degrees of freedom	2	3	4	5	6	7
Significance Level (%)	49.2	24.4	33.6	43.9	45.0	62.0
χ^2 statistic for all US Nominal Rate and Black / Official Market spread lags = 0	8.5	23.9	46.9	101.9	123.5	105.6
Degrees of freedom	3	5	7	9	11	13
Significance Level (%)	3.7	0.0	0.0	0.0	0.0	0.0

Sample: 1973:1 to 1978:12 (the starting date varies with the number of lags).
Method of estimation: OLS with White's (1986) correction for the standard errors' estimates.
* BIC and SC criteria chose LAGS = 1.

e. Robustness of the Results to Inflation Measurement

The price index used in this paper, IGP-DI (General Price Index - Domestic Supply) compiled by the Fundação Getúlio Vargas, has been computed on a calendar month basis since March 1986. It is centered on the 15th day of the month, and thus it measures the inflation from the middle of the previous month to the middle of the current month, assuming a linear approximation. The ideal inflation measure for this paper would be one measuring the inflation from the beginning of the current month to its end. Since this measure is not available, it can be approximated by a geometric mean with equal weights of the available inflation measure of periods t and $t+1$ (before March 1986, the weights were 1/3 and 2/3). I call this index IGP-DI adjusted.

All the preceding analysis is repeated with the new index. These results are not reported here. The use of the IGP-DI adjusted seems to increase noise detection in many cases, but not dramatically so. The Chow tests make a somewhat stronger case for a regime change after the Cruzado plan Nevertheless, the conclusion remains true that the Fisher model satisfactorily describes the recent Brazilian experience.

4. FISHER EFFECT FOR INDEXED SECURITIES

The indexed securities studied in this paper were indexed to the "monetary correction" index fixed by the Brazilian government. Until recently, this was the only indexation allowed (for sometime the government also allowed indexation to the U.S. dollar exchange rate). Given the exchange controls of the Brazilian economy, it is reasonable to assume that the legal indexation was the one mostly used by market participants.

Monetary correction has been an all but perfect indexation mechanism. Figure 5 shows how the monetary correction lagged behind official inflation (a spliced series composed of different price indices used by the Brazilian government). During the last three years of the '80s, the government adopted the IPC / IBGE (Consumer Price Index published by the Fundação Instituto Brasileiro de Geografia e Estatística) as the official inflation index. Since June 1987 the IPC has measured the average price level from the 15th day of one month until the 14th day of the subsequent month (Clifton, 1990). Therefore, the IPC is centered on the 30th day of the month, and thus the inflation rate calculated with the IPC of month $t+1$ is a reasonable approximation for the actual inflation of month t.

Given this one month lag in measuring inflation, it is reasonable to assume that the *promised* real rates of indexed securities should vary to accommodate discrepancies between the expected values of actual and measured inflation. In the late '80s inflation was escalating in Brazil. It is widely believed that in this period investors perceived part of the *promised* real rates of indexed securities as a compensation for inflation which would show up only on next month's index.

Figure 6 displays the *promised* net real rate of indexed securities, the net real rate when the official inflation is used, and the net real rate when the adjusted official inflation is used. Private banks began to issue indexed CDs only when the inflation level increased substantially at the beginning of the '80s (Ferreira, 1990). The gaps are due to the Cruzado plan (February 1986) and the Summer plan (January 1989), when the government tried to do away with indexation because of its perceived impact on inflation. In June 1987, under the Bresser plan, the government changed the calculation of the IPC. This distorted the calculation of the real interest rate in May and June 1987. While inflation escalates in the latter period (Figure 6), it is clear that the extremely high *promised* real rates do not correspond to high ex post real rates, when deflated by the adjusted IPC. For the whole sample, the mean *promised* real rate was 16.4 percent per annum, while the mean ex post real rate was only 2.3 percent (the deflator was the IGP adjusted).

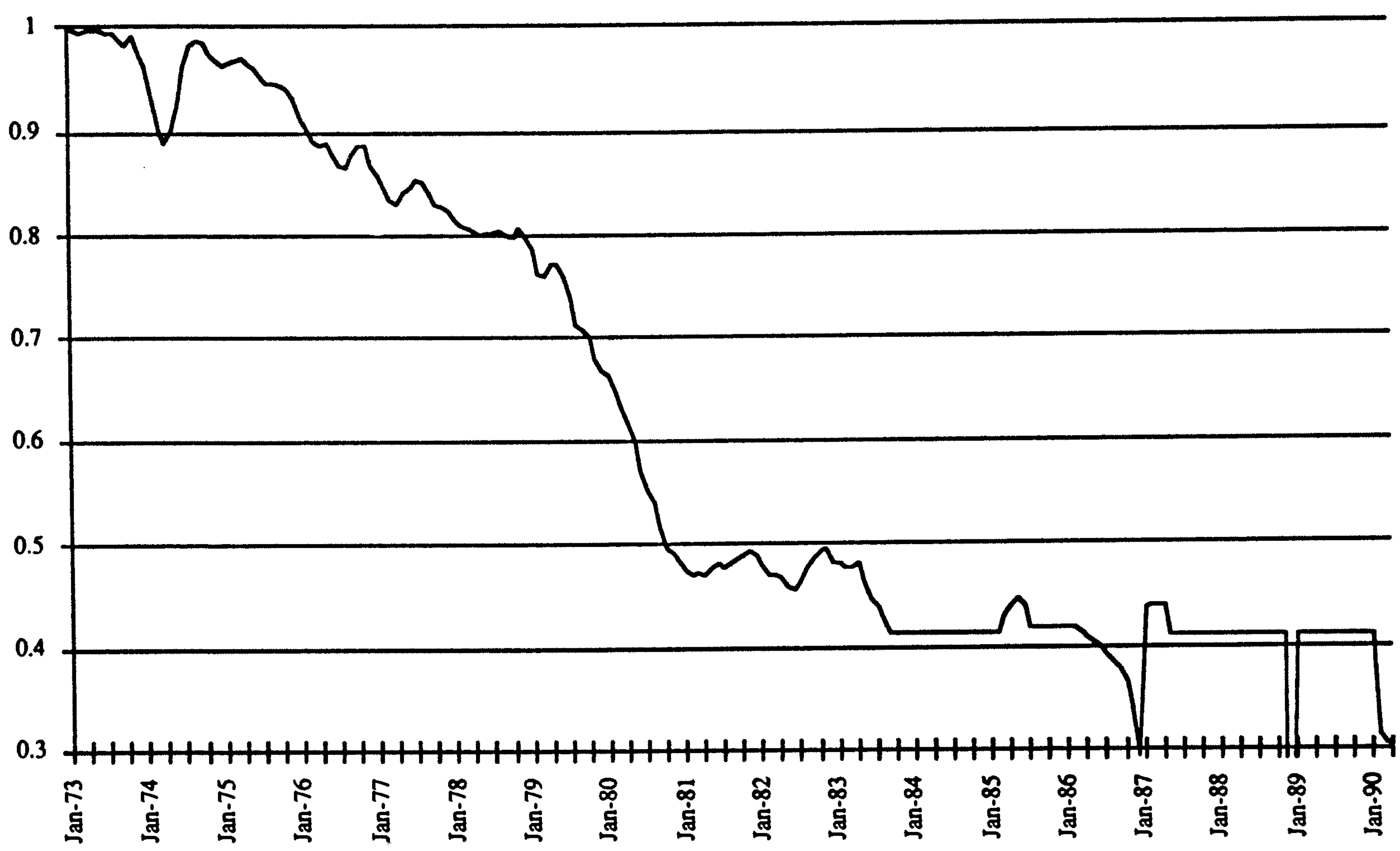

Figure 5: Evolution of the Ratio [OTN / Official Price Level]

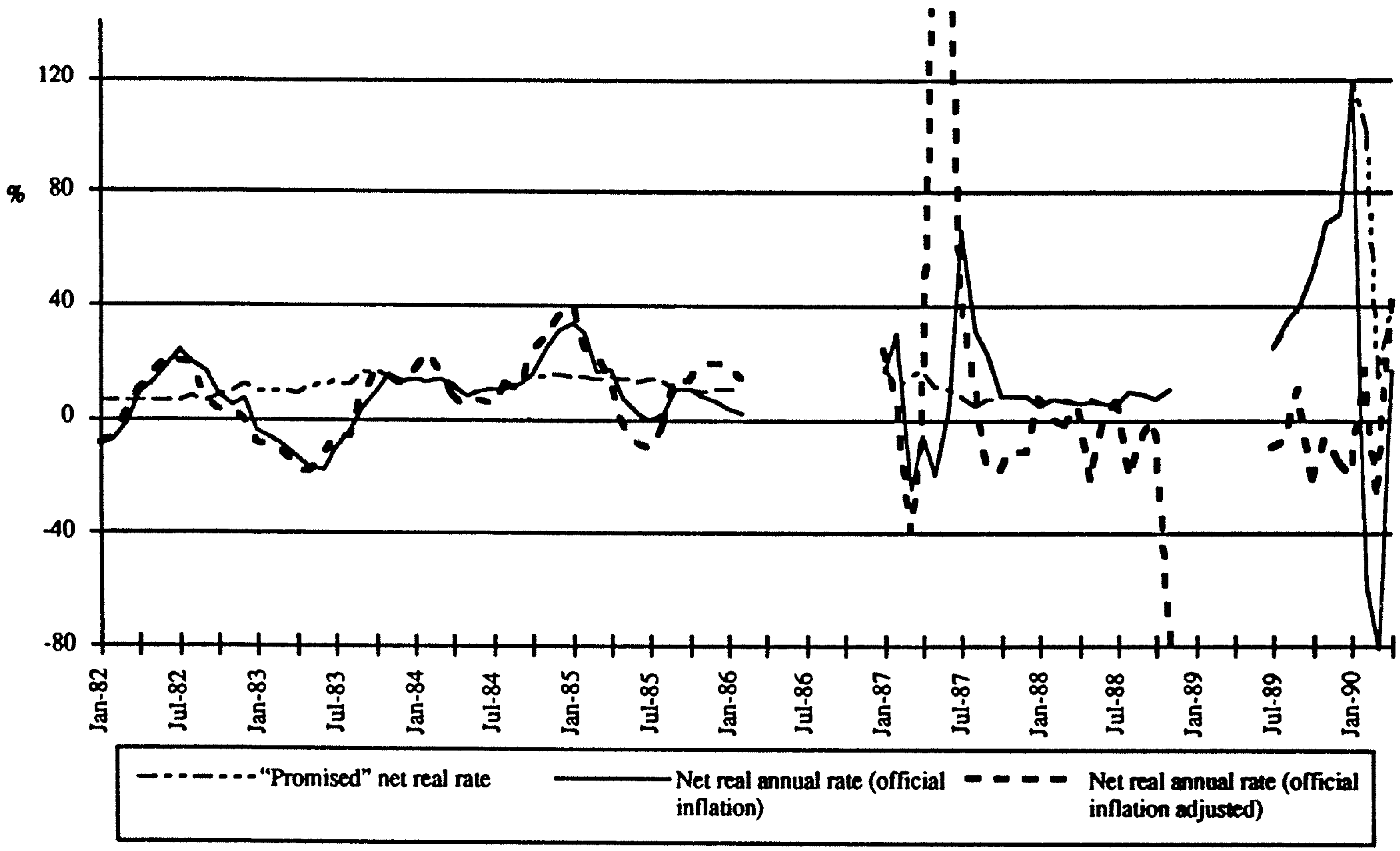

Figure 6: Indexed CDs' Net of Tax Real Rates

Equation (10) represents this Fisher effect for indexed securities. As in the previous model, it is assumed that the *expected* real rate of return is constant.

$$pr_t = \rho - MI^e_{t+1|t} + AI^e_{t+1|t} \qquad (10)$$

where pr_t = *promised* real interest rate from t to t+1 (%);
$MI^e_{t+1|t}$ = expected (at t) **measured** inflation rate from t to t+1 (%);
$AI^e_{t+1|t}$ = expected (at t) **actual** inflation rate from t to t+1 (%).

If inflation is measured by an index such as the IPC, for which $MI_{t+1} = AI_t$, equation (10) will become equation (11), the null model. Equation (12) subsumes all alternative linear models, by adding the noise term, S_t.

$$\mathrm{H}_0: \; pr_t = \rho + (MI^e_{t+2|t} - MI^e_{t+1|t}) \qquad (11)$$

$$\mathrm{H}_1: \; pr_t = \rho + (MI^e_{t+2|t} - MI^e_{t+1|t}) + S_t \qquad (12)$$

By rational expectations:

$$MI_{t+1} = MI^e_{t+1|t} + v'_{t+1} \qquad (13)$$

$$MI_{t+2} = MI^e_{t+2|t} + v''_{t+1} + v_{t+2} \qquad (14)$$

where v'_{t+1}, v''_{t+1}, and v_{t+2} are forecast errors.

Subtracting (14) from (13), and adding the result to (12) provides equation (15).

$$r_t = \rho + S_t - (v_{t+2} + v_{t+1}) \qquad (15)$$

where $v_{t+1} = v''_{t+1} - v'_{t+1}$ is a forecast error term (white noise);

r_t = $pr_t - (MI_{t+2} - MI_{t+1})$ is the actual ex post real rate of indexed securities.

Equation (15) is analogous to equation (4), except for the MA(1) term.[13] Appendix 2 presents the equivalence between the model developed for indexed securities and the one developed for nominal securities. As before, the estimation of the model is conducted by regressing the actual ex post real rate on current and lagged *promised* real rates, lagged "adjusted" inflation, and a constant. To obtain consistent estimates of the standard errors, White's (1986) heteroskedasticity-autocorrelation consistent covariance matrix estimator is employed.

Table 6 presents the results when lagged "adjusted" inflation and current and lagged *promised* real interest rates are used as regressors. The overall picture is not much different from that shown in Tables 1 and 2. Noise, although present, does not account for the bulk of the movement in the ex post real or the *promised* real rates. Lagged ex post real rates were also used as regressors. Since the longest maturity observed in the period analyzed was six months, only the sixth lag (and above) of the ex post real rate is allowed as a regressor. Adding the lagged ex post real rate does not increase noise detection. When first-differenced *adjusted* inflation is substituted for *adjusted* inflation in the regression, noise detection does not increase either. These results are not reported here.

Therefore, the econometric tests performed here did not provide strong evidence against the existence of a Fisher effect for indexed securities.

13 The MA(1) term is analogous to the one that shows up in equation (4) when the securities' maturity is longer than the observation period, only that in the current framework, a q-period security originates an MA(q) error, instead of an MA(q-1). This is due to the first-difference of inflation rates in the model.

TABLE 6

INDEXED SECURITIES: ORTHOGONALITY MEASURES AND TESTS BASED ON CURRENT AND LAGGED *PROMISED* REAL INTEREST RATES, AND LAGGED INFLATION

	NUMBER OF LAGS INCLUDED					
	1	2	3	4	5	6
NOISE BOUNDS	BIC *	SC *				
% of variance of ex post real rate	7.2	9.6	17.2	17.8	22.0	32.8
% of variance of nominal rate	13.7	17.6	26.1	24.6	30.0	39.3
NOISE TESTS						
χ^2 statistic for all Adjusted Inflation lags = 0	3.1	23.5	10.5	15.8	12.4	18.6
Degrees of freedom	1	2	3	4	5	6
Significance Level (%)	7.7	0.0	1.5	0.3	3.0	0.5
χ^2 statistic for all *Promised* Real Rate lags = 0	3.9	3.1	9.0	9.3	11.3	92.7
Degrees of freedom	2	3	4	5	6	7
Significance Level (%)	14.2	37.7	6.0	9.9	7.9	0.0
χ^2 statistic for all *Promised* Real Rate and Adjusted Inflation lags = 0	6.8	27.5	16.4	22.3	45.1	405.0
Degrees of freedom	3	5	7	9	11	13
Significance Level (%)	7.8	0.0	2.1	0.8	0.0	0.0

Sample: 1982:1 to 1986:2, 1987:1 to 1988:11, and 1989:7 to 1990:4 (the starting date varies with the number of lags).
Method of estimation: OLS with White's (1986) correction for the standard errors' estimates.
* BIC criterion chose LAGS=1; SC, LAGS = 2.

5. CONCLUSION

The Fisherian proposition that nominal interest rates adjust one-for-one to changes in inflation expectation is one of the most basic propositions learned in economics. It arises in models with risk neutral agents and an asset that pays a fixed expected real rate of return. The large literature that tests this proposition does not grant it unrestricted support.

This paper has shown that all tests of the Fisher equation can be subsumed under the signal extraction approach, as laid out by Durlauf and Hall (1989a). All tests can be reinterpreted as the projection of the ex post real interest rate on different information sets. If the model is plausible, this projection should be near zero, that is, forecast errors, not model noise, account for the bulk of the movements in ex post real interest rates. In other words, the ex ante real interest rate is approximately constant.

The recent Brazilian experience is an interesting setting in which to analyze the Fisher effect. On the one hand inflation was extremely high and volatile; therefore, its swings should dominate the adjustments of nominal interest rates. On the other hand, the ex post real interest rate varied greatly: From 1973 to 1990 it peaked at 300 percent per year, with a trough of -80 percent per year and a standard deviation of 32 percent per year.

The tests for the whole sample (January 1973 to June 1990) have shown that most variation in ex post real interest rates may be attributed to inflation forecast errors. The Fisher model seems to satisfactorily fit the Brazilian data. Very little movement on nominal rates (around 1 percent) is due to factors other than inflation expectation. This conclusion is robust to corrections in inflation measurement which account for lag in measurement and averaging in constructing the price index.

To guard against the criticism that the market for non-indexed CDs may have been too thin in the '80s, the estimation was conducted for the sub-period 1973 - 1978, for which it can be guaranteed that non-indexed CDs constituted the bulk of the CD market. For this sub-period, the noise ratios increased, meaning that factors extraneous to the Fisher model accounted for a larger part of the nominal interest rate movements. This may be attributed to the stochastic process followed by inflation, which became erratic and explosive in the '80s inducing larger forecast errors and, consequently, more volatile ex post real rates. Noise and forecast errors are the two uncorrelated components of the ex post real rate. If noise, i.e., other factors not accounted for in the Fisher model, remained the same in the '80s as in the '70s, its comparative explanatory power would decrease in the '80s vis-à-vis the '70s. From all regressors that were suggested by theoretical alternatives to the Fisher relation, the monetary aggregate M4 *gap* (a detrended and seasonally adjusted measure of the ratio M4 / Price level) was the most effective in detecting noise. The increase in noise detection has to be taken with reservations because of the effect of a large number of regressors in a small sample. When compared to a similar asset pricing model that assumes risk-neutrality – the dividend-stock price model (see Durlauf and Hall, 1989b) – the Fisher model performs substantially better. Therefore, the empirical evidence cannot reject the existence of a Fisher effect in Brazil, although other factors also influenced interest rates.

One could argue that with the extremely large inflation and nominal interest rates in Brazil, the results obtained here still leave room for the ex ante real interest rate to vary. Nevertheless, the results validate the idea that the recent Brazilian experience has been one in which a passive monetary policy led to a fairly constant ex ante real rate. If similar results hold for the rates on government debt, it is unwarranted to assume that non-indexed debt financed fiscal deficits have lower cost. Therefore, there is no easy way out of the need for fiscal control.

When inflation escalates rapidly as it did in Brazil in the second part of the '80s, the inflation index, which guides financial indexation, underestimates true inflation. With rational investors aware of this statistical problem, the *promised* real rates of indexed securities should rise to account for this fact, although the actual ex ante real rate may not vary as much. The DH test of this model did not detect an overwhelming amount of noise (not more than 33 percent), validating the existence of this Fisher effect for indexed securities.

Bearing in mind the extremely high investments rates observed in Brazil in the '70s, one could wonder how to harmonize these with, on average, negative ex post real rates (the mean was -0.6 for the period 1973 / 1979). The results presented above suggest that inflation forecast errors were the main culprits for the high variability of the ex post real rate. This, in turn, suggests that further research regarding how inflation expectations were formed is in order.

APPENDIX 1. REINTERPRETATION OF THE LITERATURE IN THE SIGNAL EXTRACTION FRAMEWORK

As explained in Section 2, all different tests of the null model are subsumed in the projection of $[R_t - \pi_{t+1}]$ on the econometrician's information set, $L_x(t)$. If model noise, i.e., other factors not taken into account by the model, is important, this projection will differ significantly from zero. In other words, all tests of equation (1), when reinterpreted in the DH framework, differ only by the $L_x(t)$ on which $[R_t - \pi_{t+1}]$ is projected. Therefore, I next characterize the $L_x(t)$ for each different test. Since there are many different tests of the Fisher equation, this appendix shall follow the taxonomy suggested by Rocha (1988) to partition the empirical literature on the Fisher effect.

a. Traditional Criterion

These tests consist of regressions of the nominal interest rate, R_t, on proxies for expected inflation and other variables known at t. The idea behind these tests is that, were π^e_{t+1} observable and H_0 true, the regression of R_t on π^e_{t+1} and a constant would yield a unitary coefficient for π^e_{t+1} and the constant would equal the unconditional expectation of the real interest rate.

Gibson (1970a) runs equation (A.1) in levels and in first-differences.

$$R_t = \alpha_0 + \alpha_1 \pi_t + \ldots + \alpha_{10} \pi_{t-9} \tag{A.1}$$

In the DH framework, this test corresponds to restricting $L_x(t)$ to the space spanned by lagged inflation rates. Section g of this appendix contains the proof of this equivalence for the simple case where inflation is an AR(1). Proofs for the tests commented on below can be similarly constructed.

Gibson (1970b) is concerned about the liquidity effect, i.e., the downward short-run impact of money growth on interest rates. He regresses R_t on a constant and current and lagged logarithms of money supply. The DH framework cannot answer Gibson's questions on the existence and timing of the liquidity effect. As far as the test of the null model (equation (1)) is concerned, his test is equivalent to restricting $L_x(t)$ to the space spanned by current and lagged money supplies.

Gibson (1972) tries to ameliorate the errors-in-variables problem involved with getting a proxy for inflation expectation by using data on inflation expectations from the Livingston survey. This test is equivalent to restricting $L_x(t)$ to the space spanned by current inflation expectation, as measured by the Livingston survey.

Sargent (1973) runs equation (A.2).

$$R_t = \alpha_0 + \sum_{i=1}^{m} w_i \cdot \pi_{t-i} \tag{A.2}$$

Sargent is actually looking for the "optimal" lag length, m. The DH framework cannot answer this question. As far as the test of equation (1) is concerned, his test is equivalent to restricting $L_x(t)$ to the space spanned by lagged inflation rates.

Lahiri (1976) follows a test similar to Gibson's (1972), by regressing R_t on a constant and on a surveyed inflation expectation. However, he uses 2SLS with lagged inflation rates as instruments. In the DH framework this is equivalent to restricting $L_x(t)$ to the space spanned by lagged inflation rates.

Levi and Makin (1979) use the surveyed inflation expectation data, output growth and the standard deviation of the surveyed inflation expectation as regressors. This is equivalent to restricting $L_x(t)$ to the space spanned by those variables.

Tanzi (1980) regresses R_t on a constant, the output gap and on the surveyed inflation expectation. Like Lahiri (1976) he uses 2SLS to estimate the coefficient for the inflation expectation. In the DH framework, this is equivalent to restricting $L_x(t)$ to the space spanned by the output gap and the instruments (lagged inflation rates).

Startz (1981) runs equation (A.3).

$$R_t = \alpha_0 + \gamma\, \pi_{t+1} + \beta\, \hat{u}_t \tag{A.3}$$

where $\hat{u}_t$ = unemployment gap.

The interpretation of this test on the DH framework is not as immediate as the other ones. This is because Startz's assumptions provide an instrumental variable for the forward projection, namely $[\pi_{t+1} - \beta\hat{u}_t]$.[12] Therefore, the equivalent DH test projects $[R_t - \pi_{t+1}]$ on the space $L_\pi(t+1) \ominus L_{\hat{u}}(t+1)$.[13]

In an appendix to his 1983 paper, Summers regresses R_t on a constant and one to eight times lagged values of π_t. He also uses 2SLS with π_{t+1} as regressor and the eight lagged values of π_t as instruments. Both tests are equivalent to restricting $L_x(t)$ to the space spanned by the eight lagged inflation rates.

Other papers in the same line have been written by Peek (1982) and Blejer and Eden (1979). The first one is equivalent to restricting $L_x(t)$ to the space spanned by the monetary and fiscal policy variables constructed by Peek, lagged output and surveyed inflation expectation; the second one, to the space spanned by surveyed inflation expectation and its variance.

b. Fama's Criterion

These tests consist of a regression of π_{t+1} on a constant, R_t, and other variables known at t. The rationale behind this test is to avoid the errors-in-variables problem associated with the unobservability of inflation expectation.

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12 Startz postulates:

H_0:	$R_t = r_t + \pi^e_{t+1}$	(a)
Rational Expectations:	$\pi_{t+1} = \pi^e_{t+1} + \nu_t$	(b)
Phillips Curve:	$\pi_{t+1} = \pi^e_{t+1} + \beta\hat{u}_t$	(c)
Real Interest Rate:	$r_t = \rho + \varepsilon_t$	(d)

To guarantee consistency he must assume orthogonality between ε_t and both π_{t+1} and $\hat{u}_t$. In his own words: "... In my regression, residuals arise from variation in the real rate of interest. Under neutrality, these are uncorrelated with both û and π (Startz (1981), p. 972)." To my knowledge, new classical macroeconomics (what Startz calls "neutrality") does not postulate that variations in the real interest must be uncorrelated with contemporaneous inflation and unemployment.

13 $L_\pi(t+1) \ominus L_{\hat{u}}(t+1)$ is the space that contains the leaded inflation rate but does not contain the leaded unemployment gap. In other words, Startz claims to be able to identify the part of leaded inflation that is orthogonal to noise.

Fama (1975) runs equation (A.4). The DH equivalent test is obtained by restricting $L_x(t)$ to the space spanned by the current nominal interest rate. Section h of this appendix presents the proof of this equivalence.

$$\pi_{t+1} = \alpha_0 + \alpha_1 R_t + \varepsilon_t \tag{A.4}$$

Joines (1977) adds lagged inflation rates as regressors in equation (A.4). This is equivalent to restricting $L_x(t)$ to the space spanned by the current nominal interest rate, and lagged inflation rates.

Other papers in this line are Carlson (1977), who includes the employment/population ratio in Fama's regression, and Nelson and Schwert (1977), who include an ARMA forecast for inflation. The equivalent DH test is constructed by allowing those variables as additional vectors spanning $L_x(t)$.

c. Mishkin's Criterion

Mishkin (1981, 82) is concerned with the constancy of real rates. He uses *ex post* real interest rates to test whether *ex ante* real rates are constant or not. Mishkin's test is already in the DH orthogonality test format, since the ex-post real interest rate is $[R_t - \pi_{t+1}]$. He runs several regressions with alternate sets of regressors, which correspond to different characterizations of the information set, $L_x(t)$, on the DH framework. Those variables are lagged inflation, time polynomials, money growth, output growth, output gap, unemployment rate, and investment/capital ratio.

Hafer and Hein (1982) use the same framework to test for the impact of real money on real rates. As explained before, the DH framework has nothing to say about the timing of this liquidity effect. As far as the null model (equation 1) is concerned, their test is equivalent to restricting $L_x(t)$ to the space spanned by once and twice lagged real money stocks.

d. Friedman's Criterion

Friedman (1978, 1980a, 1980b) sets up a model, estimated for different agents (life insurance companies, pension funds, etc), for the supply and demand for bonds. Among several other factors, these demand and supply functions depend on the nominal interest rate and on the expected inflation. The estimation of these equations provides the coefficients needed to evaluate $\partial R_t / \partial \pi^e_{t+1}$. Apart from details of estimation, this method is equivalent to restricting $L_x(t)$ to the space spanned by all the variables of his model.

e. Summers' Criterion

Summers (1983) uses band-spectrum regression in the following regression:

$$R_t = \alpha_0 + \alpha_1 \pi_{t+1} + u_t \tag{A.5}$$

where u_t = random error.

Summers' paper generated a controversy (McCallum, 1984; Summers, 1984; Barsky, 1987). None of these criticisms of Summers observed that the use of band-spectrum regression may generate inconsistent estimates. The problem is that the filtered π_{t+1} may not be orthogonal to u_t, rendering $\hat{\alpha}_1$ inconsistent. Therefore, I shall not discuss this criterion further.

f. Previous Brazilian empirical evidence

The literature on the validity of the Fisher equation for Brazil has followed the criteria just reviewed. For a review of this literature, see Rocha (1988, p. 97), who also reexamines the conventional, Fama's and Mishkin's criteria for the period 1972:12 to 1979:6, using government bonds interest rates.

Equivalence between the conventional criterion and the DH Test

This criterion depends on the stochastic process assumed for inflation. For simplicity, assume that inflation is an AR(1) process:

$$\pi_{t+1} = a\,\pi_t + v_t$$

Assume also Rational Expectations, and the Fisher Effect:

$$\pi^e_{t+1} = a\,\pi_t$$

$$R_t = \rho + \pi^e_{t+1}$$

The conventional criterion is:

$$R_t = \alpha + \beta\pi_t + \varepsilon_t$$

Under H_0, $\text{plim}(\hat{\beta}) = a$ and $\text{plim}(\hat{\alpha}) = \rho$.

The DH equivalent test is:

$$(R_t - \pi_{t+1}) = \alpha' + \beta'\pi_t + \varepsilon'_t$$

Given the above assumptions, the LHS is just $\rho + v_t$. Therefore, $\text{plim}(\hat{\beta}') = 0$ and $\text{plim}(\hat{\alpha}') = \rho$. The DH equivalent test to the traditional criterion is constructed by regressing the ex post real rate on the same variables used to proxy for expected inflation.

Equivalence between Fama's Criterion and the DH Test

This criterion does not depend on the stochastic process assumed for inflation. By Rational Expectations:

$$\pi_{t+1} = \pi^{e}_{t+1} + v_t$$

Assume also that the Fisher Effect holds:

$$R_t = \rho + \pi^{e}_{t+1}$$

Fama's regression is:

$$\pi_{t+1} = \alpha + \beta R_t + \varepsilon_t$$

Under H_0, $\text{plim}(\hat{\beta}) = 1$ and $\text{plim}(\hat{\alpha}) = -\rho$.

The DH equivalent test is:

$$(R_t - \pi_{t+1}) = \alpha' + \beta' R_t + \varepsilon'_t$$

Given the above assumptions, the LHS is just $\rho + v_t$. Therefore, $\text{plim}(\hat{\beta}') = 0$ and $\text{plim}(\hat{\alpha}') = \rho$. The DH equivalent test to the Fama criterion is constructed by regressing the ex post real rate on the nominal interest rate.

APPENDIX 2. ANALOGY BETWEEN THE NOMINAL AND THE INDEXED MODELS

NOMINAL MODEL	INDEXED MODEL
R_t	pr_t
π_{t+1}	$(MI_{t+2} - MI_{t+1})$
$(R_t - \pi_{t+1})$	$pr_t - (MI_{t+2} - MI_{t+1})$
$H_0 : R_t = \rho + \pi^e_{t+1}$	$H_0 : pr_t = \rho + (MI_{t+2} - MI_{t+1}$
$R_t - \pi_{t+1} = \rho + S_t - \nu_{t+1}$	$r_t = \rho + S_t - (\nu_{t+2} + \nu_{t+1})$

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