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The Random Walk Hypothesis and the Behavior of Foreign Capital Portfolio Flows: the Brazilian Stock Market Case*

Benjamin Miranda Tabak**

Abstract

In this paper the random walk hypothesis is tested for a set of daily Brazilian stock data given by the São Paulo Stock Exchange Index (IBOVESPA) in the period of 1986-1998. A rolling variance ratio test for different investment horizons was conducted, and it is concluded that prior to 1994 the random walk hypothesis is rejected but after that it cannot be rejected. Institutionally maturing markets, increasing liquidity and the openness of Brazilian markets for international capital can explain this increase of efficiency of the Brazilian stock market. An error-correction model is used to explain the relationship between the IBOVESPA and foreign portfolio inflows. Evidence suggests that the release of foreign capital control is one of the main determinants of increased efficiency in the Brazilian equity market.

JEL: G14, G15.

Keywords: Random Walk, Emerging Markets, Efficiency, Portfolio Inflows.

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1. Introduction

A lot of work has been done by researchers on testing the random walk hypothesis (RWH) for financial price series. Rejection of this hypothesis has serious implications for investors as they can earn profits from forecasting future prices. The rejection of the RWH suggests that stock returns contain predictable components.

Lo and Mackinlay (1988) found that stock returns do not follow random walks for the US markets using a variance ratio test. Poterba and Summers (1988) suggest that the values of variance ratios give evidence of negative autocorrelation (mean reversion) at long investment horizons and positive autocorrelations at short horizons[1].

For international markets, Frennberg and Hansson (1993) used the variance ratio test on the Swedish market and found evidence of positive auto correlated returns for short investment horizons, one to twelve months, and for longer horizons, two years or more, they found indications of negative autocorrelation, in line with research on the U.S. stock market.

Shastri and Shastri (1994) analyze stocks listed in the Tokyo Stock Exchange and found evidence of deviations from the random walk for small-sized stocks using the variance ratio test. They could not reject the RWH for medium sized and larger stocks, though.

Ayadi and Pyun (1994) show that for daily data the RWH could be rejected for the Korean Stock Exchange assuming that errors are homoscedastic. However, with heteroscedastic error terms, the RWH is rejected.

Lee, Gleason and Mathur (2000) examined the French derivatives market to assess whether financial contracts were efficient. They found evidence that the RWH cannot be rejected for these contracts.

The variance ratio test has also been used to assess if foreign exchange rates followed a random walk. Pyun, Ayadi and Chu (1994) examined three major currencies and found evidence rejecting the RWH. For more on this literature the reader is referred to Liu and He (1991), Bahmani-Oskooee (1998), Choi (1999) and the references therein.

[1] For an interesting discussion on market efficiency see Fama (1970,1991).

Some findings suggest that stock market prices and exchange rates may contain predictable components and there may be significant returns to active management while others suggest that markets may be efficient. Results in the literature are mixed.

Two recent papers focused on the Brazilian stock market (and other emerging stock markets as well). The first one, is a paper from Grieb and Reyes (1999) which used the variance ratio test in order to assess if Latin American Indexes followed a random walk. Using weekly stock returns they found that Brazil indexes show a greater tendency toward a random walk if compared to Mexico. However, the results for individual firms suggest mean reversion.

Karamera, Ojah and Cole (1999) used the Chow and Denning (1993) multiple variance ratio tests to assess if emerging equity markets were efficient. They found evidence suggesting that the RWH is consistent with the dynamics of returns in most of the emerging markets. They use monthly stock price indices and found evidence suggesting that the Brazilian stock market (if we use dollar based data) for the period from December 1987 to May 1997 followed a random walk.

The purpose of this paper is to contribute to the literature on testing the RWH by examining the RWH for the Brazilian stock market, using daily data (using a higher frequency than that of Karamera, Ojah and Cole (1999) and Grieb and Reyes (1999). This is one of the most important Latin American markets but the literature on this market is almost nonexistent. An important contribution is that the RWH is tested performing rolling variance ratio tests and tests suggest that the Brazilian equity market has become increasingly more efficient. The empirical evidence suggests that there is a structural break in the Brazilian equity market which is located around 1994.

The effects of the Brazilian foreign portfolio inflows is examined and it is found that portfolio inflows and the Brazilian stock market cointegrate, sharing a long-run relationship. Furthermore, running a vector-error correction model evidence suggests that there is bicausality between the IBOVESPA and foreign portfolio inflows, which shed some light in the previous findings that found an increase in the efficiency of the equity market.

The next section describes the data that is used for testing the RWH and motivates the use of the Lo and Mackinlay (1988, 1989) and Chow and Denning (1993) variance ratio tests. Section 3 presents the main ideas behind the variance ratio test. In section 4 results are shown. In section 5 some cointegration and causality tests are done in order to help explain results found in section 4. Section 6 concludes the paper.

2. The DATA

This paper uses the BOVESPA index to test the RWH for the Brazilian stock market[2]. The BOVESPA index (or IBOVESPA) is the most important indicator of the stock market in Brazil, representing the average behavior of prices of the main stocks at the São Paulo stock exchange. It represents approximately 85% of the total trading in all the stock exchanges in Brazil. The index is the current value of a portfolio starting from a hypothetical investment and it serves as an average indicator of the behavior of the market.

The IBOVESPA is calculated by the following formula:

$$IBOV_t = \sum_{i=1}^{n} P_{i,t} Q_{i,t} \qquad (1)$$

where

$IBOV_t$ =BOVESPA Index at instant t

n = total number of stocks that compose the theoretical portfolio

$P_{i,t}$ = last price of the stock i at instant t

$Q_{i,t}$ = theoretical quantity of the stock i on the portfolio at instant t.

The criteria for inclusion of stocks in the Index is that stocks fulfill, simultaneously, the following conditions, always in relation to the preceding 12 months:

[2] All financial data were taken from the Economatica datasource.

it must be included in a list of stocks resultant from the addition, in decreasing order, of the negotiability indexes up to 80% of the value of the addition of all the individual indexes;

its participation, in terms of value traded, must be superior to 0.1% of the total;

it must have been traded in more than 80% of the total trading session of the period.

The stock negotiability index is given by:

$$\sqrt{\frac{n}{N} \cdot \frac{v}{V}} \tag{2}$$

where

n = number of trades with the stock, carried out on the cash market (round lot) in the last 12 moths

N = total number of trades on the cash market (round lot) in the last 12 months

v = value in Brazilian currency obtained with trades with the stock on the cash market (round lot) in the last 12 months

V = value in Brazilian currency of the total amount traded on the cash market (round lot) in the last 12 months

According to the formula given above, BOVESPA calculates the index of negotiability of each of the stocks traded on the Exchange in the last twelve months and chooses the stocks until the total number of accumulated participation reaches 80 %. Then the other two criteria should be met.

In this paper we use daily market index data from 1986 to 1998. All prices are corrected for subscriptions, dividend and bonuses. Closing prices were deflated by the daily exchange rate of Brazilian reais per US dollars using the commercial exchange rate.

The real returns on the BOVESPA index do not pass the normality test. For the period in consideration the departure from normality is indicated by many statistics such as the

Kolmogorov-Smirnov and Jarque-Bera statistics[3]. This fact motivates the use of variance ratio tests to infer if the IBOVESPA follows a random walk, as these are nonparametric tests and do not depend on the assumption of normality of returns.

A rejection of the RWH for this index would be of much interest. This paper explores if similar patterns as those identified for developed countries apply to the Brazilian stock market. However, if the rejection is due to heteroscedasticity, which is present in almost all financial time series, then it would be interesting to test consistency of results applying heteroscedastic-robust statistics.

The Box and Jenkins (1976) methodology tests was implemented to model the returns of the Brazilian stock benchmark index and a "good model" for the IBOVESPA's return is an ARMA (1,1)-GARCH (1,1) for the 1986-1998 period. The Akaike's information criteria has been used to choose this model. This motivates the use of the heteroscedastic robust variance ratio statistics for testing the RWH[4]. Therefore, it is used the Lo-MacKinlay's univariate variance ratio test as it incorporates heteroscedasticity and Chow and Denning (1993) multivariate variance ratios. Using these statistics has become a commonplace in the financial literature, as they are more powerful than Dickey and Fuller unit root tests and Box and Pierce portmanteau tests.

3. The variance ratio test

The variance ratio test was used to check the plausibility of the log of IBOVESPA following a random walk. Let P_t be the log of price, μ a constant drift parameter and ε_t a random disturbance, white noise with normal distribution. Let P_t be a stochastic process satisfying:

$$P_t = \mu + P_{t-1} + \varepsilon_t \text{ , with } E[\varepsilon_t] = 0, \text{ for all } t, \tag{3}$$

or

$$r_t = P_t - P_{t-1} = \mu + \varepsilon_t \text{ ,} \tag{4}$$

[3] These results can be obtained from the authors.
[4] See Hamilton (1995).

where r_t is the return of one period.

Lo and Mackinlay (1989) exploit the fact that the variance of the increments in a random walk is linear in the sampling interval[5]. If a series follows a random walk the variance of its q-differences would be q times the variance of its first differences. That is

$$\frac{1}{q}\frac{Var\left(P_t - P_{t-q}\right)}{Var\left(P_t - P_{t-1}\right)} = 1 \qquad (5)$$

To accept the RWH this ratio should be statistically indistinguishable from one. Let the data consist of nq+1 observations, $P_0, P_1, \ldots\ldots P_{nq}$, where both n and q are arbitrary integers greater than one, then the estimators for μ and σ^2 are:

$$\hat{\mu} = \frac{1}{nq}\sum_{k=1}^{nq}\left[P_k - P_{k-1}\right] = \frac{1}{nq}\left[P_{nq} - P_0\right] \qquad (6)$$

$$\hat{\sigma}_a^2 = \frac{1}{nq-1}\sum_{k=1}^{nq}\left(P_k - P_{k-1} - \hat{\mu}\right)^2 \qquad (7)$$

The estimator $\hat{\sigma}_a^2$ is simply the unbiased sample variance of the first-difference of P_t. The unbiased estimator of the variance of the qth differences is:

$$\hat{\sigma}_b^2 = \frac{1}{m}\sum_{k=q}^{nq}\left(P_k - P_{k-q} - q\hat{\mu}\right)^2 \qquad (8)$$

with

$$m \equiv q\left(nq - q + 1\right)\left(1 - \left(\frac{q}{nq}\right)\right) \qquad (9)$$

If the process follows a random walk then

$$M(q) = \frac{\hat{\sigma}_b^2(q)}{\hat{\sigma}_a^2} - 1 \qquad (10)$$

[5] See also Campbell (1991), Campbell, Lo and MacKinlay (1997), Campbell and Mankiw (1987), Cochrane (1988), Faust (1992), Richardson (1993) and Richardson and Stock (1989).

should be close to zero. Then the standard homoscedastic Z_1 statistics is given by:

$$Z_1(q) = \sqrt{nq}\,M(q)\left(\frac{2(2q-1)(q-1)}{3q}\right)^{-1/2} \tag{11}$$

which has an asymptotically standard normal distribution. Let

$$\hat{V}(q) = \sum_{j=1}^{q-1}\left[\frac{2(q-j)}{q}\right]^2 \hat{\delta}(j) \tag{12}$$

and

$$\hat{\delta}(j) = \frac{\sum_{k=j+1}^{nq}\left(P_k - P_{k-1} - \hat{\mu}\right)^2 \cdot \left(P_{k-j} - P_{k-j-1} - \hat{\mu}\right)^2}{\left[\sum_{k=1}^{nq}\left(P_k - P_{k-1} - \hat{\mu}\right)^2\right]^2} \tag{13}$$

The heteroscedasticity-consistent standard normal test-statistics $Z_2(q)$ is:

$$Z_2(q) \equiv \sqrt{nq}\,M(q)\hat{V}^{-1/2}(q) \tag{14}$$

which is also asymptotically normal with zero mean and unit variance.

The null hypothesis of a random walk process can be tested by computing the standardized statistics, which are asymptotically standard normal.

Since the RWH requires that variance ratios (VR) for all investment horizons (q) be equal to one, Chow and Denning (1993) developed a multiple variance ratio test, which is similar to an F-test. The null is given by $H_{0_i} : M(q_i) = 0$ for $i = 1,2,...m$ and the alternative is given by $H_{A_i} : M(q_i) \neq 0$ any i. Any rejection of H_{0_i} will lead to the rejection of the RWH. The appropriate statistics are given by:

$$Z_1^*(q) = \max_{1\le i\le m}\left|Z_1(q_i)\right| \tag{15}$$

and

$$Z_2^*(q) = \max_{1 \le i \le m} \left| Z_2(q_i) \right| \tag{16}$$

Chow and Denning (1993) use the Studentized Maximum Modulus (SMM) distribution, which has a critical value of 2,491 for the 5 percent level of significance, to test the RWH.

4. Results

The Z_1-statistics for various investment horizons (2, 4, 8, 16, 32, 48 and 64 days) using daily data in a 3051-days time span from 1986 to 1998 were calculated for the BOVESPA index. The results are reported in table 2 and they show that the RWH is rejected for all investment horizons with 95% of confidence.

The test for the Z_2-statistics was also conducted. The results reported in table 1 show that the RWH is rejected for all investment horizons with 95% of confidence. This would lead to the conclusion that the Brazilian stock market could be inefficient for all investment horizons up to 64 days. Only in the heteroscedastic case for a 64 days investment horizon the RWH is not rejected when using the multivariate variance ratio critical level.

Our results contrast with those found in Grieb and Reyes (1999) and Karamera, Ojah and Cole (1999) which used weekly and monthly data and found evidence suggesting that one could not reject that the Brazilian stock market followed a random walk. A possible explanation for this contrasting result is that there could be nonsynchronous trading effects that are more pronounced on daily data than on weekly/monthly data. As we are using closing prices that are last transactions in one business day, these prices may not occur at the same time due to infrequent trading and this could induce spurious autocorrelation. This motivates further tests in order to enlighten our understanding of the empirical results found.

To check the robustness of this test a rolling variance ratio test calculating the Z_1 and Z_2-statistics, using a fixed window of 1024 days, was computed. It can be seen that for all investment horizons based on the more recent data the RWH cannot be rejected. This can be seen in figures 1 and 2. Even if we allow for the multivariate variance ratio critical levels there seems to be evidence of a structural break in the data. This could be explained by the fact that as markets become institutionally more mature and more liquid, these equity markets should become more efficient and returns could approach a random walk[6].

Evidence was found indicating that prior to 1994 the RWH should be rejected but after that it may not be rejected. The data were split into two sub-periods, one from 1986 to 1994 and the other from 1994 to 1998. In the first sub-sample the RWH is rejected for all investment horizons for both homoscedastic and heteroscedastic statistics at the conventional five percent level of significance as can be seen in table 2. In this particular case, even when allowing for a more general test such as the multivariate variance ratio test, the RWH is rejected for all horizons even when allowing for heteroscedasticity.

For the second sub-sample the RWH is rejected only with the homoscedastic statistics for a very short horizon of 2 days. However, this rejection is not supported by the heteroscedastic statistics. Table 3 shows that when more recent data is used the RWH cannot be rejected for all investment horizons. Finally, if we use Chow and Denning's (1993) test we see that the RWH is not rejected in all cases.

Leal *et alli* (1998) examined the impact of the listing of ADRs on the risk and return of underlying Brazilian stocks and found evidence of a reduction in the volatility of the underlying stocks after the beginning of the ADR trading. The Brazilian ADR program was initiated in 1992, but most companies issued ADR's from 1994 onwards. These results seem to be in line with our findings, suggesting that there could be a structural break in the Brazilian equity market around 1994.

Additional tests for RWH were performed using weekly real returns for the period 1986-1998. Evidence suggests that for short investment horizons the RWH is rejected. With the heteroscedastic version the RWH can only be rejected for 2 and 4 weeks as can be seen in table 4. However, if we use the multivariate version of the variance ratio we see that the rejection previously mentioned is due to inferential error and is not

[6] It is interesting to notice that if returns are assumed to follow a random walk then the market is said to be efficient but if we reject the RWH we cannot assert that the market is inefficient.

maintained if we use the 2.491 critical levels. These results are in line with the findings of in Grieb and Reyes (1999) and Karamera, Ojah and Cole (1999).

As can be seen in figure 3 the variance ratios exceed one and this is indicative of positive serial correlation and does not give rise to mean reversion. The long-run mean reversion that previous US studies (Poterba and Summers (1988) and Fama and French (1988)) have found has not been found in the Brazilian case[7].

The variance ratios can thus be used to affirm that the Brazilian equity market has become more efficient over time. It would be interesting to check the plausibility of these results using other portfolios (or even stocks) than the IBOVESPA. We will not pursue this study here, instead we focus on explaining why the stock market could be increasingly efficient.

The rejection of the RWH in this case may be explained by the low liquidity that many stocks in the index have[8]. Some of these stocks are not negotiated and even if the RWH is not rejected it would be doubtful that managers could earn abnormal profits on these stocks. To study more in depth this issue we analyze whether the increase of foreign portfolio inflows in the 1990's could help explain the increase in efficiency.

5. Foreign inflows and outflows

Access to the capital markets of many countries, particularly emerging markets, has been severely restricted for nonresidents for long periods. The foreign inflow and outflow data for the stock market is used to analyze if it could have had an influence on market efficiency in Brazil. Figure 4 shows daily traded volume in stocks in the BOVESPA index. An explanation for the results found could be the fact that liquidity had a huge increase after 1994. This could be due to an increase in foreign portfolio inflows. The variance ratio (VR) estimates point to a structural break in the mid of the 1990's which could be explained by increasing liquidity due to huge foreign portfolio inflows.

[7] The variance ratios are less than unity for horizons greater than two years, however they are not significant.
[8] Although, as we have explained before, the index is built using the most negotiated stocks.

A test for equality of means is performed for the period of 1990 to 1994 and for the period of 1994 to 1998. The mean inflow for the first period was US$ 420.08 million and for the latter period was US$ 2,530.02 million. As it can be seen in table 5 the null of equality of means is rejected with 99 % confidence. The same occurs to outflows, the mean is around US$ 231.83 million for the first period and US$ 2,061.68 million for the latter one.

In the nineties Brazil witnessed a revival of capital inflows (as many developing countries did). The reader is referred to Garcia and Barcinski (1998) for a description and analysis of the effects on macroeconomic variables of the foreign capital flows to Brazil in that time period.

In figures 5 and 6 portfolio inflows and outflows are depicted. Inflows and outflows had a huge increase in the last years and this fact could have add efficiency to the Brazilian market. One would still want to know whether there are some relationships between the stock market index and these foreign portfolio inflows-outflows.

It would be interesting to answer if there are any relationships between the BOVESPA index and foreign portfolio inflows. However, the series must be stationary in order for statistics to be meaningful. Unit root tests (both the Dickey and Fuller (1979) and Phillips and Perron (1988)) were done on all series and evidence suggesting that the series are not stationary on their levels but are stationary on the first difference (they are integrated of order 1) was found.

Based on these results a cointegration test is performed in order to verify if these variables have a long-run relationship. We use the Johansen's (1991,1995) methodology to test for cointegration between these series. The null of at least one cointegrating equation between the real level of the BOVESPA index and foreign portfolio inflows could not be rejected, as can be seen in table 7. We used the multivariate Schwarz information criterion to choose the lag length.

The sign of the coefficient of the cointegration relationship is negative which indicates that increases in inflows should be followed by increases in the IBOVESPA and vice-versa.

As the IBOVESPA (X) and foreign portfolio inflows (Y) are cointegrated (CI(1,1)) (a sufficient condition for causality among these series) then there exists an error correction form given by[9]:

$$\Delta(X_t) = \alpha_1 + \alpha_X (X_{t-1} - Y_{t-1}) + \sum_{i=1}^{n} \gamma_{i1} \Delta X_{t-i} + \sum_{i=1}^{m} \delta_{i1} \Delta Y_{t-i} + \varepsilon_1 \qquad (17)$$

$$\Delta(Y_t) = \alpha_2 + \alpha_Y (X_{t-1} - Y_{t-1}) + \sum_{i=1}^{n} \gamma_{i2} \Delta X_{t-i} + \sum_{i=1}^{m} \delta_{i2} \Delta Y_{t-i} + \varepsilon_2 \qquad (18)$$

The coefficients α_X and α_Y are the speed of adjustment of this system. As shown in table 8, only the coefficients on foreign portfolio inflows is significantly different from zero, which means that foreign portfolio inflows makes all the correction to eliminate any deviations from long-run equilibrium. It is important to note that the absence of Granger causality for cointegrated variables requires that lagged terms for endogenous variables are non significant and also that the speed of adjustment coefficients be equal to zero.

We can see that lagged portfolio inflows are significant and help explain the IBOVESPA average on future months. However, there is bicausality as lagged terms for IBOVESPA helps explain foreign portfolio inflows. Increases in the stock market index seems to induce foreign portfolio inflows.

6. Summary and concluding notes

Under the assumptions of homoscedasticity and heteroscedasticity the RWH was rejected for daily stock prices at the five percent level of significance for the whole sample, even when allowing for a more general test as the multivariate variance ratio statistic due to Chow and Denning (1993).

A rolling variance ratio test gives evidence that the RWH should be rejected when using older data, especially prior to 1994. The data has been divided in two sub-samples and evidence that using recent data we cannot reject the RWH was found. Another aspect

[9] The criteria for choosing lag-length is as usual the minimization of the Schwarz information criterion.

that's worth noting is that the long-run mean reversion in Brazil that previous US studies have found was not found in the Brazilian case[10].

This increase of efficiency could be explained by institutionally maturing markets, the increase of liquidity and the openness of the Brazilian markets for international investors. It is concluded that inflows of foreign portfolio capital indeed had a huge increase after 1994. Evidence of cointegration between the BOVESPA Index and inflows of foreign capital was found, using Johansen´s (1991, 1995) methodology. Indeed, in an error-correction model it is concluded that the inflow series adjusts to make the corrections necessary to maintain the long-run equilibrium relationship, and that there is bicausality between these series.

In future research a more in depth analysis of the effects of the Brazilian American Depositary Receipts (ADR) on the increase of the efficiency of the Brazilian market will be assessed. There are certainly other possible explanations for the increase of efficiency in the Brazilian stock market. This will be subject of further research.

[10] However, this comparison should be looked with caution as these previous studies used longer time spans.

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Table 1
Variance ratio test: daily real returns, 1986-1998 (3051 observations)

Investiment Horizon. q days	2	4	6	8	16	32	48	64
Z1	7.90*	7.79*	7.42*	6.99*	6.45*	4.74*	3.81*	2.79*
Z2	5.30*	5.31*	5.15*	4.90*	4.67*	3.61*	3.02*	2.28#

* Indicates that ratios are statistically different from one at the five percent level of significance for both the Lo and MacKinlay (1988, 1989) and Chow and Denning (1993) tests.
# denotes inferential error where test statistics are separately significant but jointly insignificant.

Table 2
Variance ratio test: daily real returns, 1986-1994 (2046 observations)

Investiment Horizon. q days	2	4	6	8	16	32	48	64
Z1	7.05*	7.57*	7.68*	7.48*	6.74*	4.97*	4.01*	3.02*
Z2	5.09*	5.64*	5.78*	5.74*	5.30*	4.09*	3.43*	2.65*

* Indicates that ratios are statistically different from one at the five percent level of significance for both the Lo and MacKinlay (1988, 1989) and Chow and Denning (1993) tests.

Table 3
Variance ratio test: daily real returns, 1994-1998 (1005 observations)

Investiment Horizon. q days	2	4	6	8	16	32	48	64
Z1	2.26#	0.33	- 0.84	-1.83	-0.96	-0.41	- 0.32	-0.41
Z2	1.33	0.18	- 0.45	-1.00	-0.56	-0.25	-0.21	- 0.29

# denotes inferential error where test statistics are separately significant but jointly insignificant.

Table 4
Variance ratio test: weekly real returns, 1986-1998 (650 observations)

Investiment Horizon. q days	2	4	6	8	10	12	14	16
Z1	3.96*	3.39*	3.05*	2.58*	2.30#	1.93	1.57	1.18
Z2	2.32#	2.06#	1.94	1.71	1.58	1.37	1.14	0.87

* Indicates that ratios are statistically different from one at the five percent level of significance for both the Lo and MacKinlay (1988, 1989) and Chow and Denning (1993) tests.
# denotes inferential error where test statistics are separately significant but jointly insignificant.

Table 5

Tests for equality of means

Category	t-test	Anova F-Statistic
Inflows	13.74*	188.82*
Outflows	12.10*	146.46*

* Indicates rejection of the hypothesis of equality of means between 1990-1994 and 1994-1998 with 99 % confidence.

Table 6

Unit root tests

	Augmented Dickey-Fuller	Phillips-Perron
BOVESPA	-0.97	-0.94
1st difference	-4.74*	-7.86*
INFLOWS	- 1.18	-2.23
1st difference	- 5.97 *	- 16.59 *
OUTFLOWS	-0.39	-1.13
1st difference	- 4.95 *	- 14.01 *

* Indicates rejection of the unit root hypothesis with 99 % confidence.

Table 7

Test for cointegrating relationship

Eigenvalue	Trace	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.1896	21.88	21.88	12.53	None **
0.0084	0.85	3.84	6.51	At most 1

** denotes rejection of the hypothesis at 1% significance level
L.R. test indicates 1 cointegrating equation(s) at 1% significance level

Eigenvalue	Max - Eigenvalue	5 Percent Critical Value	1 Percent Critical Value	Hypothesized No. of CE(s)
0.1896	21.03	11.44	15.69	None **
0.0084	0.8531	3.84	6.51	At most 1

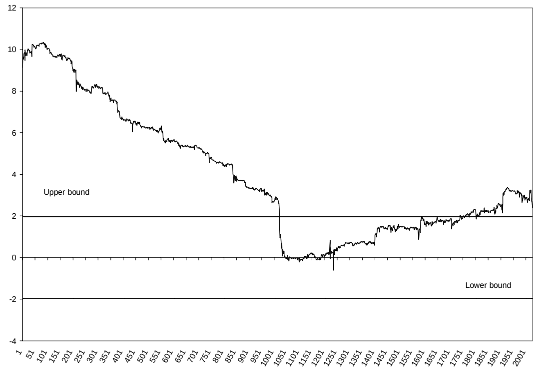
** denotes rejection of the hypothesis at 1% significance level
L.R. test indicates 1 cointegrating equation(s) at 1% significance level

In order to ensure that the tests were not biased or lacked power because of an inappropriate choice of lag length. We considered all lags from 12. The results shown above are for a 3-lags test, which has the minimum multivariate Schwarz information criteria.

Table 8
Error Correction Model

	$\Delta IBOV_t$	$\Delta INFLOW_t$
Cointegration Equation	-0.0050	0.1260*
	(0.0101)	(0.0313)
	(-0.4996)	(4.0236)
$\Delta IBOV_{t-1}$	0.2829*	0.7550*
	(0.1017)	(0.3156)
	(2.7801)	(2.3922)
$\Delta IBOV_{t-2}$	-0.2067*	0.2676
	(0.1030)	(0.3195)
	(-2.0066)	(0.8375)
$\Delta IBOV_{t-3}$	-0.0800	-0.0844
	(0.0946)	(0.2933)
	(-0.8466)	(-0.2878)
$\Delta INFLOW_{t-1}$	0.0508**	-0.4597*
	(0.0264)	(0.0821)
	(1.9211)	(-5.5997)
$\Delta INFLOW_{t-2}$	0.0326	-0.4080*
	(0.0292)	(0.0907)
	(1.1155)	(-4.4965)
$\Delta INFLOW_{t-3}$	0.0827*	-0.0624
	(0.0275)	(0.0855)
	(3.0005)	(-0.7306)
R-squared	0.1699	0.3616
Adj. R-squared	0.1163	0.3204
F-statistic	3.1725	8.7811
Log likelihood	61.9145	-51.2645
Schwarz	-0.9159	1.3476

* Rejection of the null with 99 % confidence
** Rejection of the null with 95 % confidence
Standard errors and t-statistics in parentheses.



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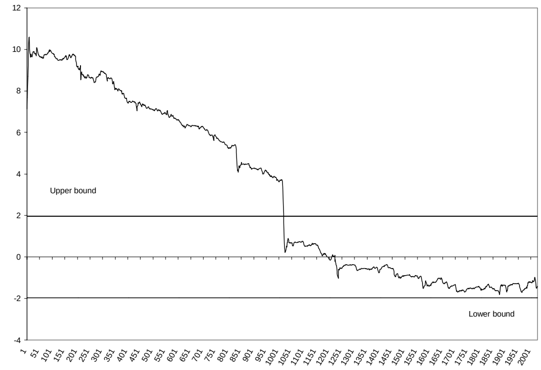

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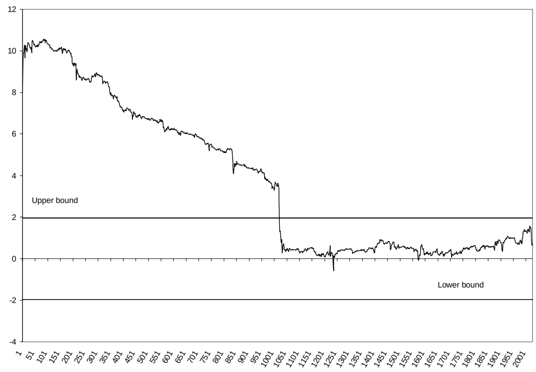

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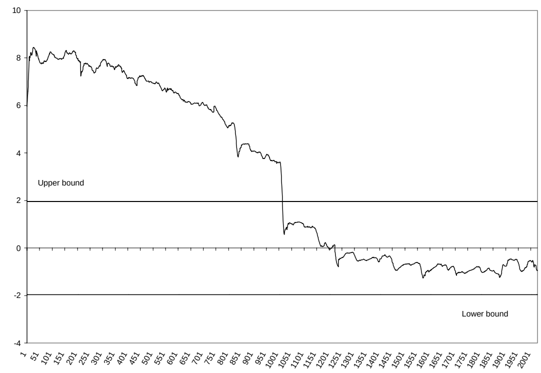
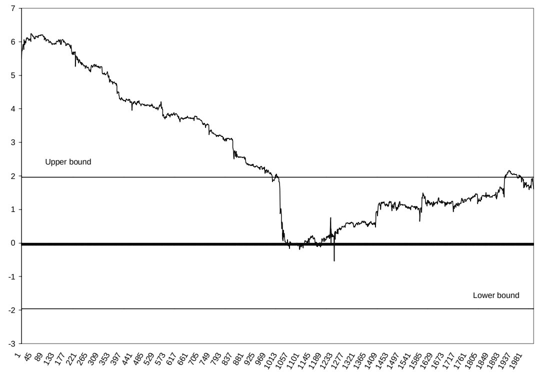


q = 16

Figure 1. Z1 estimates for exchange rate-adjusted returns over a moving 1024 days sample period starting with 01/03/86 and ending with 06/12/98, for differet investment horizons (q = 2, 4, 8 and 16 days). Solid lines around 1.96 and –1.96 represent a two-sided 95% confidence interval for Z1 under H0.



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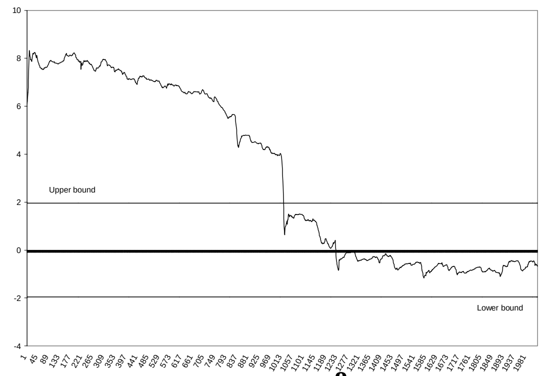

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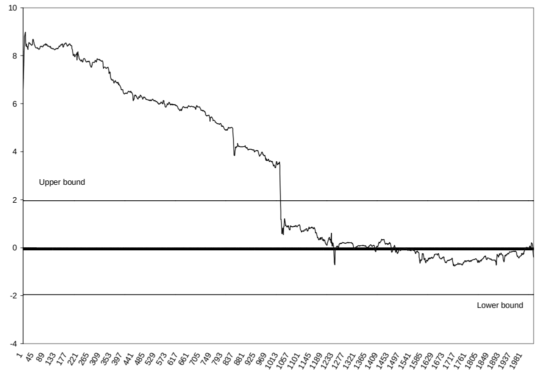

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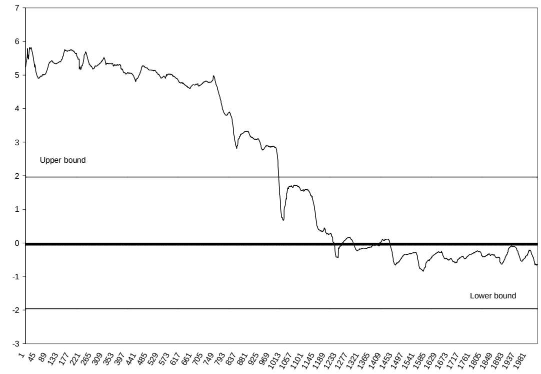

q = 16

Figure 2. Z2 estimates for exchange rate-adjusted returns over a moving 1024 days sample period starting with 01/03/86 and ending with 06/12/98, for different investment horizons (q = 2, 4, 8 and 16 days). Solid lines around 1.96 and –1.96 represent a two-sided 95% confidence interval for Z1 under H0.

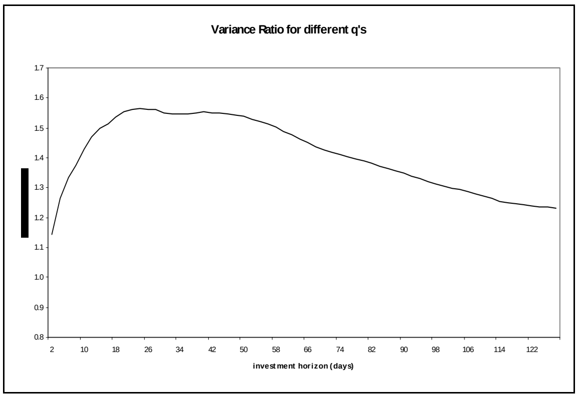

Figure 3. Variance Ratios for different investment horizons.

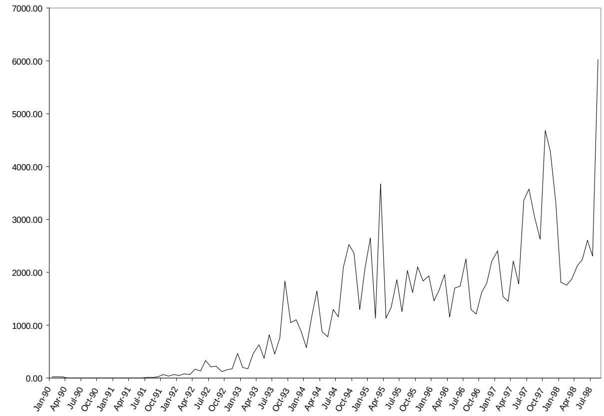

Figure 5. Portfolio Inflows from January 1990 to June 1998 (in US$ million dollars).

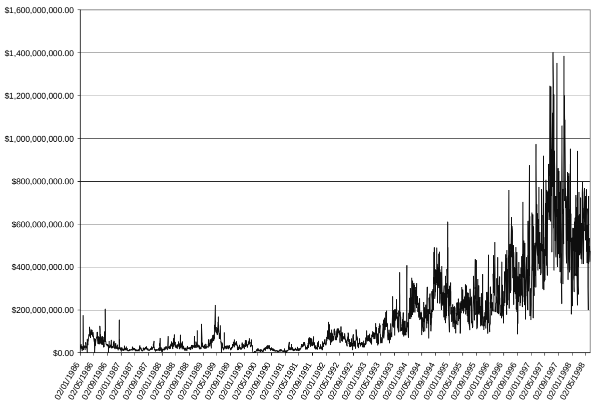

Figure 4. Daily traded volume in stocks in the BOVESPA index starting in 01/03/86 and ending in 06/12/98 (in US$).

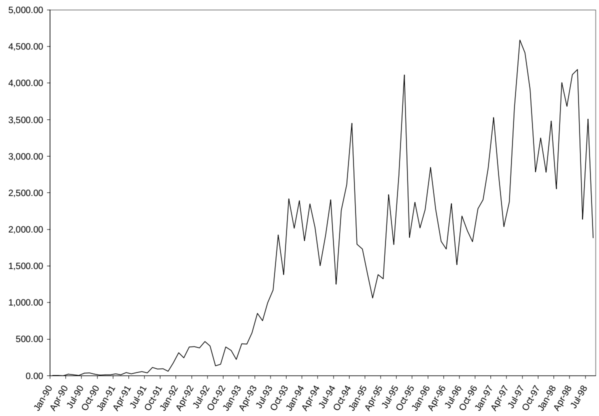

Figure 6. Portfolio Outflows from January 1990 to June 1998 (in US$ million dollars).

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