Infrequent large nominal devaluations and their impact on the futures prices for foreign exchange in Brazil*

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Summary: 1. Introduction; 2. Institutional background; 3. Unconditional unbiasedness; 4. Conditional unbiasedness; 5. Excess return under incomplete information; 6. Conditioning the probability distribution on the event of a major devaluation; 7. Kalman filter modelling of excess returns; 8. Concluding remarks.
Key words: futures market; exchange rate; exchange rate policy; generalized method of moments.

This paper discusses the behavior of futures prices for foreign exchange in Brazil during a period of high inflation and successive stabilization attempts (1989-92). After testing for futures prices unbiasedness and predictability by applying the generalized method of moments, the paper argues that the finding of excess returns may be viewed as a rational response to the frequent and unpredictable changes in the exchange rate policy during that period. This response could reflect an informational problem where the exchange rate policy is assumed to be unknown; or, a "peso" problem of rational (under) overprediction where the futures bias is the market response to the known policy of infrequent large nominal devaluations. The second line of explanation is suggested by conditioning the probability distribution of the excess return of futures contracts on the event of a major devaluation.

O artigo discute o comportamento dos preços futuros da taxa de câmbio no Brasil durante um período de inflação alta e sucessivas tentativas de estabilização (1989-92). Após testar a existência de viés na formação de expectativas e na antecipação dos preços futuros para a taxa de câmbio via o método generalizado de momentos, o artigo propõe que os retornos excessivos no mercado futuro sejam vistos como uma resposta racional à política de desvalorizações infreqüentes e/ou de mudanças não-antecipadas da política cambial no período. Esta resposta pode refletir um problema de informações, em que a política cambial é desconhecida; ou um "problema do tipo peso", isto é, super (sub) predição, onde o viés futuro é a resposta racional do mercado à política de grandes e infreqüentes desvalorizações nominais da moeda. Esta segunda linha de explicação é sugerida, condicionando a distribuição de probabilidades dos retornos no mercado ao evento de uma desvalorização nominal.

1. Introduction

An important question in economics is whether markets successfully aggregate information. It is often argued that prices in speculative markets such as futures and options possess stochastic properties that correspond to theories of efficient markets. This paper examines the behavior of future prices for foreign exchange during a period of high inflation and successive stabilization attempts in Brazil. The recent Brazilian experience with the liberalization of the market for foreign exchange and the creation of the futures contract for foreign exchange in the São Paulo Futures Board of Trade provides a good benchmark for analysis.

The market for foreign exchange in Brazil was highly unstable during the period of analysis (1989-92). During this period, Brazil followed a flexible exchange rate policy under which the exchange rate for the domestic currency was adjusted on a daily basis in terms of the US dollar. The Central Bank would stand ready to purchase exchange from the banks and

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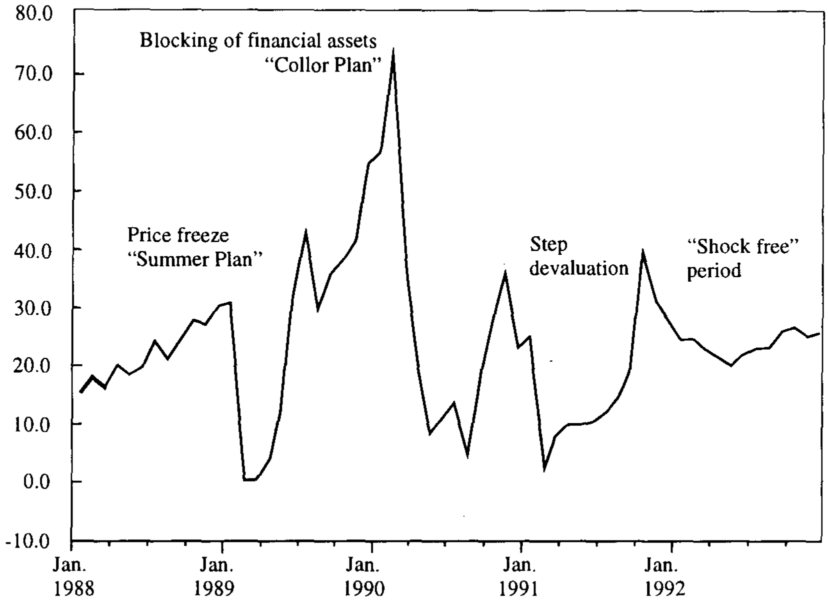
to sell it to the bank at the official rate for approved transactions. This period can be broadly divided in the following three subperiods: from March 1989 to March 1990, which begins with the introduction of the Summer Plan and ends with the introduction of the Collor Plan; the subperiod covering the Collor Plan and the devaluation of October 1991; and the year of 1992, which is basically a "shock free" period (chart 1).

Chart 1
Nominal exchange rate
(domestic currency per US$; monthly percentage change)



The introduction of the futures contracts for foreign exchange at the Futures Board of Trade reduced the impact of the exchange rate instability on trade flows by allowing market hedging strategies to be more easily implemented. However, by engaging in such operations, traders have to bear costs, essentially of two natures. On the one hand, if the market is not using publicly available information efficiently futures prices become biased predictors of future spot prices and an "information cost" may arise. On the other hand, a risk premium may be present entailing an additional cost in hedging strategies. Both of these costs would be expected to be exacerbated during periods of repeated policy changes, which should be reflected in futures prices behavior.

The paper is organized as follows. Section 2 presents background information on the institutional aspects of the futures market operations in the Commodity and Futures Exchange in Brazil. Sections 3 and 4 perform tests of unconditional and conditional unbiasedness, respectively, using the generalized method of moments. The next three sections discuss alternative explanations for the finding of unpredictable but non-zero excess returns in the market. Incomplete information by the agents is postulated in section 5, and section 6 investigates the impact on future market prices of infrequent and large domestic currency devaluations by conditioning the probability distribution of the excess return on the event of a major devaluation. Section 7 estimates a Kalman filter assuming the existence of a time varying risk premium and assesses its forecasting accuracy. Section 8 presents the concluding remarks.

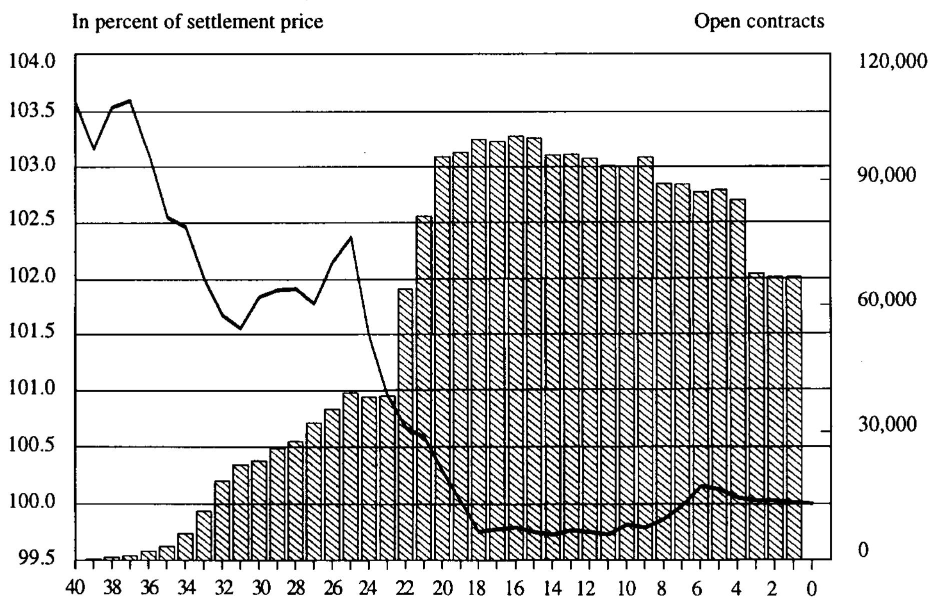
2. Institutional background

The futures market for foreign exchange was created in 1988 with the introduction of dollar contracts by the Mercantile & Futures Exchange, founded in 1986. By May 1991, as a result of the merger of this exchange with the older São Paulo Commodities Exchange (BMSP), founded in 1917, The Commodity & Futures Exchange (BM&F) was born. This exchange currently accounts for 98% of the overall Brazilian futures industry. Coincident with repeated stabilization attempts during the period of analysis and despite changes in contract specifications, there was a remarkable and consistent growth in trading volume from 1986 to 1992. Total trading on the BM&F rose from a total of about 2 million contracts in 1986 to 10 million in 1990, 19 million in 1991 and more than 40 million annually since 1992.

The main currency contract traded on the BM&F is the exchange rate of the Brazilian currency per United States dollar. The commercial US dollar, used throughout this paper, is adopted for imports, exports and financial transfers. The floating US dollar is used for other operations including gold arbitrage against the US dollar by the Central Bank. During the 1990s, the turnover for the US commercial dollar future contracts has been higher than US$ 4 billion annually, providing substantial liquidity for hedging exchange rate risk.

The contract size for the commercial US dollar is US$5,000 and every month is a delivery month. The commodity exchange trades from 10:00 am to 4:00 pm local time and the last trading day for a contract is the first business day of the delivery month. Delivery is settled cash in domestic currency in the second business day after the last day of trading. The settlement is made using the rate that prevailed at the last trading day. Chart 2 plots an illustrative contract for the dollar. Contract negotiations begin more than 40 days prior to maturity and traded volume increases substantially about 20 (business) days prior to maturity. Futures prices usually approach the price at maturity as the time to maturity reduces. The next two sections perform tests of unbiasedness using a fixed time to maturity.

Chart 2
Futures prices and volume
(illustrative dollar contract)



3. Unconditional unbiasedness

Let $F_t(T)$ be a futures price at date t of a contract maturing at date T, and S_T be the spot price at date T.[1] The difference between two future prices for the same contract, apart from each other by n periods, i.e., the measured excess return, may be due to a forecast error (ε_{t+n}) or a possible risk premium (R_{t+n}).

$$F_t(T) - F_{t+n}(T) = \{F_t(T) - E_t[F_{t+n}(T)]\} + \{E_t[F_{t+n}(T)] - F_{t+n}(T)\} = R_{t+n} + \varepsilon_{t+n}$$

Rational agents should not be expected to consistently underpredict or overpredict the future prices at maturity or at any other later date, and the forecast error over a given period should be, on average, zero. Under this assumption, any excess return in the futures market would reflect a risk premium, R_t, a compensation demanded by risk-averse investors for taking over the risk of future price changes. This section tests whether future prices are unbiased forecasts of future spot prices. Since $E_t[F_{t+1}] = E_t[F_{t+2}] = ... = E_t[F_T] = E_t[S_T]$ under the null hypothesis, the hypothesis is formulated such that by using futures prices from contracts maturing at different times the sample size can be increased substantially. The hypothesis tested is:

$$H_0: E[F_t(T) - F_{t+n}(T)] = 0$$

where $n = 1$ month or 1 week.

The data set keeps the length of the forecasting period to maturity constant and equal to either one month or one week. Excess returns were computed for the dollar for one month (chart 3) and one week horizons, using the contracts with the nearest delivery date.[2] The specification of the hypothesis is in natural logarithms. Since ε_{t+n} is generated by new information that arrives between time t and $t + n$, the residuals will be serially correlated even under the null hypothesis unless the sampling interval equals the forecasting interval and therefore $n = 1$. The method of moments as described in Hansen & Hodrick (1980) was applied to the overlapping data set (daily data with one week or one month futures rates). Estimation also uses the sampled data set with nonoverlapping observations (weekly data with one week futures rates or monthly data with one month futures rates). The sampled data set is later used for modelling the risk premium. Hansen & Hodrick (1980) provide a proof that using the overlapping data the generalized method of moments (GMM) is superior to sampling the data to produce a nonoverlapping data set to which ordinary least squares can be applied directly, but it should be noted that the one month forecast errors with daily data may produce a relatively large overlap of data.

Table 1 presents the results of the unbiasedness test for the period March 1989 to December 1992 for both sampled and non-sampled observations. The table shows the mean excess return from holding a futures contract for one week and one month and the corresponding t-statistics.

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[1] Unless explicitly stated, the variables will be expressed in terms of their natural logarithms.

[2] Consider for example the November 1991 contract. To obtain the values for November, the difference between the October prices and the September prices for the November contract were taken keeping constant the forecasting horizon (one month or one week).

Chart 3
Excess return on futures prices for the dollar
(one month forecast horizon)

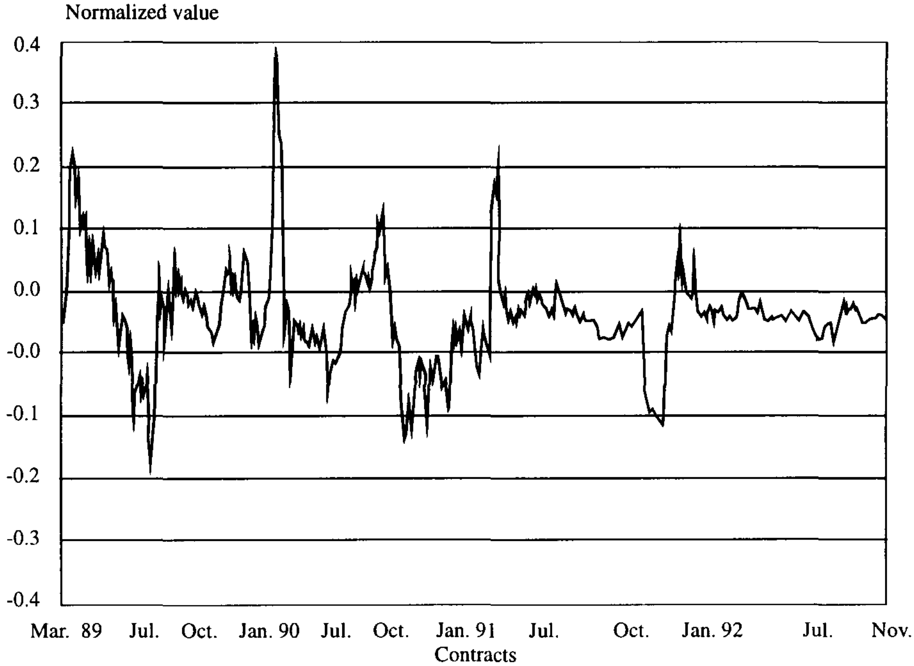


Table 1
Test of unconditional unbiasedness
Excess returns
(Complete sample: Mar. 1989 – Nov. 1992)

	Overlapping observations		Nonoverlapping observations	
Forecast horizon	1 month	1 week	1 month	1 week
Obs.	687	1,369	43	178
Mean	0.0084	0.0031	0.0189	0.0045
t-statistic	3.2706	3.3736	1.7070/1.7275	1.9220/1.9275
p-value	0.0005	0.0004	0.0480/0.0420	0.028/0.0270

The *t*-statistics use standard errors corrected for autocorrelation using the method of moments for overlapping observations. For nonoverlapping observations both the standard *t*-statistics as well as the ones computed with White's heteroskedastic consistent covariance matrix are provided.

The results in table 1 suggest that the null hypothesis of a zero bias can be rejected for the full sample. Table 2 divides the sample in three subperiods over which it is expected to display a differentiated behavior. The subperiods were defined by major macroeconomic shocks in the economy, as discussed above.

Table 2
Test of unconditional unbiasedness
Excess returns (Selected subsamples)

	Sample period: Mar. 1989 — Mar. 1990		
	Overlapping observations		*Nonoverlapping observations*
Forecast horizon	1 month	1 week	1 week
Obs.	188	388	56
Mean	0.0034	0.0105	0.0114
t-statistic	5.3148	4.4175	2.0670/2.0851
p-value	0.0000	0.0000	0.0220/0.0185
	Sample period: June 1990 — Dec. 1991		
	Overlapping observations		*Nonoverlapping observations*
Forecast horizon	1 month	1 week	1 week
Obs.	302	606	79
Mean	-0.0114	-0.0021	0.0006
t-statistic	-0.29758	-0.0517	0.1678/0.1688
p-value	0.00292	0.5206	0.434/0.4330
	Sample period: Jan. 1992 — Nov. 1992		
	Overlapping observations		*Nonoverlapping observations*
Forecast horizon	1 month	1 week	1 week
Obs.	197	375	43
Mean	0.0137	0.0037	0.0027
t-statistic	11.9170	7.0096	4.413/4.465
p-value	0.0000	0.0000	0.0000/0.0000

The t-statistics adopt standard errors corrected for autocorrelation using the method of moments for overlapping observations. For nonoverlapping observations, both the standard t-statistics as well as the ones computed with White's heteroskedastic consistent covariance matrix are provided.

The excess returns were consistently positive for each of the forecasting horizons before and after the Collor Plan. For each of these subperiods the forecasting bias is significantly different from zero. During the period June 1990 — Dec. 1991 (the "Collor Plan"), excess returns are not statistically significant. This may be due to the unusually high degree of uncertainty regarding the stabilization plan and exchange rate policy prior to, and dissipated by, the adoption of the plan. Conditional on the assumption of a zero expected forecast error, a non-zero excess return may be evidence of a non-zero (time varying) risk premium. Before attempting to further discuss this finding and to model a time varying risk premium, we perform must conditional tests of unbiasedness.

4. Conditional unbiasedness

Efficiency tests conditioned on the information available to the market and assuming a zero risk premium tests if the excess return from holding a futures contract for n periods $F_t(T) - F_{t+n}(T) = \varepsilon_t$ is correlated with information up to time t. Efficiency requires the forecasting error to be orthogonal to variables in the information set, I_t. This null hypothesis of market efficiency can be examined by testing the hypothesis that $\beta_0 = \beta_k = 0$ in the following regression:

$$F_t(T) - \mathrm{F}_{t+n}(T) = \beta_0 + X_t\beta_k + \varepsilon_{t+n}$$

where X_t is a vector of variables in the information set, I_t, and B_k is a vector of k coefficients.

Since under the null hypothesis X_t and ε_{t+n} are orthogonal, OLS will generate consistent estimates of the coefficients. However, the errors might not be homoskedastic. Second, the errors may not be uncorrelated, since the sampling interval does not necessarily equal the forecasting interval. ε_{t+n} is generated by new information that arrives between time t and time $t + n$, and the residuals are expected to be correlated even under the null. In this case, as Hansen and Hodrick (1980) show, the errors will be moving averages of order $(n - 1)$, where n is the forecast horizon. To obtain consistent estimates of the standard errors, the correct asymptotic covariance matrix is computed using the generalized method of moments proposed by Hansen (1982). The Newey-West (1987) correction to guarantee positive definiteness is applied.

Past excess returns from the futures market for the dollar were included in I_t to perform a weak form test. We basically test if investors relying on past price patterns receive any excess return.

The following equation was used for estimation with varying lags:

$$F_t(T) - F_{t+n}(T) = \beta_0 + \beta_1 [F_{t-n}(T) - F_t(T)] + ... + \beta_k [F_{t-n-k+1}(T) - F_{t-k+1}(T)] + \varepsilon_{t+n}$$

The results of the White test suggest that the null of homoscedasticity cannot be rejected at the 1% level, $\chi^2(1) = 0.8$ for monthly sampled observations and 0.9 for the weekly sampled dataset. The use of different contracts keeping constant the time to maturity of the forecasting horizon results in no change in the variance of the futures price since the amount of information gathered is presumably constant. Autocorrelation is also not detected. Alternatively, to avoid sampling, the full sample is also used and GMM estimation with White's (1984) heteroscedastic consistent covariance matrix estimator is performed (White's standard error). The overlapping data confirms the expected autocorrelation as pointed out in Hodrick (1987). The Ljung-Box statistic rejects the hypothesis of zero residual autocorrelations for several lags $\chi^2(1) = 8.17$, $\chi^2(2) = 9.20$, $\chi^2(20) = 32.21$, at the 5% level. Estimation is also performed with the autocorrelation and heteroscedastic consistent covariance matrix estimator described in Andrews (1991) and here referred as Andrews standard error. Finally, before performing the efficiency tests, tests for nonstationarity for the sampled and non-sampled observations were performed and are reported in table 3.

Table 3
Unit root tests for $F_{t+n}(T) - F_t(T)$

	Nonoverlapping		Overlapping	
Dependent variable: $\Delta(F_{t+n}(T) - F_t(T))$ Augmented Dickey-Fuller	Weekly t-test	Monthly t-test	Weekly t-test	Monthly t-test
Constant, no trend	-4.8809	-4.8293	-4.2996	-4.8809
Constant, trend	-4.9027	-4.7326	-4.2754	-4.9027

All tests rejected the existence of a unit root so standard asymptotic analysis was used to obtain the distributions of the tests statistics. Table 4 presents the point estimates of the β's and the corresponding standard deviations. Table 5 introduces the GMM estimation and the χ^2 statistic to test the null that expected returns are zero.

Table 4
Test of conditional unbiasedness
Complete sample: Mar. 1989 — Nov. 1992
Nonoverlapping observations

K = 1

Forecast horizon Obs.	Monthly 42		Weekly 177	
	β_0	β_1	β_0	β_1
Mean	0.0129	0.1193	0.0041	0.0219
s.e.	0.0112	0.1556	0.0023	0.0754
White's	0.0115	0.1284	0.0023	0.1022
p-value	0.1330	0.1790	0.0380	0.3860

K = 4

Forecast horizon	Monthly				
	β_0	β_1	β_2	β_3	β_4
Mean	0.0195	0.0824	-0.3435	-0.2228	-0.1191
s.e.	0.0120	0.1618	0.1579	0.1578	0.1615
White's	0.0118	0.1324	0.1283	0.1482	0.1115
p-value	0.0540	0.2690	0.6050	0.9290	0.8540
Forecast horizon	Weekly				
	β_0	β_1	β_2	β_3	β_4
Mean	0.0037	0.0192	-0.1003	-0.0042	-0.0037
s.e.	0.0023	0.0760	0.0761	0.0761	0.0761
White's	0.0024	0.1006	0.0652	0.0623	0.0504
p-value	0.0680	0.4240	0.0630	0.4730	0.4710

Table 5
Test of conditional unbiasedness
Complete sample: Mar. 1989 — Nov. 1992
Overlapping observations — GMM estimation
K = 1

Forecast horizon	Monthly		Weekly	
Obs.	686		1368	
$\chi^2(2)$	2503.15		370.79	
p-value	0.0000		0.0000	
	β_0	β_1	β_0	β_1
Mean	0.0005	0.9452	0.0007	0.7631
White's	0.0008	0.01933	0.0006	0.0469
Andrew's	0.0010	0.01893	0.0005	0.0408

K = 4

Forecast	Monthly					
	β_0	β_1	β_2	β_3	β_4	$\chi^2(5)$
Mean	0.0006	1.1367	-0.1764	-0.0131	-0.0168	3872.8
Andrew's	0.0009	0.0793	0.1022	0.0675	0.0308	0.0000
Forecast	**Weekly**					
	β_0	β_1	β_2	β_3	β_4	$\chi^2(5)$
Mean	0.0006	1.1367	-0.1764	-0.0131	-0.0168	3872.9
Andrew's	0.0009	0.0793	0.1022	0.0675	0.0308	0.000

Table 6
Test of conditional unbiasedness
Selected subsamples

Monthly forecast horizons — Overlapping obs. — GMM estimation

	187 obs. (Mar. 1989 — Mar. 1990)		301 obs. (June 1990 — Dec. 1991)		196 obs. (Jan. 1992 — Nov. 1992)	
$\chi^2(2)$	1033.2		1817.3		991.0	
p-value	0.0000		0.0000		0.0000	
	β_0 (s.e.)	β_1 (s.e.)	β_0 (s.e.)	β_1 (s.e.)	β_0 (s.e.)	β_1 (s.e.)
Mean	0.0020	0.9402	-0.0004	0.9470	0.0025	0.7842
White's	(0.0024)	(0.0292)	(0.0013)	(0.0299)	(0.0010)	(0.0821)
Andrew's	(0.0025)	(0.0296)	(0.0014)	(0.0222)	(0.0012)	(0.0583)

Weekly forecast horizons — Overlapping obs. — GMM estimation

	387 obs. (Mar. 1989 — Mar. 1990)		605 obs. (June 1990 — Dec. 1991)		374 obs. (Jan. 1992 — Nov. 1992)	
$\chi^2(2)$	174.0		268.3		590.6	
p-value	0.0000		0.0000		0.0000	
	β_0 (s.e.)	β_1 (s.e.)	β_0 (s.e.)	β_1 (s.e.)	β_0 (s.e.)	β_1 (s.e.)
Mean	0.0028	0.7391	-0.0004	0.7902	0.0014	0.6276
White's	(0.0017)	(0.0712)	(0.0008)	(0.0621)	(0.0005)	(0.0750)
Andrew's	(0.0016)	(0.0650)	(0.0007)	(0.0497)	(0.0006)	(0.0738)

Weekly forecast horizons — Nonoverlapping obs.

	55 obs. (Mar. 1989 — Mar. 1990)		78 obs. (June 1990 — Dec. 1991)		42 obs. (Jan. 1992 — Nov. 1992)	
	β_0 (s.e.)	β_1 (s.e.)	β_0 (s.e.)	β_1 (s.e.)	β_0 (s.e.)	β_1 (s.e.)
Mean	-0.0121	-0.0856	0.0004	0.0901	0.0026	0.0627
s.e.	(0.0054)	(0.1359)	(0.0035)	(0.1126)	(0.0008)	(0.1569)
White's	(0.0059)	(0.1336)	(0.0034)	(0.1349)	(0.0007)	(0.0890)
p-value	0.7400	0.0230	0.4730	0.2530	0.0000	0.2430

The results indicate an ambiguous evidence of the weak efficiency hypothesis. The use of overlapping observations leads to the rejection of the hypothesis for the whole period as well as for the selected subsamples. This is indicative of the *ex-ante* predictability of future forecast errors by past information when overlapping daily observations are used. However, for nonoverlapping observations, for weekly or monthly forecasting horizons, the coefficient of the lagged excess return is not significant in any of the periods.

Two lines of explanations are provided for the fact that the futures rate has been a biased predictor of subsequent dollar movements in Brazil during the sample considered. Both assume rational agents and one of them assumes some sort of market failure. The first of these approaches explores the case of a futures bias under incomplete information. Following Kaminsky and Kumar (1990), the next section provides an illustration which assumes the economy was affected by repeated policy shocks during the sample period. The second approach assumes a "peso problem" that occurs, as pointed out by Krasker (1980), when the probability of rejecting the efficiency and rationality of the market when it is true is larger than one is led to believe. An illustration of a peso problem for the Brazilian economy is provided in section 6, for the subperiod starting Jan. 1992, immediately after the last major devaluation of the domestic currency. Econometric modelling of the excess return is provided through a Kalman filter in section 7.

5. Excess return under incomplete information

Although expectations are assumed to be rational *ex-ante*, they may appear to be biased *ex-post*. If there is an informational problem, such as the lack of knowledge of the true exchange rate process, even if investors use all available information efficiently, they may make, on average, non-zero forecast errors.

Assume Brazil's exchange rate devaluation policy followed during the period of analysis a rule of devaluing the currency in each period above or below the expected inflation rate from period t to $t + n$. The futures price for foreign exchange could be written as follows:

$$F_t\,(T+n) - F_t\,(T) = \pi^e + E_t\delta_i + u_t$$

where, π^e is the expected inflation rate for n periods, and $i = 1{,}2$, such that the exchange rate policy pursues a devaluation rate either above ($\delta_i = \delta_1 > 0$) or below the inflation rate ($\delta_i = \delta_0 < 0$).

Assume that investors know that the monetary authority can change its policy but do not know with certainty when a policy change may occur. Denote θ_t the probability of switching to a policy where δ_i is positive in period t. The expected devaluation in the future exchange rate, using all available information up to period $t-1$, will be

$$E_{t-1}\,[F_t\,(T+n) - F_t\,(T)] = \pi^e + \delta_1\,\theta_t + \delta_0\,(1-\theta_t)$$

If a rate of devaluation above expected inflation is not pursued for several periods then the futures rate could be written as

$$F_t(T+n) = F_t(T) + \pi^e + \delta_0 + v_t$$

and, therefore, the rational overprediction (agents expected a higher rate of devaluation) will be given by the forecast error:

$$E_{t-1}[F_t(T+n)] - F_t(T+n) = (\delta_1 - \delta_0)\,\theta_t - v_t$$

which is on average different from zero, even though expectations were completely rational *ex-ante*. If the central bank alternates periods of faster and slower rates of devaluation, agents with incomplete information regarding the government policy may make mistakes and persistence may arise as a result of a cost of adjustment on information gathering. The significantly different from zero excess return may at times be persistently negative or positive respectively during subsamples. Over longer sample periods, however, rational agents would be expected to learn the government behavior, and no significant excess return should be observed. Under those assumptions, both the stabilization attempts and the medium sized devaluation during the sample period could explain the significance of the unbiasedness tests for most of the sample period. The significance of the excess return for the period starting after the last major devaluation of October 1991, however, does not seem to be explained by any major unanticipated change of government policy. This question is addressed in section 7, when we discuss the "peso problem."

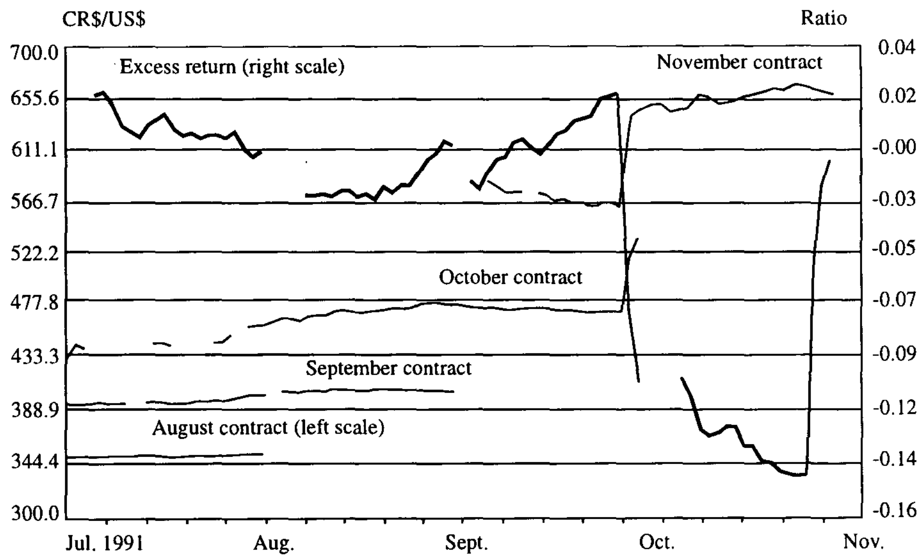
6. Conditioning the probability distribution on the event of a major devaluation

Although no major shock occurred during 1992, after the last major devaluation of October 1991 the excess return during this period is significantly different from zero for both sampled and non-sampled observations. As Hodrick (1987) puts it, whenever there are potential changes in government policy processes that have not occurred in the sample the data are not ergodic. Ergodicity is also a problem if there are events that occur during the sample but not with the appropriate frequency to correspond to their *a priori* probability. Krasker (1980) proposed that if the sampling distribution is difficult to determine, the use of the statistics asymptotic distribution may lead to misleading results. The classical inference problem is the case of the Mexican peso which was consistently sold at a discount on the forward market (Lizondo, 1983). Participants were anticipating the possibility of a sizeable devaluation but, in moderate sized samples, the event did not occur, making it seem as though expectations were biased. This is indeed a description of how a type *I* error can occur. Krasker points out that "it may be argued however that the efficiency tests in these cases are invalid because the sampling distribution of the test statistic may not be well approximated by their asymptotic distribution under the null" (e.g., the sampling distribution may not be normal). The distribution of excess returns may have heavy tails and may be skewed if the participants in the market perceive a small probability, in each period, that there would be a major devaluation or a stabilization attempt which would end the inflationary process causing the excess return as defined so far to be unusually positive in the first case and negative in the second.[3]

[3] According to the definitions used so far, if futures prices have a rising trend, the risk premium is said to be negative (as if agents underpredicted futures prices), and those who are short in domestic currency and long in US$ are compensated for the risk they bear, such as the risk that a stabilization will be attempted and the currency devaluation will be much lower than the inflation rate.

Instead of assuming that the distribution of the excess return is approximately normal, we will assume, as in Krasker (1980), that the conditional distribution of the excess return, given that no major devaluation occurs during the sample period, is approximately normal. The test statistic to deal with the peso problem was developed to assess the significance of non-zero returns for the mark/pound during the German hyperinflation. We modify the test for excess returns and apply it to assess the existence of a peso problem with respect to the probability of a major domestic currency devaluation. The sample contains 44 weekly observations of the last subperiod, from January 1992 to November 1992. If there had been a major devaluation during the sample period at time t, it is assumed that the cruzeiro/dollar rate would have increased by a factor proportional to a cumulative index of the real exchange rate with the benchmark base period set as the beginning of the sample, since the sample itself starts after a major "correction" of the exchange rate. Such a devaluation would imply a drop in F_t/F_{t+n} by about the same factor and if the probability of such an event is not considered, simply because the event did not happen during the sample period, agents would probably end up underpredicting F_{t+n} at some point in time, as they did in October 1991 (chart 4).

Chart 4
Futures prices and "unanticipated" exchange rate develuation



Let A_t ($\breve{A}t$) denote the event that a major devaluation does not (does) occur in period t so that $A = \cap_{t=1}^{t} A_t$ is the event that a major devaluation does not occur during the sample period. We assume that $Pr\{A_t / A_1, A_2, \ldots, A_{t-1}\} = \theta_t = e^{-(\alpha + \beta t)}$ for some $\alpha, \beta > 0$. Define δ_t as the size of the possible "corrective" devaluation based on the real dollar exchange rate index that begins with the sample itself. Then, the expected value defined for the above probability and conditional on the event of a major devaluation is:

$$E\,[F_t(T) - F_{t+n}(T) / \{A_1, A_2, \ldots, A_{t-1}, \breve{A}_t\}] =$$

$$\delta_t\, E\,[F_t(T) - F_{t+n}(T) / \{A_1, A_2, \ldots, A_{t-1}, A_t\}]$$

Therefore,

$$E[F_t(T) - F_{t+n}(T) / A] = \frac{1}{T} \sum_{t=1}^{T} \frac{1}{e^{-(\alpha+\beta t)} + (1 - e^{-(\alpha+\beta t)})\, \delta_t} - 1$$

Let $\hat{r}_t$ and $\hat{\sigma}_t$ be the observed value of $F_{t-n}(T) - F_t(T)$ and standard deviation during the sample period. The test statistic is $r\chi$ where χ is the characteristic function of the set A. $r\chi = r$ if A occurs, and 0 otherwise.We reject the null, $E[r_t] = 0$, if $1 - Pr_{\alpha,\beta}\,\{r\chi \leq \hat{r}_t\chi\} \leq 0.05$, i.e., if the observed value lies too far in the tail of the distribution of $r\chi$. Assuming the conditional distribution of r_t is approximate normal,

$$Pr_{\alpha,\beta}\;(r\chi \leq \hat{r}\chi) \simeq \phi\;\left(\frac{\hat{r}\chi - E_{\alpha,\beta}[r / A]}{\hat{\sigma}}\right)\; e^{-\alpha T - \beta T(T+1)/2}$$

where Φ is the cumulative distribution function for $N(0,1)$.

An illustration of the fact that the rejection of the null under the asymptotic distribution does not necessarily need to lead to the case where only the conditional distribution of r is assumed to be normal is provided in table 7. The table plots, for different sample periods of the same subperiod from January 1992 to November 1992, the results of the standard tests for nonoverlapping weekly observations under the column *Normal distribution* and the respective p-values, and the results of the test when the conditional distribution is assumed to be normal with the respective p-values.

Table 7
Rejection of the null under conditional normality
Excess returns — sample: Jan. 1992 — Nov. 1992

Nonoverlapping weekly observations

Obs.	δ_t	r_t	σ_t	Asymptotic test statistic	Normal p-value	Conditional prob. θ_t of devaluation	Normal p-value	H_0 $E[r_t] = 0$
10	1.008	0.0031	0.0048	2.089	0.033	0.090	0.024	reject
15	1.056	0.0032	0.0040	3.087	0.004	0.150	0.013	reject
20	1.073	0.0027	0.0036	3.377	0.002	0.200	0.021	reject
25	1.040	0.0025	0.0036	3.488	0.001	0.249	0.032	reject
30	0.999	0.0027	0.0043	3.414	0.001	0.299	0.046	reject
35	0.969	0.0027	0.0041	3.837	0.000	0.349	0.061	no
40	0.939	0.0028	0.0039	4.494	0.000	0.399	0.077	no
44	0.879	0.0027	0.0040	4.466	0.000	0.439	0.090	no

Imposing a type I error probability of 0.05, we can see that while we reject the null in every subsample, reasonable values of α and β do not lead to the rejection of the null for all subsamples. When the probability of a major devaluation of a magnitude of 4% reaches 0.35 we do not reject the null. It is rational for the participants in the market to "overpredict" the futures rate anticipating the occurrence of a possible devaluation of the cuzeiro.

7. Kalman filter modelling of excess returns

This section develops a Kalman filter approach to model the weekly forecast error observed in the nonoverlapping sample. Assume

$$\frac{F_{t,k} - S_t}{S_t} - \frac{E_t\,(F_{t+1,k-1} - S_t)}{S_t} - p_t + \varepsilon_t$$

where:

S_t = spot exchange rate;

$F_{t,k}$ = rate at t of a futures contract with k days to maturity;

$F_{t+1,k-1}$ = rate at $t + 1$ of a futures contract with $k - 1$ days to maturity;

p_t = normalized risk premium.

Hodrick and Srivastava's (1986) version of the generalized method of moment tests the implications of the above equation for all t. Let

$$p_t = \alpha\, p_{t-1} + \eta_t$$

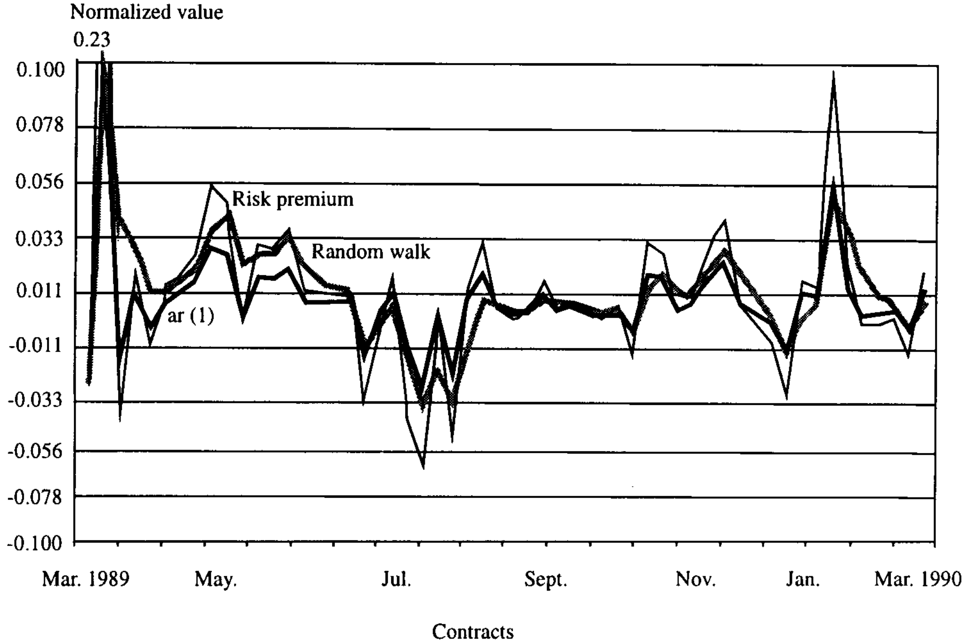
where p_t is the risk premium, ε_{t+1} is the innovation in the logarithm of the exchange rate, S_{t+1}, and η_t is the innovation in the risk premium. Following McCurdy and Morgan (1986), the estimated Kalman filter uses a logarithmic transformation of the data in order to avoid dividing the equation by a common stochastic variable. Following Wolff (1985), this section postulates this state space model, uses a Kalman filter approach to estimate the unobservable risk premiums and employs a Theil's U statistic (the ratio of the root mean-square error of the forecast to the root mean square error of the naive forecast of no parameter change) to evaluate the ability of the Kalman filtering experiment to improve the forecast of the futures rate.

The estimation used the sample of nonoverlapping observations as well as the full sample with a weekly forecasting horizon. The first sample contains 178 weekly observations, from March 1989 to November 1992. The full sample contains 1,297 weekly observations. Ordinary least squares were fitted to the fixed parameter regression model, and by maximum likelihood the first order autoregressive stochastic parameter model and the random walk model were estimated to the sampled observations as discussed in Harvey (1980) and seen on chart 5.[4]

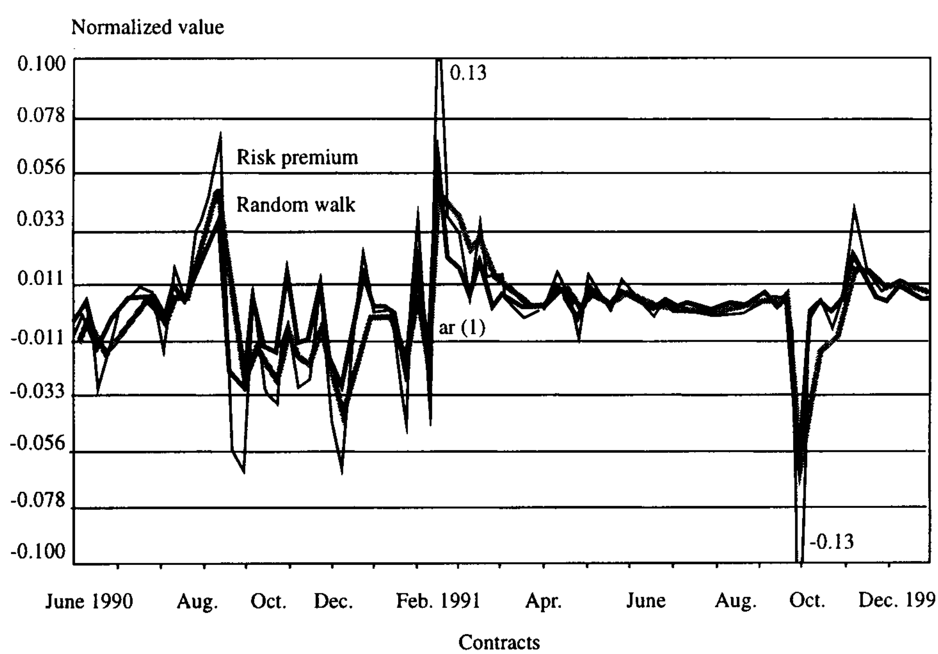
[4] The vector of prior coefficients is calculated from a regression in the initial observations of the sample.

Chart 5
Excess return — Kalman estimation
Weekly forecasting horizon

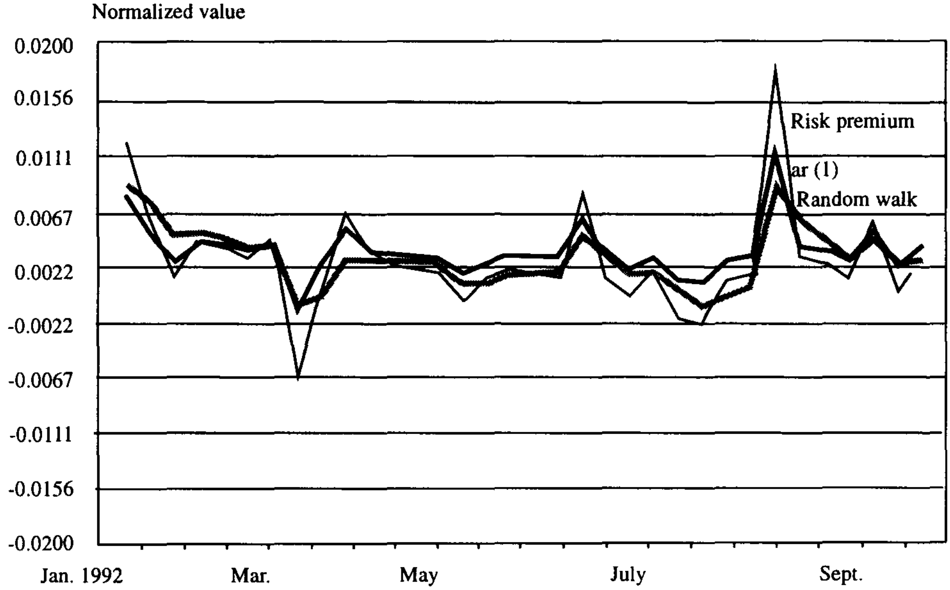
Mar. 1989 — Mar. 1990



June 1990 — Dec. 1991



Jan. 1992 — Nov. 1992



• *Full sample*[5]

Fixed parameter

$$F_{t+1}(T) - F_t(T) = \underset{(0.0024)}{0.0041} + \varepsilon_{t+1} \quad ; \qquad s_e^2 = 0.0009$$

Durbin-Watson statistic = 1.95156
Log of likelihood function = 351.12

Random walk (naive) parameter

$$F_{t+1}(T) - F_t(T) = p_t + \varepsilon_{t+1} \quad ; \qquad s_e^2 = 0.0008$$

$$p_t = p_{t-1} + \eta_t \quad ; \qquad s_\eta^2 = 0.38\ s_e^2$$

Final p_t = 0.0034
(0.0127)

Log of likelihood function = 342.57

[5] Figures in parentheses beneath coefficient estimates are the corresponding estimated standard errors, and s_e^2 is the estimate of the error variance σ_e^2.

Autoregressive stochastic parameter

$$F_{t+1}(T) - F_t(T) = p_t + \varepsilon_{t+1} \quad ; \quad s_e^2 = 0.0005$$

$$\underset{(0.0024)}{p_t - 0.0043} = \underset{(0.0767)}{0.0218}\ (p_{t-1} - \underset{(0.0024)}{0.0043}) + \eta_t \quad ; \quad s_\eta^2 = 0.0005$$

Final $p_t = \underset{(0.0153)}{0.0052}$

Log of likelihood function = 348.93

• *Full sample: overlaping observations*

The autoregressive stochastic parameter model when sampling of the data is avoided results in:

$$F_{t+1}(T) - F_t(T) = p_t + \varepsilon_{t+1} \quad ; \quad s_e^2 = 0.0004$$

$$\underset{(0.0026)}{p_t - 0.0031} = \underset{(0.0179)}{0.7630}\ (p_{t-1} - \underset{(0.0026)}{0.0031}) + \eta_t \quad ; \quad s_\eta^2 = 0.0002$$

Final $p_t = \underset{(0.0126)}{0.0055}$

Log of likelihood function = 2904.9

The point estimates of the fixed parameter model imply that the original weekly risk premium is estimated to be 0.41% throughout the period. The autoregressive stochastic parameter model implies that the risk premium varies around a mean of 0.43, according to a first order autoregressive model with coefficient 0.02. The same model, when sampling of the observations is avoided increasing the degrees of freedom and reducing the forecast interval, implies a mean of 0.31 and a first order coefficient of 0.76. Therefore, as it should be expected, the correlation between risk premiums in adjacent weeks depends greatly if we consider overlapping observations or not. Finally, the random walk model estimates a risk premium of 0.34 by the end of the sample period. Chart 5 plots the risk premium estimates of each of the models *versus* the actually observed excess return for three subsamples.

The forecasting ability of the stochastic coefficient models was compared to the fixed parameter model by means of the U-statistic as defined in Theil (1966). In the case of the full sample with overlapping observations, the inequality coefficient was compared to the U-statistic obtained with the naive forecast represented by the random walk. The autoregressive model seems capable of outperforming the fixed parameter model or the naive model (random walk) for specific forecasting horizons. Defining the variance of the error of the transition equation for the naive model as $\sigma_\eta^2 = Q\,\sigma_e^2$, the Kalman filter was applied varying the ratio of the variance of η_t to the variance of ε_{t+1} from 10^{10} to 0 with "step" reductions of 0.1. The search over the variance of the transition equation resulted in $Q = 0.38$, the greatest Q to produce an inequality coefficient lower than the one under the fixed parameter model. Theil's coefficient was calculated as 0.9281 for such Q. The innovation variance in futures prices according to this model is 2.6 times as variable as the innovations in risk premiums implying that almost 40% of the variability of the futures rate forecast error are due to risk premiums. The fixed parameter model resulted in a coefficient equal to 0.9288. The inequality coefficient obtained for the stochastic parameter model with sampled observations was 0.9208, inferior to the one obtained with either the fixed parameter model or with the naive forecast of the random walk model. Alternatively, with overlapping data, the *AR* parameters result in a Theil's coefficient of 0.8876.

8. Concluding remarks

The hypothesis of unconditional unbiasedness was shown to be rejected for the futures rates for the dollar in the São Paulo Board of Trade for the period 1989-92. Weekly and monthly samples and various subperiods of the March 1989 to November 1992 period were used. This result is robust to the case of overlapping observations according to the GMM estimation performed.

The results of testing the hypothesis of conditional unbiasedness give weak evidence of predictability of excess returns. The use of lagged forecast errors fails to significantly predict the behavior of excess returns for the nonoverlapping observations. The fact that sampling may discard important short lived periods of turbulence and information and the observation that most of the forecast errors were caused by unpredictable action by the government justify that finding. The the generalized method of moments allows the use of overlapping data, avoiding the severe loss of information involved in discarding enough data to exactly match the maturity time of the contract with the sampling interval. It also avoids the inconsistency of GLS when the error term follows a moving average and lagged dependent variables are included as instruments of the prediction equation for the forecast error.

The first explanation provided for the fact that the futures rate has been a biased predictor of subsequent dollar movements in Brazil during the sample considered assumes an informational problem. The frequent changes in exchange rate policy could potentially account for persistent forecast errors by the participants in the futures market for foreign exchange. A second line of explanation explores the occurrence of a "peso problem", where the probability of rejecting the efficiency and rationality of the market when it is true is larger than one is led to believe. The subperiod starting on January 1992, immediately after the last major domestic currency devaluation of the Cruzeiro in October 1991 is used to show that no rejection of the hypothesis occurs if the probability distribution of the excess return is conditioned on the possible event of a major devaluation. In this case, rational overprediction by the part of the agents may be a natural outcome. Modelling of the excess returns was provided by means of an application of the Kalman filter that is showed to improve the forecasting ability over a specific sample period.

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