CENTRO DE ESTUDIOS MONETARIOS LATINOAMERICANOS

2005

MONEY AFFAIRS

VOLUME XVIII, NUMBER 1, JANUARY-JUNE

MONEY AFFAIRS is a bi-yearly publication of the Centre for Latin American Monetary Studies (CEMLA), Durango nº 54, Mexico City, D. F., 06700. ISSN-0187-7615.

MONEY AFFAIRS is regularly listed in the International Current Awareness Service: Economics. Selected material is indexed in the International Bibliography of Economics.

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Concentration and price rigidity: evidence for the deposit market in Chile

I. INTRODUCTION

For the conduction of monetary policy it has an outstanding importance the effect of changes of the base rate over the market interest rates. When the effects of monetary policy over prices and output are evaluated, it is often assumed that there is a complete and quick pass-through. However, there is international evidence that supports the fact that there is important sluggishness of market interest rates.[1] It might be presumed that the predictability and effectiveness of a change on the policy rate would depend significantly on the flexibility of market interest rates. Addition-

[1] Hannan and Berger (1991), Newmark and Sharpe (1992), Scholnick (1996), Heffernan (1997), Blinder (1998), Mizen and Hofmann (2002).

Paper prepared by Solange Berstein, of the Superintendence of Pension Fund Administrators, and Rodrigo Fuentes, of the Central Bank of Chile, and presented at the IX Meeting of the Central Bank Researchers Network, held in San José, Costa Rica. We thank helpful comments by Leonardo Hernández, Fernando Parro and the participants to the internal seminar at the Central Bank. The views and conclusions presented in this paper are exclusively those of the authors and do not necessarily reflect the position of the Central Bank of Chile, nor of the Superintendence of Pension Fund Administrators.

ally, in the case of Chile as in many other countries, market concentration on the banking industry has increased considerably over the last years, which according to Hannan and Berger (1991) would imply stronger price rigidity.

Price stickiness can be a consequence of a collusive behavior as it is modeled by Hannan and Berger (1991), or menu costs, as in Blinder (1994), or durable relationships between banks and customers as a result of switching costs (Newmark and Sharpe, 1992). It is also the case that differences are observed between banks and even between different products offered by the same bank.

The analysis presented in this article includes a time series examination of the deposit interest rates, testing the effects of concentration over price rigidity. In addition, panel data estimation at the bank level is exposed, which considered the effects of bank characteristics over the speed of adjustment. The results support the fact that there is some sluggishness in deposit interest rates and that the stickiness increases with market concentration. At the bank level we found that certain characteristics of banks as solvency, market share and credit risk jointly with market concentration are the determinants of the speed of the deposit rate adjustment to changes in the monetary policy rate. The inclusion of variables like credit risk and solvency try to capture whether market discipline has anything to do with the transmission of the monetary policy rate to deposit rates.

The paper proceeds as follows. Section II provides with a review of the literature, stressing the conceptual framework of the analysis. Afterwards, Section III has a data description and Section IV shows the time series results. Subsequently, Section V presents the panel data estimation, including some methodological issues and the results. Finally, Section VI concludes.

II. LITERATURE REVIEW AND CONCEPTUAL FRAMEWORK

There is a broad literature that relates deposit interest rates stickiness with market concentration. One of the seminal papers that study this relationship is Berger and Hannan (1989). This article tries to identify the structure-performance hypothesis from the efficiency-structure hypothesis. The former would mean that there is collusion in a certain market and the second one would mean that firms with different levels of efficiency would survive in a concentrated market. In this last case, as firms that are more efficient would have a higher market share, a study that relates profits with concentration will conclude that there is a positive relationship, but the reason would be that there are more efficient firms in the market, and not necessarily a collusive behavior. So the policy implications are different from the case where the structure-performance hypothesis prevails. To identify this, instead of looking to the profit concentration relationship, they study the price concentration relationship by using a panel of U.S. banks in different markets, for the period that goes from September 1983 to December 1985. The paper gives evidence that supports the fact that more concentrated markets imply lower deposit rates than less concentrated markets.

The same authors, Hannan and Berger (1991), provide a stylized model of monopolistic competition that illustrates how firms with market power not necessarily change prices when there is a change in costs. This theoretical model shows how firms decide to change prices or not by comparing costs and benefits of such decisions; moreover, they allow for differences between down-pricing and up-pricing decisions. For the U.S. banking industry they found that there is greater price rigidity in more concentrated markets, and the stickiness was higher when there was a stimulus to increase deposit rates.

A later paper, Newmark and Sharpe (1992), explores evidence of price rigidity for the banking industry by using a different methodology. This article argues that there are long run relationships between banks and its customers, which would imply certain degree of stickiness in prices. The evidence found in this paper supports the facts stated Hannan and Berger (1991), this is that higher concentration imply more rigidity and that decreases in deposit rates are faster than increases. However, Jackson III (1997) argues that there is a non monotonic relationship between concentration and price rigidity, the paper provides an empirical estimation based on the model taken from Worthington (1989). A different approach is presented in Sharpe (1997), this paper considers Klemperer’s (1995) switching costs model for the case of bank deposit interest rates, arguing that in the presence of switching costs banks have monopoly power which imply lower deposit rates. The authors identify the effects of switching costs by separating locations with high presence of movers where it is assumed that movers have no switching costs, so that locations with high portion of movers would have higher deposit rates.

Each of the above studies uses panel data analysis for different time periods, different methodologies and data of different locations for the U.S. economy. There is a smaller amount literature that explores this subject for other countries. In fact, studies for other countries investigate the dynamics of deposit interest rates, by using time series analysis instead of panel data. These papers focus on deposit interest rate pass-through without directly estimating its relationship with concentration, but interpreting any findings of stickiness as a signal of collusion. This is the case of Scholnick (1996), which estimates speed of adjustment for Malaysia and Singapore. The methodology considers an Error Correction Model (ECM) that it is estimated for both countries, and explores price rigidity and possible asymmetries between increases and decreases of deposit rates. For the U.K. Heffernan (1997), by also using an ECM finds significant differences between banks, products and over time, even between products offered by a same bank. Mizen and Hofman (2002) also study the UK, by using an ECM allowing for asymmetries between increases and decreases of deposit rates, but assuming also non-linearities, arguing that there might be a different response depending on the size of the change. They found that there is complete pass-through in the long run for deposit rates and the speed of adjustment increases when the gap between the retail rate and the base rate is widening and it get slower when the movements are in the direction of automatically closing the gap. Another interesting finding is that the speed of adjustment was affected by expectation and interest rate volatility, but not concentration.

Summarizing, the literature supports the fact that there is interest rate stickiness and that concentration implies lower deposit rates and more rigidity. Moreover, there seems to be differences across markets, banks and products that might be explained by other factors, not only market concentration. In this sense, there is another line of literature that analyzes market discipline in depositors behavior. According to this literature, interest paid should be higher for banks that show lower performance, because they would appear to be riskier. Therefore, these banks would be penalized in a world where there is less than 100% deposit insurance. If this were the case, banks that show lower performance not only would pay higher interest rates but potentially might be the case that the pass-through of changes in the policy rate would be different according to bank characteristics. In fact, Cook and Spellman (1994) show that deposit interest rates respond to individual bank risk factors, even in the case there is 100% insurance.

Peria and Schmukler (2001) also provide evidence of market discipline for Argentina, Chile and Mexico. Finally, Budnevic and Franken test the market discipline hypothesis for Chile and found stronger evidence for interest rates than for the quantity of deposits. For testing market discipline the study considers a CAMEL (Capital Adequacy, Assets Quality, Management, Earnings and Liquidity) indicator for each bank.

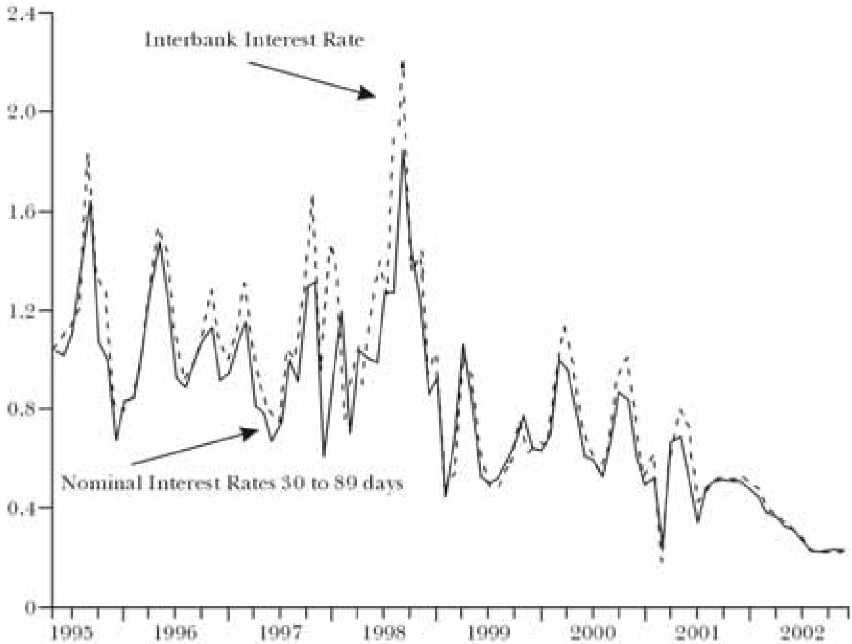
Thus, this paper studies the deposit interest rate stickiness at an aggregate level and afterwards we look for differences across banks by using panel data estimation. In the analysis it is considered the effect on price rigidity of concentration, so that we test for the possibility of a positive relationship between these two variables, which would be consistent with previous findings. Additionally we include bank characteristics, to capture any effect of these variables over price sluggishness, which might be a consequence of market discipline.

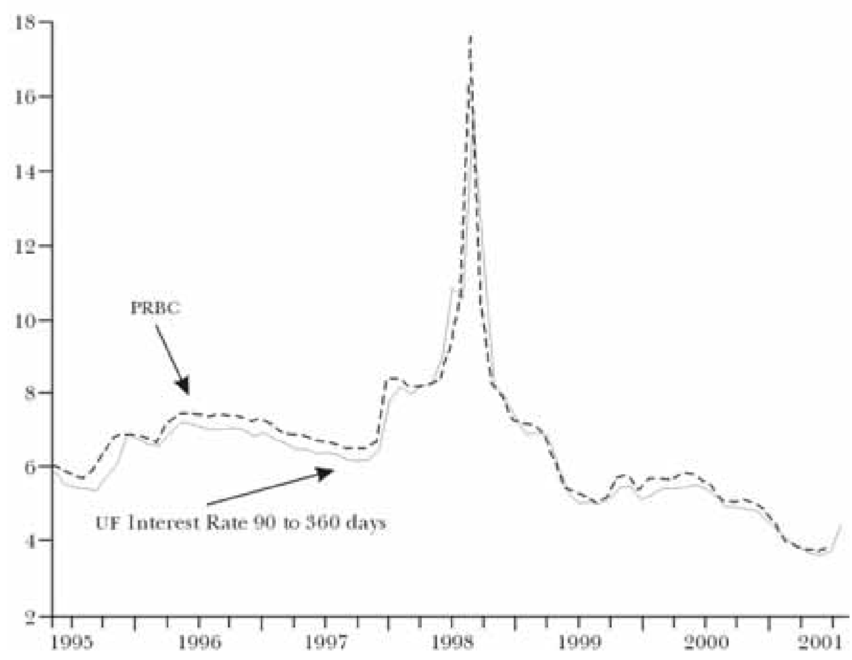
III. DATA DESCRIPTION

The data required for the analysis are basically deposit interest rates of different denominations and maturities. These interest rates are the effective interest rates for transactions that take place during a specific month. It is important to notice that in the Chilean financial system there are three units of account that coexist: peso, US dollars and UF. The UF is a unit of account indexed to the previous month inflation. It varies daily with the past inflation but the one-month variation is exactly the previous month inflation. Deposits and financial instruments, in general, of short-term maturity (less than 90 days) are usually expressed in pesos, while medium and long-term deposits (90 days and above) are denominated in UF terms. Dollars denominated deposits have small share of total deposits (less than 10% on average in our sample period, but increasing over time).

The monetary policy is announced using a monetary policy rate in UF terms. However, in practice, since May 1995, the monetary policy is implemented using the money market rate, which is a nominal rate. In August 2001, the Central Bank modified the denomination of the monetary policy rate from indexed to nominal rate. This change has several short-term consequences on the financial market and two important long-term effects. First, the volatility of the nominal rates decreases, but as counterpart the volatility of indexed interest rate increases. Second, it help-

FIGURE III.1. EVOLUTION OF THE INTEREST RATES, 1995-2002





ed to implement a more expansive monetary policy in the last two years in Chile.[2]

This paper analyzes the relationship between the monetary policy and the deposit rate exploiting monthly data at the aggregate level of the banking industry as well as at the individual bank level. By a first look at the aggregate data, one notices that the deposit rate follows closely the money market rates (Figure III.1). For the UF deposit rate we use the interest rate on 90 days Central Bank promissory notes, which is denominated in UF until august 2001. After that period due to the nominalization process, these promissory notes were issued in nominal terms.

For the bank level analysis, specific bank characteristics were required. For this purpose it was collected information from banks balance sheets. The variables chosen were solvency, liquidity, risk and size. Solvency was computed as capital over total assets. Liquidity is measured as liquid fund over demand deposit. Concerning size the variable was defined as market share defined over total deposit. Finally, risk is measured as non-performing loans over total loans. Different measures of concentration were used C3, C5 and Herfindhal index in terms of total deposits.

IV. TIME SERIES RESULTS WITH AGGREGATE DATA

An empirical model that intends to capture the effect on the deposit rate adjustment of concentration to changes in the policy rate using data of the banking industry for the time period between May 1995 and December 2002 for the case of nominal rates is estimated.[3] Thus the equation to be estimated is:

$$y_t = \delta + \sum_{j=1}^{m} \beta_j y_{t-j} + \sum_{k=0}^{n} \alpha_k z_{t-k} + \sum_{l=0}^{p} \gamma_l \Delta MPR_l + \varepsilon \qquad (1)$$

where y represents the bank-deposit rate, z the money market or interbank rate, ΔMPR the change in the monetary policy interest rate, and ε is the error term that is assumed to be white noise. The difference between the money market or interbank rate and the monetary policy rate is that the first two are interest rate determined in the market, while the latter is set by the Central Bank as a target value. In Chile monetary policy is conducted, as in many other countries, by managing liquidity such that the inter-

[2] See Fuentes et al. (2003).

[3] For the case of indexed rates the data is from May 1995 to July 2001.

bank or money market rate is in line with the policy rate. One of the coefficients of interest is α_0, which measures the impact effect of a change in the money market rate on the deposit rate. The other coefficient of interest is the one that measures the long run effect:

$$\lambda = \frac{\sum \alpha_k}{1 - \sum \beta_j} \tag{2}$$

To complete the model we establish a relationship between the coefficients of interest and our measure of concentration given by the Herfindhal index (H). We assume that α_0 is a linear function of H, and each coefficient in (1) is a linear function of H. Thus the long-term coefficient is a non-linear function of H. That is:

$$y = 1_T \delta + X\mathrm{B} + Z\mathrm{A} + R\Gamma + \mu \tag{3}$$

where 1_T is a vector of ones, B, A, Γ are vectors of parameters. X is a $T \times 2l$ matrix comprised by lags of the dependent variables and the interaction variables. Z is a $T \times (2l+2)$ matrix of the contemporaneous and lags values of the money market rate and the interaction of *z* and the Herfindhal index. R is a $T \times (2l+2)$ matrix of the contemporaneous and lags values of the monetary policy rate and the interaction of MPR and the Herfindhal index. Each element of X, say x_{ij}, is defined as:

$$x_{ij} = \begin{cases} k_{t-j} & j = 1,\ldots,l \\ k_{t-j+l} H_{t-j+l} & j = l+1,\ldots,2l \end{cases} \tag{4}$$

Where in the case of Z and R the variable *k* is replaced by the money market rate and the MPR. Note that the number of lag *l* in each case could be different and they are chosen in order to make μ white noise. It is worth noticing that the model is estimated in levels, because there is no economic reason for unit roots for the variables used and in any case unit root tests are included in the Appendix.

Table IV.1 shows the estimation results of equation (3) for short-term deposits that received a nominal interest rate. Model [1] does not control for the year 1998, and it shows a smaller impact coefficient than model [2], where the year 1998 is controlled for. This result implies that in an unusual year the banks do not pass through the jump in the interest rate to the deposit rate. In any case the coefficient is not very different and it varies from

0.75 to 0.88, meaning that banks modify the deposit rate in 75% or 88% when they face a change in the interbank interest rate.

It is interesting that our measure of concentration does not affect the size of the impact coefficient. However concentration affects the coefficient of the lags variables and thus it affects the long run coefficient. Table IV.2 shows the value of this coefficient when concentration is evaluated in the mean, the median, the maximum and the minimum of concentration, as a way to see the effect of the Herfindhal index on the long-run parameter. This exercise shows that at the mean or the median the coefficient is statistically equal to 1. But, market concentration affects negatively the interest rate pass through.

TABLE IV.1. NOMINAL RATE FOR 30 TO 89 DAYS

	[1]	*[2]*
Constant	-0.217	0.053
	[0.154]	[0.028]
Interbank Rate	0.755	0.884
	[0.028]***	[0.023] ***
Interbank Rate (-1)	0.749	0.514
	[0.249]***	[0.117] ***
Interbank Rate (-5)	-0.165	
	[0.071]**	
Nominal Rate 30ds (-4)	0.130	0.184
	[0.055]**	[0.033] ***
Nominal Rate 30ds (-5)	0.151	
	[0.084]*	
Nominal Rate 30ds (-6)	0.114	
	[0.037]***	
DTPM(-2)	0.055	0.030
	[0.011]***	[0.007] ***
Herf	0.028	
	[0.016]*	
Herf(-1)*Interbank Rate (-1)	-0.106	-0.054
	[0.026]***	[0.012] ***
Herf(-1)*Nominal Rate 30ds (-1)	0.049	
	[0.010]***	
Herf(-2)*Nominal Rate 30ds (-2)	-0.039	-0.017
	[0.008]***	[0.004] ***
Herf(-3)*Nominal Rate 30ds (-3)	0.013	
	[0.005]**	
Adjusted R-squared	0.972	0.979
SE of regression	0.053	0.046
Durbin-Watson statistic	2.190	1.707

Standard deviations in brackets. In model [2] we control for year 1998.
* Significant at 10%; ** significant at 5%; *** significant at 1%.

TABLE IV.2. IMPACT AND LONG RUN COEFFICIENTS FOR NOMINAL INTEREST RATE 30 TO 89 DAYS

	Model [1]		*Model [2]*	
Concentration	*Impact*	*Long Run*	*Impact*	*Long Run*
Mean	0.755	1.05**	0.884	0.97**
Median	0.755	1.03**	0.884	0.96**
Maximum	0.755	0.66	0.884	0.84
Minimum	0.755	1.28*	0.884	1.06**

In model [2] we control for year 1998. Chi-Square (1) in brackets for λ=1.
** Can't reject at 5%, * at 10%.

TABLE IV.3. 90 DAYS INDEXED INTEREST RATE

	[1]	*[2]*
Constant	-0.078	-5.009
	[0.087]	[1.93]**
PRBC	0.774	1.511
	[0.008]***	[0.29]***
PRBC (-1)	-0.338	
	[0.141]**	
PRBC (-4)	0.052	0.061
	[0.027]*	[0.030]**
UF 90 ds 1yr (-1)	0.691	0.318
	[0.183]***	[0.010]***
UF 90 ds 1yr (-2)	-0.206	-0.100
	[0.046]***	[0.011]***
Herf		0.593
		[0.219]***
Herf * PRBC		-0.098
		[0.032]***
Herf (-3)*UF 90 ds 1yr (-3)	0.012	0.007
	[0.003]***	[0.001]***
Herf (-4)*UF 90 ds 1yr (-4)	-0.010	-0.008
	[0.003]***	[0.004]**
Adjusted R-squared	0.983	0.997
SE of regression	0.237038	0.103747
Durbin-Watson statistic	1.969082	1.959175

Standard deviations in brackets. In model [2] we control for year 1998.
* Significant at 10%; ** significant at 5%; *** significant at 1%.

To check the robustness of our results we cut the sample in July 2001, to isolate the process of nominalization. Our results did not change much. We tried other measures of concentration like C3 and C5, but the results did not change in a qualitative manner. We also explored for the existence of asymmetrical effects between ups and downs of the interbank interest rate. For doing so we introduced a dummy variable that takes a value equal to 1 when the interbank rate increases. We test for changes in every slope coefficient, but we couldn't find evidence of asymmetries.

Using indexed deposit interest rate we estimated equation (3). Now the money market rate was associated to the 90 days Central Bank promissory note. The results are shown in Table IV.3. When the 1998 effect is not controlled, concentration does not affect the impact coefficient. But after controlling for that year effect the relationship between the impact coefficient and concentration become significantly negative.

Table IV.4 shows the result for the relation between market concentration and the pass through coefficient. Again, at the average level of concentration the long-term coefficient is statistically equal to 1 in model [1], but not in model [2]. In this case when controlling for year 1998 effect, concentration affect negatively both coefficients, meaning that more concentrated makes slower pass through interest rate movements, even in the long run. This result is consistent with the international evidence. However in the case of Chile we could not find evidence of asymmetries in the pass through between ups and downs, which has been the case of previous studies for other countries.[4] In fact a dummy variable that takes value equal to 1, when the interbank rate increases cannot find to have an economically significant coefficient.[5]

TABLE IV.4. IMPACT AND LONG RUN COEFFICIENTS FOR UF INTEREST RATE 90 DAYS TO 1 YEAR

	Model [1]		*Model [2]*	
Concentration	*Impact*	*Long Run*	*Impact*	*Long Run*
Mean	0.774	0.991**	0.666	0.930
Median	0.774	0.992**	0.656	0.913
Maximum	0.774	1.001**	0.493	0.849
Minimum	0.774	0.985**	0.782	1.066**

In model [2] we control for year 1998.

** Chi-square (1) λ=1 Can't reject at 5%, * at 10%.

[4] Hannan and Berger (1991).

[5] Espinoza and Rebucci (2003) found no evidence of asymmetries for Chile. They also found that Chile do not have different pass through coefficient when comparing with a group of OECD countries.

V. PANEL DATA ESTIMATION

In the previous section using aggregate data we explored the relationship between market concentration and the pass through interest rate coefficient. In this section we study this relationship using data at the individual bank level. The advantage of doing so is twofold. On the one hand, it allows for controlling by specific bank characteristics. In an environment of market discipline, depositors will choose carefully where they are making their deposits. On the other hand, the panel data analysis gives equal weight to all banks. With aggregate data, large banks may drive the results.

A similar analysis conducted by Berstein and Fuentes (2003) found that bank characteristics matter for the pass through of the monetary policy rate to bank lending rates. They also found that the short run coefficient was around 0.7 and the long run tends to be equal to 1.

In this paper we construct a panel data using nominal and indexed interest rate at the bank level as dependent variables. The explanatory variables are those used in the previous section plus banks characteristic defined in section III (liquidity, solvency, size and risk portfolio). For short-term deposit our sample includes 21 banks, and 20 banks for the 90 to 360 days deposit. Recall that short-term deposits are denominated in pesos and longer-term deposits are in UF.

1. Methodological Issues

The literature on dynamic panel data estimation, as our empirical model presented in section IV, has been revitalized in the second half of the nineties. Anderson and Hsiao (1981) presented the well-known problem of inconsistency of the least square dummy variable estimate in dynamic panel data. They proposed a method based on instrumental variable, which consist of taking first differences of the equation to eliminate unobserved heterogeneity and then use instrumental variables to estimate consistently the parameters of the lag dependent variables.

For instance, let's assume that the following equation is to be estimated using panel data:

$$y_{it} = \rho\, y_{it-1} + \beta\, x_{it} + \eta_i + u_{it} \quad (t = 1,...,T\,; i = 1,...N) \tag{5}$$

Where y_{it} represents the lending interest rate, x_{it} represents a dependent variable like the interbank interest rate, η_i is the unobserved heterogeneity. Taking first difference the equation to be estimated is:

$$y_{it} - y_{it-1} = \rho(y_{it-1} - y_{it-2}) + \beta(x_{it} - x_{it-1}) + u_{it} - u_{it-1} \quad (6)$$

Anderson and Hsiao propose $y_{i,t-2}$ or ($y_{i,t-2}$- $y_{i,t-3}$) as instrument for ($y_{i,t-1}$- $y_{i,t-2}$). But Arellano (1989) showed that $y_{i,t-2}$ is a better instrument for a significant range of values of the true ρ in equation (6).

Arellano and Bond (1991) proposed an alternative methodology based on GMM estimators. This method used several lags of the variables included as instruments, so it is especially efficient when T is small and N is large.[6] The method is based on T(T-1)/2 moment condition and it is consistent for fixed T or for T that grow to a slower path than N when both go to infinity. The application of this method requires that $T-1 \leq N$.

In a recent paper Alvarez and Arellano (2003) show the asymptotic property of the within group, GMM and LIML estimators. An important result for our case is that, regardless the asymptotic behavior of N, when T goes to infinity the estimator of ρ is consistent. Moreover, if lim(N/T)=0 as N and T goes to infinity, there is no asymptotic bias in the asymptotic distribution of the within group estimator, while in the opposite case if lim(T/N)=0, as N and T goes to infinity, there is no asymptotic bias in the asymptotic distribution of the GMM estimator. In the case of our panel T is large and it will increase as the time goes by, while N will remain relatively fixed, thus the traditional within group estimator will provide better results.

2. Panel Data Estimation

Using the data described above and the methodology, which was explained in the previous section, we estimate equation 3. We assume that the responsiveness of deposit interest rate to monetary policy rate is affected by the level of concentration of the banking industry, the market share of each bank (as a proxy of size) and solvency. The hypothesis for concentration was explained in section II. Market share is used for testing whether large banks are able to pay lower interest rate on deposits. Sol-

[6] Judson and Owen (1999) provided evidence that for small T, GMM is a better estimator than Anderson and Hsiao's methods under the mean square error criterion. But for unbalanced panel data and T around 20 is unclear what method is better.

vency may affect the speed of adjustment of deposit rate, since one should expect that more solvent banks will not pass through the monetary policy rate at the same speed as less solvent bank. This hypothesis comes from the market discipline literature.

Table V.1 summarizes the results for the coefficients of interest in the case of nominal interest rate and the appendix shows the results in greater detail. To isolate the effect of each variable on the coefficients, we evaluated the value for each coefficient moving one variable at the time. For instance, we estimate the impact coefficient for the maximum and the minimum values of solvency, assuming that the other variables are equal to their sample mean. According to our results, concentration and market share do not affect the short-term or impact coefficient. However, solvency affects negatively, consistent with Cook and Spellman (1994), the long term and the impact pass through coefficient, i.e. banks that are more solvent adjust the deposit rate to a lower path. As shown the pass through coefficient varies between 0.573 and 0.886.

TABLE V.1. IMPACT AND LONG RUN COEFFICIENTS FOR NOMINAL INTEREST RATE 30 TO 89 DAYS

	Impact	*Long Run*
Concentration		
Mean	0.849	0.965
Median	0.849	0.975
Maximum	0.849	0.835
Minimum	0.849	1.089
Market Share		
Mean	0.849	0.965
Median	0.849	0.963
Maximum	0.849	0.998
Minimum	0.849	0.958
Solvency		
Mean	0.849	0.965
Median	0.871	0.970
Maximum	0.573	0.841
Minimum	0.886	0.973

In the long run the three variables are relevant determinants of the stickiness in deposit rate. In any case, taking the maximum and the minimum of each variable at the time (the other two are set equal to their mean value) the long-term coefficient fluctuates from 0.83 to almost 1.1, showing a higher degree of flexibility in the long-term than in the short term. As expected, according to

Hannan and Berger (1991), concentration has a positive effect on the degree of stickiness of the deposit rate. The opposite was found for market share, large banks tend to show a more flexible deposit rate, which could be due to higher level of efficiency. When considering solvency, the long-run coefficient shows the same pattern than the impact coefficient. Banks that are more solvent tend to pass through the monetary policy rate more slowly since they are more reliable. This is consistent with the market discipline hypothesis.

Table V.2 presents the results for the indexed interest rate for 90 days to 1 year. In this case market share did not turn significant, while concentration and solvency remain as important explanatory variables of the speed of adjustment. These two variables have similar effect as in the case of nominal rates for the long-run coefficient. However for the impact coefficient, concentration was important but not solvency. Additionally, in this exercise credit risk became important to explain the speed of adjustment. Those banks with a riskier (ex - post) loan portfolio tend to exhibit a higher degree of sluggishness. This result will go against the market discipline hypothesis, since one should expect that the degree of stickiness increase with the level of bank risk. A plausible explanation for our finding is that the risk variable is a measure of ex – post risk, and what is relevant for the depositor is the ex – ante risk. On the other hand, from the bank's point of view

TABLE V.2. IMPACT AND LONG RUN COEFFICIENTS FOR NOMINAL INTEREST RATE 90 TO 360 DAYS

	Impact	*Long Run*
Concentration		
Mean	0.746	0.92
Median	0.754	0.93
Maximum	0.565	0.86
Minimum	0.826	1.11
Risk		
Mean	0.746	0.923
Median	0.751	0.929
Maximum	0.651	0.805
Minimum	0.780	0.964
Solvency		
Mean	0.746	0.923
Median	0.746	0.932
Maximum	0.746	0.856
Minimum	0.746	0.949

FIGURE V.1. PASS THROUGH COEFFICIENTS AND CONCENTRATION: NOMINAL RATE 30 DS, 1997-2002

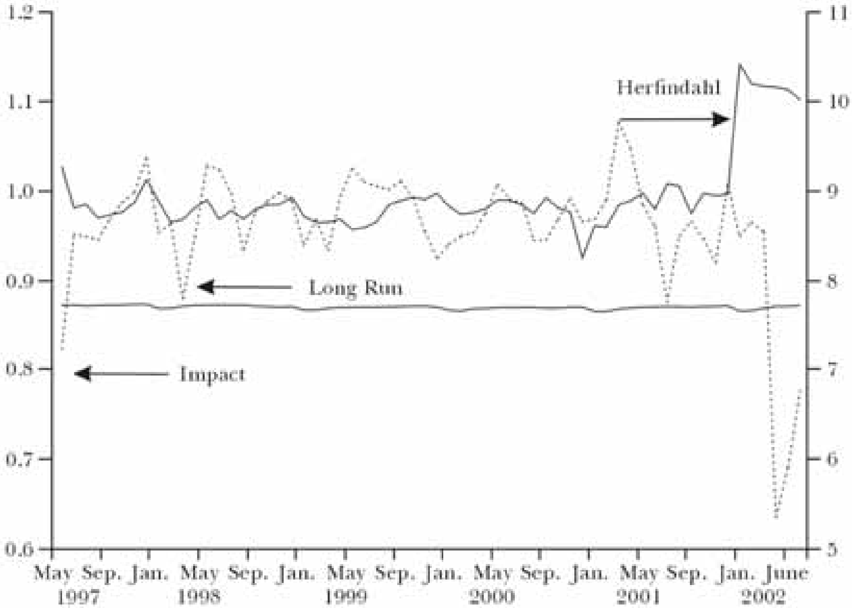
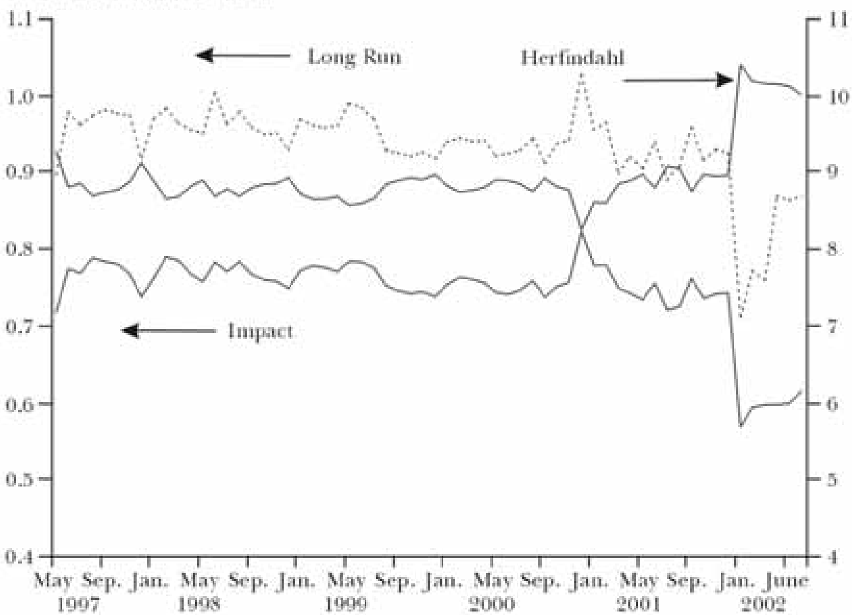


FIGURE V.2. PASS THROUGH COEFFICIENTS AND CONCENTRATION: UF RATE 90 DS TO 1 YR, 1997-2002



when it faces a larger amount of unpaid loans, the bank would need a higher spread to cover from that risk. Therefore it will tend to pay a lower deposit rate and to have a slower adjustment in this interest rate.

Figures V.1 and V.2 show the overtime evolution of the aggregate impact and the long-term coefficients. This evolution is determined by the effect of concentration, solvency, market share, and credit risk on the pass through coefficients. In the same graph, on the left-hand side axis we show the evolution of our measure of concentration. Concentration does not imply movements on the impact coefficient for the nominal interest rate, but it does in the case of UF denominated deposits. At the end of the period the Herfindahl increased due to a merge between two large banks that drastically reduce the long-term pass through coefficient for nominal and indexed deposit rates. Besides, for the indexed interest rate concentration seems to drive the results at the end of the period where there is an increase in concentration and a reduction of both the impact and the long-term coefficient. Note that in this case the deposit rate is for longer-term deposit, which is different for the nominal case.

VI. CONCLUSIONS

There is consensus with respect to the importance of market interest rate flexibility for the conduction of monetary policy. When the effects of monetary policy over prices and output are evaluated it is often assumed that there is a complete and quick pass-through. However, there is international evidence that supports the fact that there is important sluggishness of market interest rates.[7] In the case of Chile there is evidence of sluggishness of adjustment in the case of lending interest rates; however, compared to other countries it appears to be more flexible than average.[8]

In terms of deposit interest rates in many other countries it has been found that there is significant rigidity and that it is closely related to market concentration on the banking industry.[9] Moreover, concentration of these industries around the world has in-

[7] Hannan and Berger (1991), Newmark and Sharpe (1992), Scholnick (1996), Heffernan (1997), Blinder (1998), Mizen and Hofmann (2002).

[8] Berstein y Fuentes (2003).

[9] Hannan and Berger (1991).

creased considerably over the last years, which is also the case for Chile.

The evidence presented in this article supports the fact that there is some rigidity for deposit interest rates and that it is significantly related to concentration. For instance, as concentration has increased over the last years, sluggishness of deposit interest rates has also increased. In addition, panel data estimation at the bank level supports this finding and also allows identifying the effects of bank characteristics over the speed of adjustment.

In the case of short run nominal rates it was found that larger banks tend to show a more flexible deposit rate, which could be due to higher level of efficiency. When considering solvency, banks that are more solvent tend to pass through the monetary policy rate more slowly since they are more reliable. For indexed interest rates, market share did not turn significant, while solvency continues to be significant with a similar effect to the case of nominal rates for the long-run coefficient. These findings are consistent with the market discipline hypotheses, in the sense that banks that are more trustworthy would have lower deposit interest rates and adjust at a slower path.

Appendix A

UNIT ROOT TEST FOR DEPOSIT RATES AND POLICY RATES (1995-2001)

	ADF	*DF-GLS*	*Phillips-Perron*	*Phillips-Perron Ng Mzt*
PRBC	-1.928	-1.949*	-2.630	-1.995*
Interbank Nominal Rate	-3.733*	-3.175*	-4.364**	-3.135*
UF 90 ds. to 1 year	-2.179	-2.085*	-2.172	-1.999*
Nominal 30 to 89 days	-5.380 **	-5.421**	-5.250**	-4.224**

* No stationarity rejected at 5%; ** No stationarity rejected at 1%.

Appendix B

Panel data estimation

NOMINAL RATE FOR 30 DAYS

	[1]	*[2]*
Interbank Rate	0.724 [0.019]***	0.905 [0.029]***
Interbank Rate(-1)	4.732 [1.137]***	1.091 [0.151]***
Interbank Rate (-2)	-2.657 [1.162]**	
Interbank Rate (-3)	-0.808 [0.186]***	-0.340 [0.132]**
Interbank Rate (-4)	0.246 [0.113]**	-2.480 [0.824]***
Interbank Rate (-5)	3.068 [0.918]***	0.305 [0.031]***
Nominal Rate 30ds (-1)	-2.579 [1.212]**	0.268 [0.134]**
Nominal Rate 30ds (-2)	3.203 [1.293]**	
Nominal Rate 30ds (-5)	-4.822 [1.036]***	1.776 [0.914]*
Nominal Rate 30ds (-6)	0.791 [0.127]***	0.447 [0.104]***
DTPM(-1)	0.015 [0.004]***	0.017 [0.005]***
Herf	0.051 [0.008]***	0.035 [0.006]***
Herf (-1)*Interbank Rate (-1)	-0.555 [0.129]***	-0.160 [0.016]***
Herf (-2)*Interbank Rate (-2)	0.292 [0.132]**	
Herf (-3)*Interbank Rate (-3)	0.107 [0.021]***	0.053 [0.015]***
Herf (-4)*Interbank Rate (-4)	-0.030 [0.014]**	0.014 [0.002]***
Herf (-5)*Interbank Rate (-5)	-0.355 [0.104]***	0.272 [0.093]***
Herf (-1)*Nominal Rate 30ds (-1)	0.327 [0.137]**	
Herf (-2)*Nominal Rate 30ds (-2)	-0.361 [0.146]**	-0.033 [0.015]**
Herf (-4)*Nominal Rate 30ds (-4)	0.007 [0.003]**	-0.015 [0.003]***
Herf (-5)*Nominal Rate 30ds (-5)	0.536 [0.116]***	-0.197 [0.103]**
Herf (-6)*Nominal Rate 30ds (-6)	-0.076 [0.015]***	-0.046 [0.012]***
Risk (-5)*Interbank Rate(-5)	0.810 [0.431]*	2.653 [1.091]**
Risk (-5)*Nominal Rate 30ds(-5)	-1.493 [0.572]***	-3.376 [1.322]**
Risk (-6)*Nominal Rate 30ds(-6)	0.614 [0.257]**	0.675 [0.376]*
Solvency *Interbank Rate	-0.286 [0.081]***	-0.379 [0.125]***
Solvency (-1)*Interbank Rate (-1)	0.152 [0.060]**	0.296 [0.0912]***
Solvency (-3)*Interbank Rate (-3)	-0.209 [0.095]**	-0.451 [0.142]***
Solvency (-3)*Nominal Rate 30ds(-3)	0.230 [0.131]*	0.487 [0.201]**
Market Share (-3)*Interbank Rate(-3)	-0.008 [0.004]**	-0.012 [0.003]***
Market Share (-3)*Nominal Rate 30ds(-3)	0.007 [0.004]*	0.012 [0.004]***
Market Share (-5)*Nominal Rate 30ds(-5)	0.005 [0.002]***	0.005 [0.001]***
Market Share (-6)*Nominal Rate 30ds(-6)	-0.003 [0.002]**	-0.004 [0.001]***

(continued)

NOMINAL *(conclude)*

	[1]	*[2]*
R-squared	0.9620	0.972
S.E. of regression	0.063	0.053
Log likelihood	1755.7	1876.3

Standard deviation in brackets in model [2] we control for year 1998.

* Significant at 10%; ** significant at 5%; *** significant at 1%.

UF RATE 90 DAYS TO 1 YEAR

	[1]	*[2]*
PRBC	2.735 [0.573]***	1.861 [0.313]***
PRBC (-1)	-12.119 [3.563]***	
PRBC (-2)	15.041 [3.765]***	
PRBC (-3)	-5.580 [2.347]**	-6.901 [2.324]***
PRBC (-6)	0.171 [0.042]***	-0.051 [0.018]***
UF 90 ds 1yr (-1)	13.446 [3.758]***	
UF 90 ds 1yr (-2)	-16.297 [3.999]***	
UF 90 ds 1yr (-3)	6.305 [2.471]**	7.411 [2.444]***
UF 90 ds 1yr (-4)	0.296 [0.093]***	
UF 90 ds 1yr (-5)	0.194 [0.102]*	
UF 90 ds 1yr (-6)	-0.440 [0.113]***	
dtpm	-0.264 [0.086]***	
Herf	1.143 [0.356]***	0.721 [0.212]***
Herf*PRBC	-0.214 [0.064]***	-0.121 [0.035]***
Herf (-1)*PRBC(-1)	1.362 [0.404]***	
Herf (-2)*PRBC (-2)	-1.731 [0.427]***	-0.021 [0.007]***
Herf (-3)*PRBC (-3)	0.626 [0.266]**	0.777 [0.262]***
Herf (-4)*PRBC (-4)	0.015 [0.004]***	
Herf (-1)*UF 90 ds 1yr (-1)	-1.467 [0.425]***	0.026 [0.003]***
Herf (-2)*UF 90 ds 1yr (-2)	1.850 [0.451]***	0.017 [0.008]**
Herf (-3)*UF 90 ds 1yr (-3)	-0.701 [0.279]**	-0.831 [0.275]***
Herf (-4)*UF 90 ds 1yr (-4)	-0.049 [0.012]***	
Herf (-5)*UF 90 ds 1yr (-5)	-0.024 [0.011]**	-0.002 [0.001]***
Herf (-6)*UF 90 ds 1yr (-6)	0.031 [0.011]***	0.010 [0.002]***
Risk *PRBC	-1.934 [0.660]***	-1.593 [0.609]***
Risk (-6)*UF 90 ds 1yr (-6)	-0.789 [0.289]***	
Risk	17.154 [4.334]***	11.954 [3.629]***
Solvency (-1) *PRBC(-1)	0.883 [0.305]***	0.816 [0.254]***
Solvency (-2)*PRBC (-2)	1.185 [0.364]***	1.053 [0.393]***
Solvency (-1)*UF 90 ds 1yr (-1)	-1.637 [0.399]***	-1.327 [0.341]***
Solvency (-2)*UF 90 ds 1yr (-2)	-1.066 [0.382]***	-0.986 [0.407]**

(continued)

UF RATE *(conclude)*

	[1]	*[2]*
Market Share	0.053 [0.026]**	0.040 [0.009]***
Liquidity (-2)*PRBC(-2)	0.005 [0.000]**	
R-squared	0.977	0.986
S.E. of regression	0.340	0.255
Log likelihood	-315.4	-32.5

In model [2] we control for year 1998.
* Significant at 10%; ** significant at 5%; *** significant at 1%.

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Modelling and forecasting exchange rate dynamics in Jamaica: an application of asymmetric volatility models

1. INTRODUCTION

Given the small, open and import dependent nature of the Jamaican economy, the exchange rate is probably the most important asset price. It has been found to be an important element in the monetary transmission process in Jamaica[1] and movement in this price has a significant pass-through to consumer prices.[2] Against this background, understanding and forecasting exchange rate behaviour is important to monetary policy. More importantly, because of the thinness and related volatility of the market, policy makers have to take into account the information content of short-term volatility. That is, while the medium to

[1] See Robinson and Robinson (1997), Allen and Robinson (2004).
[2] See Robinson (2000a and 2000b) and McFarlane (2002).

Paper prepared by Rohan Longmore and Wayne Robinson of the Research Services Department, Bank of Jamaica. This paper was published as Working Paper WP2004/03, Bank of Jamaica and presented at the IX Meeting of the Central Bank Researchers Network held in San José, Costa Rica, November 1-3, 2004. The views expressed are those of the authors and do not necessarily reflect those of the Bank of Jamaica.

long-term outlook is important, policy makers are also concerned about exchange rate movements in the very short run in deciding intervention policy. Despite this, very few studies have attempted to model and forecast exchange rate dynamics in Jamaica, particularly in the very short run. This paper attempts to fill this gap.

Traditionally, exchange rates have been explained largely by macroeconomic variables or economic fundamentals. Market fundamentals are economic variables that models with rational behaviour predict are the determinants of asset prices. In the flexible price monetary model for example,[3] the relative supplies of domestic and foreign currency determine the exchange rate. In this model, prices adjust so that PPP holds and there is equilibrium in the goods market. Domestic and foreign currency assets are perfect substitutes and with perfect capital mobility, Uncovered Interest Parity (UIP) holds. Dornbusch (1976) introduced sticky prices and showed that unanticipated monetary disturbances lead to exchange rate overshooting. Relaxing the assumption of perfect substitutability of assets and incorporating non-monetary assets, the portfolio balance approach shows how the equilibrium allocation of wealth is determined by exchange and interest rates. More recent models allow for the fact that domestic agents do hold foreign currency. In this setting, the greater the rate of substitution between both currencies, the greater the deviation of expected future exchange rate from the spot rate.

However, following the seminal paper of Meese and Rogoff (1983), the ability of economic fundamentals to predict exchange rates has been questioned, particularly in the short run.[4] In particular, the random walk model was found to out-perform the traditional structural models. This suggests that the exchange rate is influenced by factors other than macroeconomic fundamentals, and as such, these models would not adequately explain the short run characteristics of asset markets.

In seeking to address the shortcomings of the structural approach in explaining the dynamics of exchange rates, there is a growing body of literature that focuses on the market microstruc-

[3] See Frenkel (1976) and Mussa (1976).

[4] Meese and Rogoff (1983) concluded that for some assets, in particular, exchange rate, prices and fundamentals are largely disconnected. Mark (1995) suggests that fundamentals have significant explanatory power only over long horizons.

ture with the use of high-frequency data. This approach encapsulates issues relating to information asymmetries, heterogeneity of participants and market configurations.

This paper adopts an eclectic approach in that it incorporates both market microstructure, as well as, macroeconomic fundamentals, namely changes in the supply of high-powered money, in modelling and forecasting Jamaica's exchange rate at the daily frequency. The closest work to that being undertaken was done by Walker (2002), who focused only on a limited set of microstructure variables in a linear GARCH model. This paper, however, accounts explicitly for the asymmetric response of the market to shocks and the significant leptokurtosis in the exchange rate. This, as it is known that the standard linear GARCH model is inadequate in the presence of these factors and, as such, this paper employs nonlinear asymmetric GARCH models. We also revisit the mixture of distribution of hypothesis theorem, which explores the relationship between volume and volatility[5] and assess the significance of co-volatility across markets. Finally, we test the forecasting performance of the various models over a 30-day ahead horizon.

The paper proceeds as follows. Empirical regularities of assets returns and the characteristics of the Jamaican foreign exchange market are discussed in section 2. A brief review of the models of time varying heteroscedasticity is presented in section 3, followed by the estimation results in section 4. The forecasting performance of the models is assessed in section 5, followed by some concluding comments in section 6.

2. EMPIRICAL CHARACTERISTICS OF THE JAMAICAN FOREIGN EXCHANGE MARKET

Asset markets are known to possess some common features or regularities. While these features are mostly found in stock markets, there is some evidence that they are also present in foreign exchange markets. This section examines the presence of such regularities in the Jamaican foreign exchange market, which is relatively small and less developed. We first discuss these common features and the international evidence and then examine whether they are present in the Jamaican data. This will then determine the relevant statistical distribution and model for the exchange rate.

[5] See Walker (2002) and Galati (2000).

2.1 Empirical Regularities

2.1.1 Volatility Clustering and Non-Normality

The variance of speculative prices or asset returns is not constant over time. In fact, asset prices are commonly characterised by volatility clustering[6] – large changes followed by large changes and small changes followed by small changes.[7] While some researchers suggest that this phenomenon is a result of speculative activities, others attribute it to uncertainty or risk. Uncertainty about future market fundamentals, associated with current and expected fiscal and monetary policies, as well as corporate decisions, generate bouts of increased volatility.[8]

More importantly, Gokcan (2000) suggested that volatility is related to the stage of market development. Risk or the uncertainty of returns in emerging markets is typically higher than those in developed markets. As such, volatility in emerging markets is generally larger and more persistent than in developed markets. One explanation is the difference in the speed and reliability of information available to investors, which is associated with modes of telecommunication and possibly the accounting system in place. As such, a small and relatively under developed foreign exchange market such as Jamaica's should exhibit relatively larger volatility clustering.

Volatility clustering or non-constant variance gives rise to thick tails or leptokurtosis.[9] This, as if the unconditional kurtosis of the innovations, ε_t, is finite then the moment condition $E(\varepsilon_t^4)/E(\sigma_{\varepsilon t}^2)^2 \geq E(z_t^4)$ holds with strict equality only if σ_t is constant. The presence of excess kurtosis or thick tails in asset returns implies that estimations based on the assumption of identical and independently distributed (i.i.d.) errors are inappropriate for asset returns. Further, there is strong evidence in the finance literature linking volatility in asset returns with higher order serial correlation.[10] Against this background, the empirical distribution of asset returns is typically highly non-normal.[11]

[6] Mandelbrot (1963)

[7] This is one feature of asset returns which structural models fail to capture.

[8] Connolly and Stivers (1999) argue that volatility clustering may not be due to 'news' but heterogeneity of beliefs.

[9] Bollerslev, Engle and Nelson (1994).

[10] See Kim (1989).

[11] See Fama (1965), Kim and Kon (1994).

2.1.2 Non-trading Periods, Daily Seasonality and Regular Events

It is widely accepted that information accumulates during periods when financial markets are closed and is reflected in the price when the markets reopen. As such variances are higher following weekends and holidays. This leads to the observation of daily seasonality in asset returns also known as the days-of-the-week effect (DOW). The release of important information is also found to be associated with high volatility. Harvey and Huang (1992), for example, find that foreign exchange volatility is higher when there is news of heavy central bank trading or there is a release of macroeconomic news. Patell and Wolfson (1979) document similar evidence for the stock market. The pattern of volatility during the trading day is also found to be predictable, in that, volatility is typically higher at the open and close of trading in both stock and foreign exchange markets.[12]

The impact of news, as well as the magnitude of the DOW effect depend on the efficiency of the market, which is a function of the level of market development. Thus *a priori*, DOW effects may not be as significant in the Jamaican market.

2.1.3 Asymmetry and Leverage Effects

It is known that the magnitude of the response of asset prices to shocks depends on whether the shock is negative or positive. Black (1976) attributes asymmetry to a *leverage effect*. That is, when a stock price falls, the value of the associated company's equity declines and, as such, its leverage or the debt to equity ratio rises. The consequence of which is a perception of increased risk that translates into higher volatility. In this context, negative surprises increase predictable volatility in asset markets more than positive surprises.

Another explanation of asymmetry is the *volatility feedback hypothesis.*[13] This was developed to explain stock price volatility. A negative shock to volatility increases the future risk premia. This would cause the stock price to fall if the future dividends are expected to remain the same. When applied to the foreign exchange market, a shock, which increases the volatility of the market, increases the risk of holding the currency. This induces a

[12] See Baillie and Bollerslev (1992).
[13] See Campbell and Hentschel (1992).

portfolio shift out of the currency, leading to a depreciation of the exchange rate.

While asymmetry has been found in stock returns there is very little evidence that it exists in foreign exchange returns.[14] This could be due to the fact that such studies generally focus on highly developed markets. In the case of a thin market such as Jamaica's, adverse shocks can have more persistent effects than positive shocks.

2.1.4 Co-movement in Volatility

The earliest observation of correlation in volatility can be found in Black (1976) for stock markets. Harvey et al (1992) provide similar evidence for stock markets and Engel et al (1990) for the US bond market. The co-movement in volatility not only holds across different assets within a market but also across markets. For example, King et al (1994) and others have found co-movement across international markets. Co-movement in volatility reflect the extent to which investors consider assets in different markets as substitutes and the availability of relevant information. Most market analysts regard the substitution effect as the overriding condition. That is, if assets in separate markets are considered as substitutes by an investor, new information in either market may necessitate portfolio adjustments to minimize risk exposure.

The degree of co-volatility would be lower in lesser-developed financial systems where stock and bond markets are less liquid and secondary trading is low. In this context the ability of investors to capitalize on arbitrage opportunities would be limited.

2.1.5 Market Microstructure and Volatility

Following the failure of macroeconomic variables and fundamentals to adequately explain volatility in speculative prices, microstructure variables have been identified as factors that can explain volatility. For the foreign exchange market, thinness indicated by volumes traded and the bid-ask spreads, market fragmentation measured by the degree of concentration, heterogeneous expectations and volatility spill over from other markets have been identified as some of the main explanatory variables.[15]

[14] See for example Engle and Ng (1990) and Kisinbay (2003).

[15] See Galati (2000) and Walker (2002) for an exposition on these relationships.

Empirical tests using microstructure variables find a positive relationship between volatility and spreads and volumes traded[16]. Spreads reflect the cost of transacting in foreign currency and as such will fluctuate in the same direction as risk, which is reflected in volatility. Volume and volatility are said to be influenced by the same process of information arrival and as such should be positively correlated.[17] However, there is the view that the relationship between volumes and volatility will depend on whether the market is fully developed as against an emerging market. In this context, Tauchen and Pitts (1983) argue that the relationship can be negative.

A highly concentrated market is likely to exhibit more volatility. As a market becomes more concentrated, the greater the likelihood of the action of one investor influencing prices. This suggests a positive relationship between market concentration and volatility.

2.2 The J$/US$ Exchange Rate

We examine the daily average dollar spot exchange rate of Jamaica[18] over the period 2 January 1998 to 12 February 2003, a total of 1280 trading days after removing weekends and holidays. The plots of the autocorrelation function (ACF) and the partial autocorrelation function (PACF) shown in figure 1 in the appendix suggest that the series is non-stationary. The correlogram dies out very slowly indicative of a long memory, or long-term dependence process. This suggests that there is a persistent temporal dependency between observations over various displacements. The presence of a long memory process in the exchange rate series indicate that shocks to the exchange rate have some amount of permanence and the mean and variance of the series are time dependent.[19]

To test for stationarity we use the Phillips and Perron (1988) test for unit root, which involves estimating the test regression.

[16] See Clark (1973) and Frankel and Froot (1990).

[17] This relationship describes what Clark (1973) posits as the mixture of distribution hypothesis.

[18] Dollar exchange rate is defined as units of notional currency per US dollar.

[19] The presence of long memory process in the asset prices contradicts the weak form of market efficiency, which stipulates that, conditioned on historical information, future returns or movements in prices are unpredictable.

$$s_t = \mu + \beta(t - \mathrm{T}/2) + \alpha s_{t-1} + \varepsilon_t$$

where s_t is the log of the exchange rate and ε_t the innovation. The hypotheses are $H_o^1 : \alpha = 1$ and $H_o^2 : \beta = 0, \alpha = 1$ which are tested using the $Z(t_a)$ and $Z(\Phi_3)$.[20] The computed test statistics are -1.24 and 1.09 for the $Z(t_a)$ and $Z(\Phi_3)$, respectively, against the 5% critical values of –3.41 and 6.25, respectively. Thus the unit root hypothesis cannot be rejected.

It is known, however, that these tests do not account for the presence of near unit roots or a long memory process. That is, these tests cannot detect an order of integration, d, which is less than unity. In which case the series is said to be fractionally integrated and has to be differenced by $d<1$ times to be made stationary. As such, we also employ the Geweke and Porter-Hudak (1983) spectral regression test for fractional integration.

A time series **y** follows and autoregressive fractionally integrated moving average process of order (p,d,q) (ARFIMA(p,d,q)) if

$$\Phi(L)(1-L)^d(y_t - \delta) = \Theta(L)\varepsilon_t, \quad \varepsilon_t \sim iid(0, \sigma_3^2) \tag{1}$$

where L is the lag operator, $\Phi(L) = 1 - \phi_1 L - \ldots - \phi_p L^p, \Theta(L) = 1 + \vartheta_1 L + \ldots + \vartheta_q L^q$, and $(1-L)^d$ is the fractional differencing operator defined by

$$(1-L)^d = \sum_{k=0}^{\infty} \frac{\Gamma(k-d)L^k}{\Gamma(-d)\Gamma(k+1)} \tag{2}$$

with $\Gamma(\cdot)$ denoting the gamma, or generalized factorial function. The stochastic process **y** is both stationary and invertible if all roots of $\Phi(L)$ and $\Theta(L)$ lie outside the unit circle and $|d| < 0.5$. When $d \in (-0.5, 0.5)$ and $d \neq 0$ the autocorrelation function of an ARFIMA process decays hyperbolically to zero as $j \to \infty$, at a much slower rate than the exponential decay of a stationary ARMA process (i.e., $d=0$).

Geweke and Porter-Hudak (1983) suggested a semi-parametric procedure to obtain an estimate of the fractional differencing parameter, d, based on the slope of the spectral density function around the angular frequency $\xi = 0$. More specifically, let $I(\xi)$ be the periodogram of **y** at frequency ξ defined by

$$I(\xi) = \frac{1}{2\pi T}\left|\sum_{t=1}^{T} e^{it\xi}(y_t - \bar{y})\right|^2. \tag{3}$$

[20] See Phillips and Perron (1988) for the precise form of these statistics.

Then the spectral regression is

$$1n\{I(\xi\lambda)\} = \beta_0 + \beta_1 1n\left\{sin^2\left(\frac{\xi_\lambda}{2}\right)\right\} + \eta_{\lambda,} \quad \lambda = 1,\ldots,v \tag{4}$$

where $\xi_\lambda = \frac{2\pi\lambda}{T}(\lambda = 0,\ldots T-1)$ denotes the harmonic ordinates of the sample, T is the number of observations, and $v = g(T) << T$ is the number of harmonic ordinates included in the spectral regression. The OLS estimate of the slope coefficient in (4) provides an estimate of *d*. The plot of the estimates of *d* is shown in Figure 2 in the appendix. The estimates appear to settle down around 0.83, thereby supporting the hypothesis of fractional integration.

The immediate implication of these results is that the exchange rate follows a random walk and hence the use of a martingale or random walk model for short-run exchange rate movements. A random walk model of exchange rate changes, however, may not be appropriate in the presence of conditional heteroscedasticity. In this context we study the properties of the estimated innovations from a simple random walk model with drift. We use both the first difference, which corresponds to a single unit root and the fractional differenced series. With respect to the latter, the series was differenced 0.83 times using the binomial operator in equation 2. The plots of both differenced series are shown in Figure 3 in the appendix and the relevant statistics in Table 1.

TABLE 1. SUMMARY STATISTIC FOR THE RETURN SERIES

Series	*Mean*	*Variance (%)*	*Skewness*	*Excess Kurtosis*	*Jarque-Bera Normality Test*[1]	*Ljung-Box Test Q(40)*	*Ljung-Box Test Q^2 (40)*
1st difference	0.000267	0.0004	0.696	8.38	1332.031 (P = 0.0)	(P = 0.000)	(P = 0.000)
fractional diff.	0.0047	0.0014	1.773	4.11	1091.1 (P = 0.0)	(P = 0.000)	(P = 0.000)

From Figure 3, the return series exhibits bouts of intense volatility followed by periods of tranquillity, which is consistent with the volatility-clustering hypothesis in the finance literature. Periods of intense volatility are followed by further observations of high volatility. The volatility in the Jamaican foreign exchange market could be a reflection of the stage of market development as Gokcan (2000) suggests, in combination with a latent perception of risk or low confidence, related to weak economic fundamentals.

Although the number of players, the depth of the market and consequently the degree of competition has increased over the sample period, relatively few large players largely dominate the market. Further, market trades are segmented between large contract transactions and smaller over the counter trades. The trading infrastructure and the vehicle for information dissemination is in an early stage of development and as such doesn't permit an efficient dissemination of information among traders and end users. Thus there may well be an autocorrelated news-generating process from time to time in the market.[21]

Within a thin market, any negative news or fad, however generated, can generate large swings in a context where the perceived risk of holding Jamaica Dollars may (or may not) be higher than what is warranted by fundamentals. The level of uncertainty reflects the effect of historical shocks, particularly in the early part of the 1990s, which from the unit root analysis, tend to be long lived.

The statistics in Table 1 indicate the presence of significant excess kurtosis, particularly for the first differenced series. In fact, the excess kurtosis is larger than that found for the currencies of countries with developed markets in Baillie and Bollerslev (1992). The Ljung and Box (1978) test statistic for the *k*th-order serial correlation suggest significant higher order autocorrelation. The distribution is significantly skewed towards the left, particularly for the fractionally differenced series. The Ljung and Box $Q^2(k)$ statistics reject the hypothesis of conditional homoscedasticity. The kernel plots (Figures 4a and 4b) of the unconditional distribution confirm that the unconditional distribution of the return series is not Gaussian.

In summary, the preceding analysis indicates that the empirical distribution of returns in the foreign exchange market is non-normal, with very thick tails. The leptokurtosis reflects the fact that the market is characterised by very frequent medium or large changes. These changes occur with greater frequency than what is predicted by the normal distribution. The empirical distribution confirms the presence of a non-constant variance or volatility clustering. The degree of leptokurtosis when compared to other markets may be reflective of the thinness of the Jamaican

[21] See Walker (2002) for further details.

foreign exchange market, where small movements or shocks tend to get magnified overtime. It may also reflect uncertainty regarding Government policies and other economic fundamentals

The presence of long memory implies that shocks are long lived. The more significant result, however, is that of the asymmetry of the distribution, which implies that positive shocks (i.e. shocks that lead to a depreciation) are more likely than negative shocks. The response of the market may, however, differ depending on the nature of the shock (i.e. positive or negative). The thick left tail means that the density of a return distribution is asymmetric with a sharper fall on the right tail, which is an indication of the fact that market declines occur with greater frequency than increases. In this context, the risk of holding long position in foreign exchange is relatively small. In other words, short positions are relatively more expensive.

3. MODELS OF TIME VARYING CONDITIONAL HETEROSCEDASTICITY

This section presents an overview of the models used, given the statistical properties of exchange rate returns found above, namely asymmetry and significant fat tails. Generally, time varying heteroscedasticity is modelled by the linear GARCH model of Bollerslev (1986) i.e.

$$\Delta s_t = b_0 + \varepsilon_t , \qquad \varepsilon_t \big| \Omega_{t-1} \sim D(0, h_t)$$

where,

$$h_t = w + \sum_{i=1}^{p} \phi_j \varepsilon^2{}_{t-i} + \sum_{j=1}^{q} \beta_j h_{t-j} \tag{5}$$

where w, ϕ_j and β_j are constant and non-negative parameters. This specification allows for the conditional variance to be dependent on past information, which will induce variability over time. More specifically, the conditional variance is explained by past shocks and past variances.[22] The key features of this specification are that if $p=0$, the process reduces to an ARCH (q) process

[22] Engle and Ng (1991) examined the implied relationship between past errors and the conditional variance. The graphical representation of this relationship is termed the news impact curve. The exact shape of this curve is dependent on the specification of h_t.

and ε_t a white noise process when $p = q = 0$. To ensure stationarity and to prevent negative variances the restriction $\sum_{i=1}^{p}\phi_j + \sum_{j=1}^{q}\beta_j < 1$[23] must hold.

Generally, the linear GARCH (*p, q*) model, based on the conditional normal distribution, captures thick tails and other stylised facts such as non-trading periods and regular events. Notwithstanding the apparent success of linear GARCH models, Engle and Ng (1991), Bollerslev, Chou and Kroner (1992) and other leading researchers have suggested that there are features of the data that this model cannot capture. For example, it doesn't always account for significant fat tailedness in the unconditional distribution.[24] Further, it is found to be deficient in correcting bias in the forecast and forecast error variance associated with a skewed distribution. Against this background, there have been a number of extensions to the GARCH (*p, q*) model to explicitly account for skewness. We have restricted our analysis to the more popular models of asymmetric volatility. These include, the exponential GARCH (EGARCH) model, Glosten, Jogannathan, and Rankle (1992) GJR-GARCH model, asymmetric power ARCH (APARCH), Zakoian (1994) threshold ARCH (TARCH). The TS-GARCH advanced by Taylor (1986) and Schwert (1990), the t-GARCH and the generalized version of Higgins and Bera (1992) non-linear ARCH (NGARCH) are included to capture the information content within the thick tails of the return distribution. Given the excess kurtosis in the returns series a simple GARCH would be inappropriate and as such we considered only the linear models, which account for this feature.

3.1 Models of Fat Tails

3.1.1 t-GARCH

The most common approach used when the error distribution tends to have significantly fatter tails than the normal distribution is to adopt the Student-*t* distribution. This gives rise to the t-GARCH model where the degrees of freedom are also estimated. The functional form of h_t remains the same, however, the normal density function in the log likelihood is replaced by the Student-*t*.

[23] See Bollerslev (1986) for a comprehensive discussion on the need for these restrictions.

[24] See for example Baillie and Bollerslev (1989) and Hseih (1988).

3.1.2 NGARCH

The NGARCH model is a generalization of the Higgins and Bera (1992) non-linear ARCH, which only contained ARCH lags. The NGARCH has the following structure:

$$h_t = w + \sum_{i=1}^{p} \phi\left(\varepsilon_{t-i} + \gamma h_{t-1}^{1/2}\right)^2 + \sum_{j=1}^{q} \beta_j h_{t-j} \quad (6)$$

In this model the asymmetric effect depend upon the standard deviation.

3.1.3 TS-GARCH

The TS-GARCH model developed by Taylor (1986) and Schwert (1990) is another popular model used to capture the information content in the thick tails, which is common in the return distribution of speculative prices. The specification of this model is based on standard deviations and is as follows:

$$\left(h_t^{1/2}\right) = w + \sum_{i=1}^{p} \left(\phi_j \left|\varepsilon_{t-i}\right|\right) + \sum_{j=1}^{q} \beta_j (h_{t-j}^{1/2}) \quad (7)$$

3.2 Non-linear GARCH models

3.2.1 EGARCH

The exponential GARCH (EGARCH) model advanced by Nelson (1991) is the earliest extension of the GARCH model that incorporates asymmetric effects in returns. The variance component, h_t, is an asymmetric function of past ε_t's and is defined as follows:

$$log\ (h_t) = w + \sum_{i=1}^{p} \phi_j \left[\frac{\left|\varepsilon_{t-i}\right|}{\sqrt{h_{t-i}}} - \sqrt{2/\pi} \right] + \gamma \frac{\varepsilon_{t-1}}{\sqrt{h_{t-i}}} + \sum_{j=1}^{q} \beta_j \log h_{t-j} \quad (8)$$

where w, ϕ_j, γ and β_j are constant parameters. Unlike the GARCH (p, q) model, the form of the EGARCH (p, q) equation indicates that the conditional variance is an exponential function, thereby removing the need for restrictions on the parameters to ensure positive conditional variance. The asymmetric effect of past shocks is captured by the γ coefficient, which is usually negative, that is, *ceteris paribus*, positive shocks generate less volatility than negative shocks. This feature permits the capture of the sign effect by allowing positive and negative innovations to have different effects on the volatility. If $\gamma = 0$, positive and negative shocks have the same effect on volatility. The size effect is captured by ϕ_j and is expected to be positive. Shocks are measured relative to its standard deviations. The use of absolute shocks and logs in this parameterisation allows us to capture the size effect, in that it increases the impact of large shocks on the next period conditional variance.

3.2.2 GJR-GARCH

The GJR-GARCH (*p, q*) model is another volatility model that allows for asymmetric effects. This was introduced by Glosten, Jagannathan and Runkle (1993). The general specification of this model is of the form:

$$h_t = w + \sum_{i=1}^{p}\left(\phi_j \varepsilon^2{}_{t-i} + \gamma S^-_{t-i}\varepsilon^2{}_{t-i}\right) + \sum_{j=1}^{q}\beta_j h_{t-j} \tag{9}$$

where the only difference from the general GARCH (*p, q*) model is $\sum_{i=1}^{p}\left(\gamma S_{t-i}\varepsilon^2{}_{t-i}\right)$. S^-_t is a dummy variable that is equal to 1 if $\varepsilon_{t-i} < 0$ and zero otherwise. It is this extra term, which allows for the asymmetric effect, as the impact of ε^2_{t-i} on h_t depends on whether the shock is negative or positive. In the event that a negative shock is realized, then the impact on volatility will be $\phi_j + \gamma$ and ϕ_j when the shock is positive. The parameter γ, which captures the asymmetric effect, is expected to be positive.

3.2.3 APARCH

The asymmetry power ARCH (APARCH) model of Ding, Granger and Engle (1993) allows for asymmetric effects of shocks on the conditional variance in a more general framework. Formally, the general specification of the APARCH (*p, q*) model is:

$$\left(h_t^{1/2}\right)^{\delta} = w + \sum_{i=1}^{p}\left(\phi_j\left|\varepsilon_{t-i}\right| - \gamma_i\varepsilon_{t-i}\right)^{\delta} + \sum_{j=1}^{q}\beta_j (h_{t-j}^{1/2})^{\delta} \tag{10}$$

where *w*, ϕ_j, δ and β_j are constant, positive parameters, and $-1 < \gamma_i < 1$.

Asymmetry or leverage effect in this model is captured by the γ_i term. For an APARCH (1,1) model, when $\gamma > 0$, negative shocks lead to higher volatility and vice versa. APARCH (*p, q*) models of asymmetric differs from other GARCH type volatility models with the introduction of the power term, δ, which is to be estimated. The introduction and estimation of the power term is an attempt to account for the true distribution underlying volatility. The idea behind the introduction of a power term arose from the fact that in modelling financial data, the assumption of normality, which restricts δ to either 1 or 2, is often unrealistic due to significant skewness and kurtosis.[25] Allowing δ to take the form of a free parameter to be estimated, removes this arbitrary restriction.

3.2.4 TARCH

The threshold ARCH or TARCH (*p*, *q*) model introduced by Zakoian (1994) is very similar to the GJR-GARCH model with the exception that it is the conditional standard deviation that is modelled and not the conditional variance. In this regard, the parameterisation is similar to equation (5) with the difference being that h_t is replaced with the square root of h_t.

3.3 Model Selection

Selecting the most appropriate model, particularly when the true underlying distribution is unknown can be complicated. Most researchers tend to rely on the traditional Akaike and Schwartz information criteria. However, the statistical properties and hence reliability of these information criteria are unknown in the context of time varying volatility. As such, model selection criteria typically focus on the estimation of loss functions for alternative models. When applied to models with time varying volatility such loss functions depend on the squared residuals and the variance. This paper uses three such measures:

$$L_2 = \sum_{t=1}^{T} (\varepsilon_t^2 - \sigma_t^2)\sigma_t^{-4}$$

$$L_3 = \sum_{t=1}^{T} (ln(\sigma_t^2) + \varepsilon_t^2 \sigma_t^{-2})$$

and

$$L_4 = \sum_{t=1}^{T} (ln(\varepsilon_t^2 \sigma_t^{-2}))^2 .$$

[25] Under the assumption of normality, the entire distribution can be defined by the first two moments. That is, δ is either 1 or 2. When $\delta = 1$, the APARCH model evaluates the standard deviation. $\delta = 2$, the variance is modelled.

3.4 Co-Volatility

Multivariate GARCH models, first introduced by Kraft and Engle (1983), allows for co-movements between asset returns. Subsequent studies have advanced various modifications to the original specifications. The more popular models include the VECH model by Bollerslev, Engle and Wooldridge (1988), the CCC model of Bollerslev (1990), Engle, Ng and Rothschild (1990) factor model and the BEKK model of Engle and Kroner (1995). The basic intuition behind all these models is that conditional co-variance matrix H_t of the ε_t's is dependent on past information (Ω_{t-1}). Where the models differ is in relation to the parameterisation of H_t.

This paper uses the BEKK multivariate framework, which is given as:

$$\mu_t = E(y_t|\Omega_{t-1}) = \Gamma X_t + \varepsilon_t \qquad \varepsilon_t|\Omega_{t-1} \sim D(0, H_t)$$

and,

$$H_t = C + A\varepsilon_{t-1}\varepsilon_{t-1}^{'}A^{'} + BH_{t-1}B^{'} \tag{11}$$

where C, A and B are all N*N parameter matrices. y_t is a vector of returns on N assets during period t, and μ_t, the conditional mean vector. The BEKK allows the conditional covariance matrix to be determined by the outer product matrices of the vector past return shocks. Because most multivariate GARCH models introduce an unmanageable number of parameter, in practise, it is necessary to place restrictions on the off-diagonal elements to ensure interpretability. In the case of the BEKK model, the number of parameters is given by the following expression, ($5/2N^2 + N/2$).

4. RESULTS

4.1 Data

Due to data availability the main variable used to capture macroeconomic fundamentals was the changes in the monetary base. This is an indicator of the changes in Jamaican Dollar liquidity conditions. A more appropriate measure of changes in liquidity conditions would be the movement in the overnight or inter-bank rates. However, data was not available for the full sample period. The microstructure variables employed include, trading volumes, bid-ask spread and dummies representing each day of the week. The inclusion of volumes captures the thinness of the market.

Given the presence of long memory, the analysis focuses mainly on the results for the fractionally differenced series in Table 2. Those for the first differenced series are reported for comparative purposes in Tables 3[26] and the model selection criteria are given in Tables 4 and 5.

The results support the use of GARCH models to explain exchange dynamics.[27] The information criteria shown in the bottom panel of Tables 2and 3 suggest that the best model is the GJR. This is further supported by the model selection criteria based on the loss functions. There is also support for the TS-GARCH model, which treats with fat tails. Generally, the asymmetric GARCH models for the fractional unit root series and to a lesser extent the series with a single unit root perform best in terms of minimizing the loss functions.

With the exception of the APARCH model in Table 2, the gammas are significant, confirming the importance of asymmetry in exchange rate volatility. Only the coefficients in the GJR and the APARCH models have the correct sign, which implies that negative news lead to greater volatility in the Jamaican foreign exchange market.

We postulate that this asymmetry reflects the volatility feedback hypothesis. That is, any shock to exchange rate volatility, which itself is an indicator of risk, increases the currency risk premia (as well as the cost of international trade), which unless compensated for by a higher interest rate differential, leads to greater volatility as investors freely adjust their portfolios to hedge against currency risk.

Mean return for each day of the week is generally positive and statistically significant. The reported chi-squared statistics, which test for equality of return across the days of the week, suggest that daily mean returns are significantly different from each other. Though the deviations from Monday's returns are on average just 0.1 per cent, the evidence confirms the presence of a DOW effect in Jamaica's foreign exchange returns. The presence of DOW effect in asset returns is evidence against market efficiency, as investors are able to predict with reasonable certainty move-

[26] The NGARCH models did not converge and as such were excluded from this section.

[27] The sum of the ARCH and GARCH coefficients in the EGARCH exceeds 1 and as such the results from this model may not be reliable.

TABLE 2. PARAMETER ESTIMATES FOR THE GARCH MODELS (FRACTIONAL)

	Fat Tail Models		*Asymmetric Models*			
Variables	*TGARCH*	*TS-GARCH*	*EGARCH*	*APARCH*	*GJR-GARCH*	*TARCH*
Mean						
Constant	0.074	0.117	0.026	0.050	0.054	0.036
	(48.14)	-6.51	(21.43)	(3.38)	(32.76)	(29.37)
Tuesday	0.002	0.001	0.001	0.002	0.001	
	(18.32)	90.69)	(7.96)	(20.20)	(3.95)	
Wednesday	0.001	0.002	0.000	0.004	0.001	
	(21.93)	(1.42)	(5.09)	(39.50)	(4.76)	
Thursday	0.001	-0.001	0.000	0.004	0.001	
	(11.55)	(-0.74)	(3.83)	(39.03)	(3.26)	
Friday	0.001	-0.003	0.000	0.003	0.000	
	(7.09)	(-2.53)	(3.38)	(32.22)	(1.64)	
Dlbase{+1}	-0.035	0.007	0.002	0.012	0.009	0.003
	(-13.07)	(0.33)	(0.85)	(6.40)	(2.41)	(0.72)
Lvolumes	-0.004	-6.50	-0.001	-0.003	-0.003	-0.002
	(-47.31)	(-0.01)	(-19.61)	(-0.24)	(-31.58)	(-27.40)
Variance						
Constant	0.000	0.003	-0.079	0.016	0.000	0.000
	(5.26)	(1.71)	(-0.13)	(0.58)	(11.69)	(12.84)
ARCH (Alpha)	0.335	0.002	0.918	0.034	-0.136	0.000
	(20.23)	(4.86)	(57.47)	(0.46)	(-2.05)	(11.6)
GARCH (beta)	0.244	0.019	0.287	0.880	0.054	0.001
	(13.74)	(4.43)	(6.12)	(64.95)	(2.51)	(13.64)
Gamma			0.137	0.990	0.100	-0.851
			(5.73)	(1.20)	(3.04)	(-22.80)
Delta				0.070		
				(0.42)		
Spread	0.000	-0.014		0.074	0.000	0.000
	(14.25)	(-1.34)		(0.06)	(7.41)	(1.80)
Lvolumes	0.000	-0.001	-0.285		0.000	0.000
	(-14.37)	(-12.46)	(-6.45)		(-12.83)	(-12.86)
Lvolumes {+1}	0.000					
	(2.42)					
Lwsxr {+1}		0.003	0.984		0.000	
		(5.90)	(4.27)		(-8.22)	

(continued)

TABLE 2 *(conclude)*

	Fat Tail Models		*Asymmetric Models*			
Variables	*TGARCH*	*TS-GARCH*	*EGARCH*	*APARCH*	*GJR-GARCH*	*TARCH*
Dlbase						0.000 (0.66)
SBC	-330.37	-330.40	-330.37	2705.40	-326.90	2753.04
AIC	-349.40	-349.43	-349.42	2769.60	-347.42	2703.62
Chi-Squared (4)	427.8 [p = 0.0]	239.9 [0.0]	8056.2 [0.0]	78.26.7 [0.0]	123.9 [0.0]	
Nº Observations	1034.00	1034.00	1034.00	1034.00	1034.00	1034.00
Function Value	4946.13	3711.84	6023.77	-22649.10	5641.61	5766.78

NOTES: t-Statistics are reported in parentheses. Chi-squared values correspond to the joint F-statistic, which test whether the coefficients on the days of week are significantly different from zero. The maximum likelihood function value is also reported. {+1} represent a one period lead.

ments in the exchange rate. Market efficiency suggests that there are no ex-ante regularities in asset returns; otherwise investor can employ trading rules to earn abnormal returns.

Mean returns in Table 2 decline as the volumes traded increase. There is also evidence that the market responds to expected changes in the monetary base. Return is generally higher when the market expects an expansion in the base. This condition could be a result of speculative tendencies associated to with excess Jamaica Dollar liquidity.

A positive relationship is noted between spreads and volatility. One interpretation of this relationship is that the risk of holding US dollar, particularly in periods of uncertainty, is not easily diversified. As a means of offsetting possible losses, the spread will rise. This could be reflective of the extent to which trading in the foreign exchange market is largely concentrated in one currency. Therefore, higher spreads in volatile periods reflect compensation for additional risk faced by holders of foreign currency. The largely negative contemporaneous relationship between volumes and volatility suggest one of three things; 1) the local foreign exchange market is still in a developmental stage and as such any change in the number of participant may affect price; 2) volumes traded relative to demand dictate price variability as such the more liquid the market, the less the likelihood of extreme volatility; 3) the market is often characterised by prolonged bouts of severe volatility, which normally results in depreciation. In the latter case holders of foreign currency would benefit from lowering

TABLE 3. PARAMETER ESTIMATES FOR GARCH MODELS (1ST DIFFERENCE)

	Fat Tail Models		*Asymmetric Models*			
Variables	*TGARCH*	*TS-GARCH*	*EGARCH*	*APARCH*	*GJR-GARCH*	*TARCH*
Mean						
Constant	0.0028	0.0016	-0.0002	0.0502	-0.0002	-0.0003
	(3.15)	(0.96)	(-4.45)	(26.52)	(-3.63)	(0.23)
Tuesday	0.0008	0.0009	0.0005	0.0005	0.0005	0.0009
	(6.934)	(6.50)	(7.09)	(12.83)	(6.62)	(0.42)
Wednesday	0.0005	0.0009	0.0003	0.0003	0.0004	0.0008
	(4.71)	(7.76)	(4.38)	(7.38)	(4.23)	(0.11)
Thursday	0.0004	0.0007	0.0003	-0.0270	0.0003	0.0030
	(3.63)	(4.30)	(3.33)	(-25.23)	(2.90)	(0.01)
Friday	0.0006	0.0013	0.0003	0.0003	0.0003	0.0008
	(4.65)	(5.72)	(4.08)	(11.89)	(3.84)	(0.52)
Dlbase{+1}	0.0066	0.0026		0.0039	0.0029	0.0001
	(2.24)	(0.74)		(4.49)	(1.73)	(4.61)
Dlbase			0.0019			
			(1.16)			
Lvolumes	-0.0002	-0.0001	0.0003	-0.0005	0.0003	-0.0002
	(-3.59)	(-0.85)	(3.07)	(-21.77)	(2.53)	(-0.22)
Variance						
Constant	0.0000	0.0002	-3.4495	-0.0255	0.0000	0.0116
	(0.88)	(0.53)	(-3.66)	(-1.88)	(1.13)	(4.59)
ARCH(Alpha)	0.2756	0.0061	0.8298	0.0162	0.1479	0.0000
	(8.22)	(7.71)	(42.27)	(1.39)	(2.13)	(-0.20)
GARCH (beta)	0.4808	0.0266	0.5275	0.8697	0.7556	0.0040
	(12.82)	(7.60)	(13.60)	(68.53)	(40.49)	(0.72)
Gamma		0.0140	-0.0089	-0.5712	0.0634	
		(0.003)	(-0.32)	(-4.43)	(1.59)	
Delta				0.1971		
				(3.63)		
Spread	0.0001	0.0259	67.3464	0.8413		0.0015
	(5.82)	(2.84)	(5.60)	(2.19)		(0.22)
Lvolumes	0.0000	0.0002	-0.4820	-0.0041	0.0000	-0.0001
	(-8.12)	(5.93)	(-7.42)	(-2.12)	(-3.35)	(-1.25)
Lvolumes {+1}	0.0000					
	(-4.10)					
Lwsxr {+1}	0.0000	-0.0010	2.3422	0.0227	0.0000	-0.0029
	(10.78)	(-4.39)	(6.48)	(2.23)	(2.38)	(-6.39)

(continued)

TABLE 3 *(conclude)*

	Fat Tail Models		*Asymmetric Models*			
Variables	*TGARCH*	*TS-GARCH*	*EGARCH*	*APARCH*	*GJR-GARCH*	*TARCH*
SBC	-330.18	-333.65	-331.32	-326.72	-330.37	-330.18
AIC	-350.70	-352.70	-351.80	-348.70	-350.70	-350.70
Chi-Squared (4)	145.3 [p=0.0]	87.81 [0.0]	2679.7 [0.0]	0.75 [0.9]	0.00001 [1.0]	0.22 [0.9]
Nº Observations	1034.00	1034.00	1034.00	1034.00	1034.00	1034.00
Function Value	5946.95	3825.76	6368.75	6338.47	6333.29	509.00

NOTES: t-Statistics are reported in parentheses. Chi-squared values correspond to the joint F-statistic, which test whether the coefficients on the days of week are significantly different from zero. The maximum likelihood function value is also reported. {+1} represent a one period lead.

TABLE 4. MODEL SELECTION: FRACTIONAL MODELS

	Fat Tail Models		*Asymmetric Models*			
Loss Function	*TGARCH*	*TS-GARCH*	*EGARCH*	*GJR-GARCH*	*TARCH*	*APARCH*
L_2	2669.67	999.71	1064.33	880.09	1028.80	1989.47
L_3	-9724.66	-7434.43	-12304.00	-12796.10	-6860.16	-10728.45
L_4	39288.29	34452.77	18206.21	12682.29	66762.03	N/A

volumes traded in the volatile period. High volatility may also discourage existing investors and new entrants to the market and as such may result in lower volume.[28] Expectation regarding future volumes are important for volatility, however, signs differ according to the assumption about unit roots.

The results in Tables 2 and 3 suggest that participants in the foreign exchange market build into current trading, expectations regarding the future level of the exchange rate. This coefficient is generally positive which suggests that as the market approaches an

TABLE 5. MODEL SELECTION: 1ST DIFFERENCE MODELS

	Fat Tail Models		*Asymmetric Models*			
Loss Function	*TGARCH*	*TS-GARCH*	*EGARCH*	*GJR-GARCH*	*TARCH*	*APARCH*
L_2	3245.81	833.00	1023.45	768.48	1033.79	1445.93
L_3	-13037.60	-11946.00	-11977.50	-13363.10	-6777.06	-11554.00
L_4	4816.16	19348.00	30910.60	15229.97	N/A	N/A

[28] See Pagano (1989).

expected future rate, volatility will tend to increase. One possible explanation is that trades are executed at the reservation price and or investors have some notion of a threshold rate, with any deviation from such a rate resulting in enhanced price variability due to uncertainty.

4.2 Multivariate GARCH

Tables 6 and 7 show the results from estimating a trivariate

TABLE 6. PARAMETER ESTIMATES FOR MULTIVARIATE GARCH (FRACTIONAL)

Parameters	*Estimates*	*Standard Errors*	*t-Statistic*
Mean			
constant 1	0.00	0.00	31.51
constant 2	-0.03	0.01	-2.94
constant 3	-0.09	0.19	-0.48
Variance			
constant 1	0.47	0.09	5.50
constant 2	0.01	0.03	0.27
constant 3	0.35	0.04	9.33
constant (1,2)	0.00	0.00	-0.11
constant (1,3)	0.00	0.00	0.42
constant (3,2)	0.01	0.00	7.16
error 1	0.78	0.09	8.38
error 2	1.07	0.09	11.78
error 3	-0.42	0.14	-2.96
error (1,2)	0.00	0.00	3.55
error (1,3)	0.00	0.00	2.65
error (3,2)	-0.01	0.00	-2.94
variance 1	0.00	0.00	3.64
variance 2	0.13	0.01	16.28
variance 3	2.64	0.19	13.87
variance (1,2)	0.00	0.00	-4.69
variance (1,3)	0.00	0.00	0.79
variance (3,2)	-0.03	0.02	-2.11
No. of Observation		222.00	
Function value		1373.9	
variable 1 = dlwsxr			
variable 2 = dlcross rate			
variable 3 = dir			

NOTE: dlwsxr, dlcross rate and dir represent change in the exchange rate, change in the cross rate and the change in the interest rate, respectively.

TABLE 7. PARAMETER ESTIMATES FOR MULTIVARIATE GARCH (1ST DIFFERENCE)

Parameters	*Estimates*	*Standard Errors*	*t-Statistic*
Mean			
constant 1	0.00	0.00	0.82
constant 2	-0.01	0.01	-0.60
constant 3	0.08	0.23	0.34
Variance			
constant 1	0.00	0.00	0.09
constant 2	0.19	0.00	80.29
constant 3	4.74	0.21	22.78
constant (1,2)	0.00	0.00	-0.38
constant (1,3)	0.00	0.00	-0.44
constant (3,2)	-0.13	0.01	-9.97
error 1	0.31	0.06	4.85
error 2	-0.06	0.06	-1.08
error 3	0.04	0.04	0.83
error (1,2)	0.00	0.00	0.02
error (1,3)	0.00	0.00	1.33
error (3,2)	0.01	0.00	5.61
variance 1	1.00	0.41	2.47
variance 2	0.10	0.02	5.73
variance 3	-0.04	0.15	-0.26
variance (1,2)	0.00	0.00	0.71
variance (1,3)	0.00	0.00	0.06
variance (3,2)	-0.01	0.01	-1.28
No. of Observation		222.00	
Function value		1256.7	
variable 1 = dlwsxr			
variable 2 = dlcross rate			
variable 3 = dir			

NOTE: dlwsxr, dlcross rate and dir represent change in the exchange rate, change in the cross rate and the change in the interest rate, respectively.

GARCH model using the daily exchange rate, USD/Pound exchange rate (cross rate) and the private money market interest rate over the period 7 February 2002 to 12 February 2003, a total of 250 trading days. The parameters from the multivariate GARCH model, which includes the fractionally difference exchange rate series, shown in Table 6, are mostly statistically significant.[29] The

[29] The results for the 1st differences series reported in Table 7 are largely insignificant.

mean equations, with the exception of the interest rates equation, suggest that a simple random walk model is sufficient to capture the dynamics of the variables under consideration.

With regard to the variance equations, all three markets are affected by shocks and volatility originating within the respective markets. There is also strong evidence supporting the existence of contagion. For the foreign exchange market, the positive and significant error and variance suggests that the higher the level of past foreign exchange rate volatility and shocks, the higher the level of current volatility. The significance of past shocks validates the long memory and volatility-clustering feature of foreign exchange return. In addition, this result speaks to inefficiencies, in particular, the rate of information arrival and the extent of market concentration. In a very concentrated market it may not be possible to diversify or minimise exposure to past shocks.

Shocks originating in the money and cross rate markets increase volatility in the foreign exchange market. Past volatility in the money market has no effect on volatility in the foreign exchange market. However, volatility in the cross rate market tend to provide a stabilizing effect on both the money and foreign exchange markets.

5. FORECASTS

In evaluating the forecasting power of the various GARCH models, the measure of volatility is important. As in Chong et al. (1999) the following measure of the volatility for the foreign exchange market is used

$$\sigma_t^2 = (r_t - \bar{r})^2$$

where σ_t^2 is the volatility, r_t, is the actual daily return for day t, and $\bar{r}$ is expected return for day t. The expected return over 30 days is measured by calculating the arithmetic average of daily returns from day 1 to day 29. The expected return on day 31 is calculated by taking the arithmetic average daily returns from day 2 to day 30. This is repeated for the entire forecast period. Squaring the difference between the actual and moving average returns generates the implied volatility indicated by the above equation.

In order to obtain the one-period ahead forecast error for the different GARCH models, the following equation is employed

$$\mu_{t+1} = \sigma_{t+1}^2 - \hat{h}_{t+1}$$

where $\hat{h}_{t+1}$ is the forecasted variance generated by the GARCH models.

We report the traditional mean square error statistics. We also run the following three forecast evaluation regressions:

$$\text{(i)}\ \sigma_{t+s} = a_0 + a_1 \hat{\sigma}_{t+s}^{t-GARCH}$$

$$\text{(ii)}\ \sigma_{t+s} = b_0 + b_1 \hat{\sigma}_{t+s}^{alternative\ model}$$

$$\text{(iii)}\ \sigma_{t+s} = c_0 + c_1 \hat{\sigma}_{t+s}^{t-GARCH} + c_3 \hat{\sigma}_{t+s}^{alternative\ model}$$

By comparing the R^2s from these models, we assess forecasting performance of the models relative to the linear t-GARCH (1,1) model.

The mean square error statistics for the models are reported in Table 8. The non-linear (asymmetric) forecasting models outperform the linear models with the fractionally differenced series. (Generally, there is no significant difference in the forecasting performance of the various models when a unit root is assumed in the data.)

TABLE 8. RESULTS FOR 30-DAY FORECAST (FRACTIONAL)

	TGARCH	*EGARCH*	*GJR*	*TS-GARCH*	*TARCH*	*APARCH*
RMSE	0.526	0.019	0.003	0.604	0.008	0.003
MAE	0.528	0.019	0.003	0.614	0.008	0.003
THEIL	0.950	0.680	0.512	1.000	0.487	0.470

Tables 9 to 10 report the results of the forecast encompassing test.[30] The second column on the right side of the table records the results of the test that the benchmark, t-GARCH model, en-

TABLE 9. FORECAST ENCOMPASSING TESTS FULL GARCH MODELS (FRACTIONAL)

	Forecast Horizon(s)	Ra^2	Rb^2	Rc^2
TGARCH vs. EGARCH	30	0.026	0.070	0.199
TGARCH vs. GJR-GARCH	30	0.026	0.001	0.030
TGARCH vs. TS-GARCH	30	0.026	0.030	0.050
TGARCH vs. APARCH	30	0.026	0.000	0.110
TGARCH vs. TARCH	30	0.026	0.005	0.058

[30] Results for the simple GARCH models are provide in the Appendix.

TABLE 10. FORECAST ENCOMPASSING TESTS FULL MODELS (1ST DIFFERENCE)

	Forecast Horizon(s)	Ra^2	Rb^2	Rc^2
TGARCH vs. EGARCH	30	0.000	0.017	0.017
TGARCH vs. GJR-GARCH	30	0.000	0.002	0.003
TGARCH vs. TS-GARCH	30	0.000	0.942	0.942
TGARCH vs. APARCH	30	0.000	0.013	0.013
TGARCH vs. TARCH	30	0.000	0.030	0.034

NOTES: The first column records standard errors for the TGARCH model which represent the null hypothesis that the TGARCH model encompasses the alternative. The second column test whether the alternative model encompasses the TGARCH model. The third column reports the coefficient of determination for the TGARCH model, the fourth column for the alternative model and the final column the combination of the two models.

compasses all the information in the alternate models. The third column reports the result for the test that the alternatives contain all the information in the benchmark model. The final column records the result of the tests of whether the combination of both models would lead to an improvement of the forecast generated by either models. The results from the fractional model indicate that any combination of the benchmark model and the alternative models result in an improvement of the forecasts. The TS-GARCH model performs best over the forecast horizon.

6. CONCLUSION

The empirical distribution of the exchange rate reflects or validates the tendency of agents to hold relatively long positions in foreign exchange. The key indicators of future market conditions are Jamaica Dollar liquidity conditions, spread and the volume of trade.

While the theoretical relationship between volume and volatility (i.e. the mixture distribution hypothesis) does not hold in Jamaica, the expected rate of depreciation one day ahead, influences current volatility. An increase in supply to the market does lower the rate of movement in the exchange rate. Expected liquidity conditions one day ahead were found to be important for the mean level of returns but not for volatility. Further, current liquidity conditions, proxied by the change in the monetary base, do not seem to be important for market volatility. There is evidence of a DOW effect. This suggests that caution has to be exercised when reacting to movements in the exchange rate as the market may simply be reflecting information generated during periods when the market is closed.

There is evidence of a long memory process, which means that shocks to the exchange rate persists for a long period. Further, consistent with expectation, there are spill over effects from the money market and international currency markets. This supports the monetary authority's emphasis on ensuring stable conditions in the financial markets and indicates that policy has to respond quickly to shocks to the foreign exchange market. This is further reinforced by the results that the market response to shocks is asymmetric, reflecting the volatility feedback hypothesis. In contrast to the results for more developed markets, it was found that generally, models, which account for non-linearities in the Jamaica Dollar exchange rate provide a better out-of sample forecasts.

Appendix A

FIGURE 1. AUTOCORRELATION AND PARTIAL AUTOCORRELATION OF LOG EXCHANGE

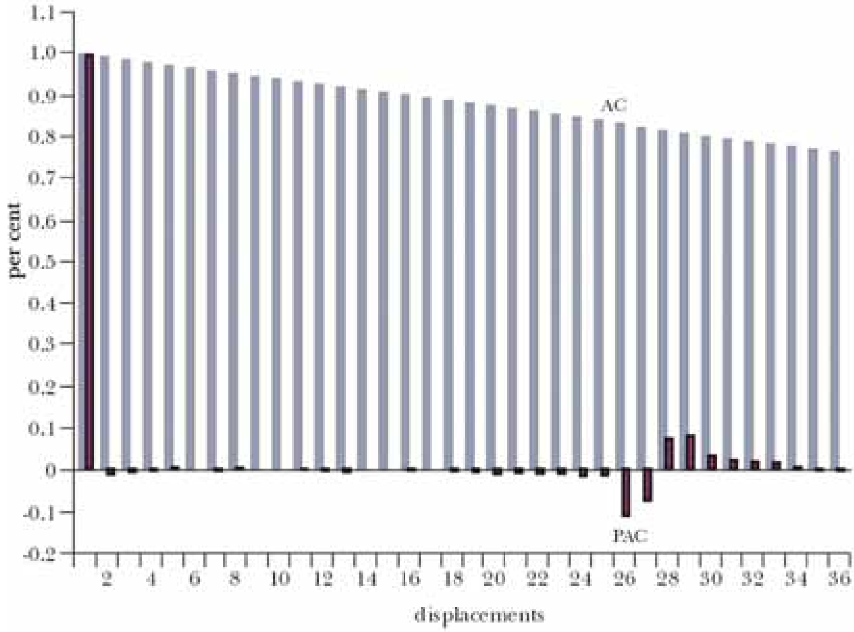


FIGURE 2. DEGREE OF INTEGRATION

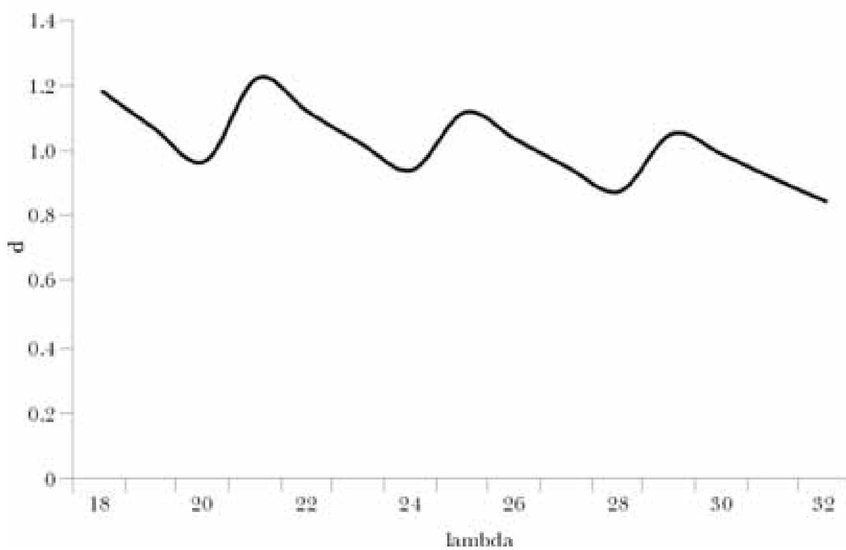


FIGURE 3a. DAILY EXCHANGE RATE RETURN (J$/US$)

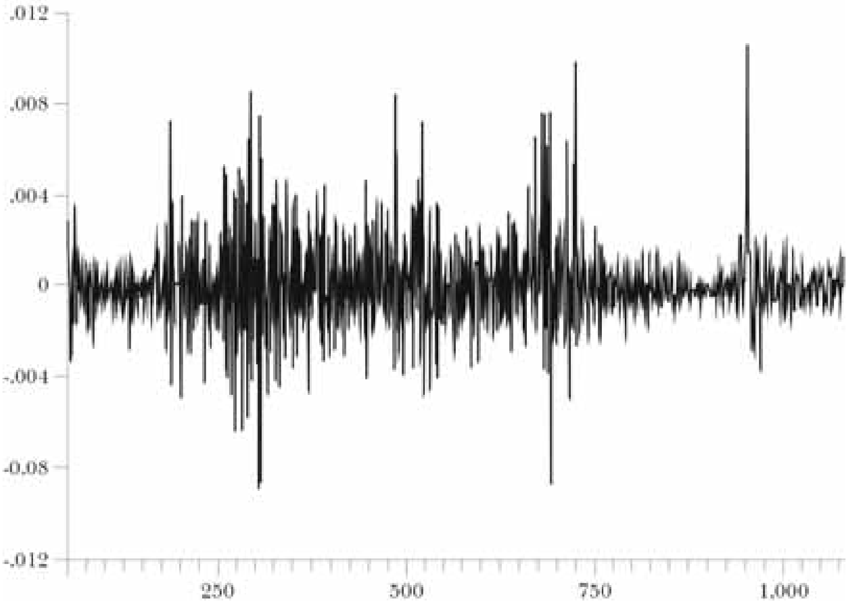


FIGURE 3b. DAILY FRACTIONALLY DIFFERENCE EXCHANGE RATE RETURN

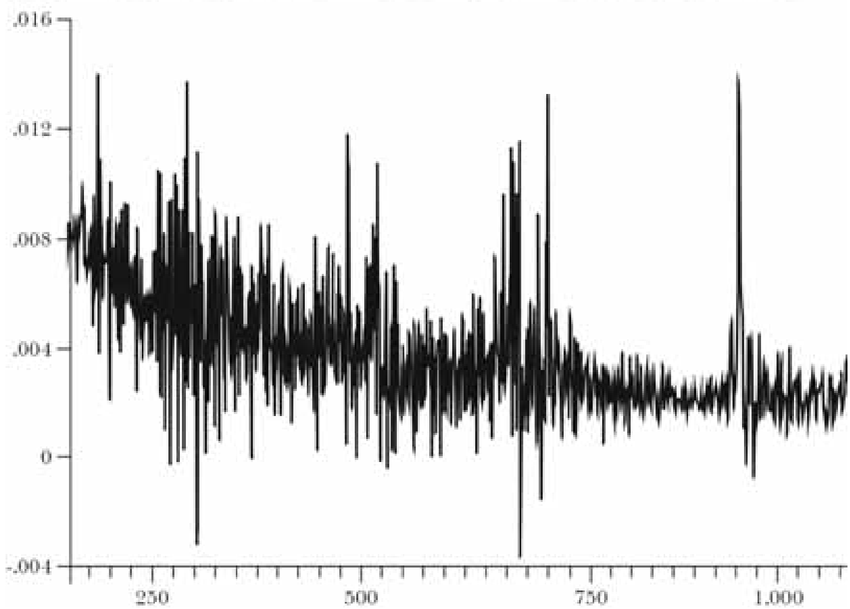


FIGURE 4a. DISTRIBUTION OF FRACTIONALLY DIFFERENCE RETURN SERIES

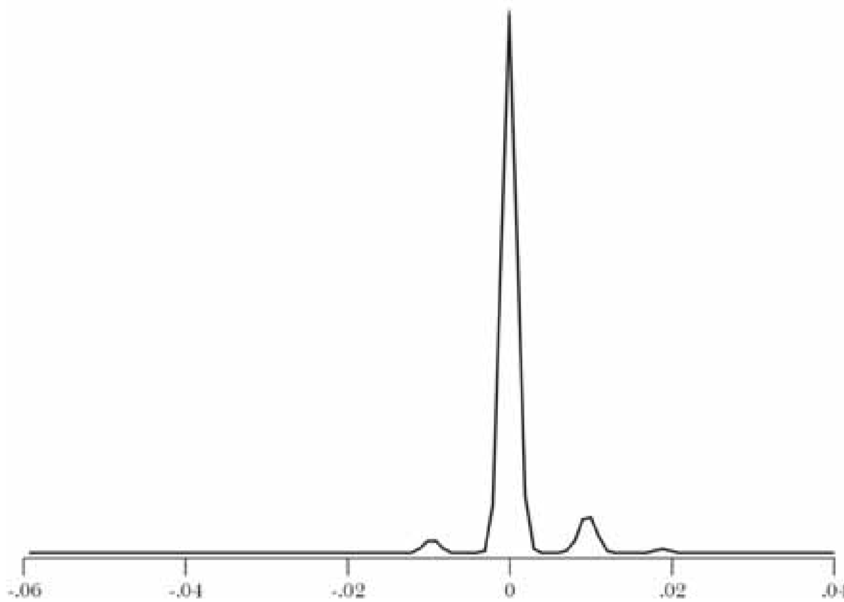
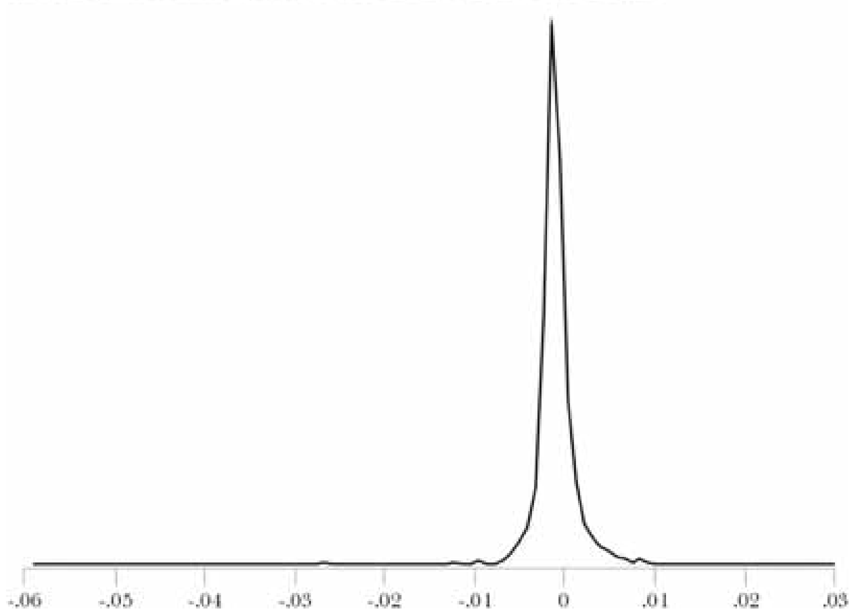


FIGURE 4b. DISTRIBUTION OF DIFFERENCE RETURN SERIES



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Víctor Olivo

Interest rate rules *vs.* money growth rules: some theoretical issues and an empirical application for Venezuela

1. INTRODUCTION

There is currently little discussion about monetary policy based on monetary aggregates. In his paper *Recent Developments in the Analysis of Monetary Policy* (1999), Bennett McCallum states:

> "The nearly standard framework at the NBER and Riksbank conferences is a quantitative macroeconomic model that includes three main components. These are:
>
> – An IS-type relation (or set of relations) that specifies how interest rate movements affect aggregate demand and output;
> – A price adjustment equation (or set of equations) that specifies how inflation behaves in response to the output gap and expectations regarding future inflation; and

Paper prepared by Víctor Olivo, Economist of the Vice-presidency of Research, Central Bank of Venezuela and presented at the IX Meeting of the Central Bank Researchers Network, held in San José, Costa Rica. The views expressed in this paper are those of the author and should not be interpreted as those of the Central Bank of Venezuela.

– A monetary policy rule that specifies each period's settings of an interest-rate instrument.

These settings typically are made in response to recent or predicted values of the economy's inflation rate and its output gap."

Later in that paper McCallum asserts:

> "So what is actually being assumed implicitly, by analyses that exclude m_t (i.e. $m_t - p_t$) from the relation 1, is that the effects of money holdings on spending are quantitatively small (indeed negligible). This is a belief with a long tradition, and I am inclined to think that it is probably justifiable, but the whole matter needs additional study."

Another argument frequently exposed to favor an interest rate rule over a monetary aggregate rule is that the latter is more prone to monetary shocks, particularly due to unexpected fluctuations in money demand.

With respect to McCallum's assessment about the relevance of money, Meltzer (2001) and Nelson (2002) present interesting theoretical and empirical (for the US and British economy) arguments that supports the importance of the real monetary base for aggregate spending decisions.

Regarding the susceptibility of money rules to money demand shocks, Walsh (2003, p. 488) points out:

> "Changes in the short-term interest rate that serves as the operational target for implementing monetary policy will affect aggregate spending decisions only if longer – term rates of interest are affected. While the use of an interest – rate – oriented policy reduces the importance of money demand in the transmission of policy actions to the real economy, it raises to prominence the role played by the term structure of interest rates."

This is a very relevant issue, because the relationship between the short – term interest rate, used as operational target by the monetary authority, and longer - term interest rates may be affected by financial innovations and other shocks, just as the demand for money. Moreover, in developing countries this link may be further weakened by the presence of shallow financial markets, unstable fiscal policy, and central banks with poor track records in providing monetary stability.

In addition, Neumann and von Hagen (2002), and Ball and Sheridan (2003) have presented interesting evidence that shows that the disinflation process observed in many countries during the 90s occurred under different monetary policy arrangements, not only inflation targeting or other interest-rate oriented monetary policy strategies.

This paper main theme is that the arguments against the use of money (i.e. money growth rate rules) in the conduct of monetary policy are not so strong, particularly for less developed economies. This topic is analyzed in two ways: i) using some simple theoretical forward-looking macro models and evaluating their inflation and output variance under interest rate and monetary aggregates rules; ii) setting up models similar to the theoretical ones, but with more complex dynamics, assigning values to the parameters, and solving them for different kind of shocks under interest rate and monetary aggregates rules.

Before proceeding with the detailed analysis, it is important to clarify certain basic assumptions from the outset:

- The models developed are not derived from the solution of the dynamic optimization problem in representative agent models. Their structure, however, is very similar to the linearized versions obtained from these models. In particular, the inclusion of forward-looking variables is intended to capture some of the main features of the models based on micro-foundations.
- The models are basically of short/medium – run nature, so they do not include capital accumulation relations. This follows McCallum and Nelson (1999) that contend that for monetary analyses in this time horizon, fluctuations in the stock of capital do not play a major role.
- Aggregate demand shocks may have a fiscal origin, but we assume the absence of fiscal dominance.
- The analysis does not try to determine optimal policy rules, instead it sets up ad-hoc simple rules that serve as benchmarks for monetary policy and may facilitate transparency and communication. The rules are, however, specified to guarantee some basic theoretical requirements: for example, in the interest rate rule the Taylor principle is maintained to ensure the existence of a determinate monetary equilibrium under rational expectations.
- It is assumed implicitly that the central bank's concern with social welfare is represented by its aim to minimize a loss function similar to the one employed in the Barro-Gordon rules *vs.* discretion discussion, with inflation and output deviations from some target values as arguments: $L = (\pi_t - \pi_t^*)^2 + \lambda \tilde{y}_t^2$. This is a

reasonable theoretical assumption, but again of an ad-hoc nature.[1]

– No attempt is made to model explicitly the open economy sector. This is an exercise particularly difficult for the Venezuelan economy, which has experienced during the 90s, frequent modification of its exchange rate regime. Instead, we rely on the fact that some versions of the open economy models built on the foundations of optimizing agents and sticky prices, could be reduced to a form that is isomorphic to the typical closed economy new Keynesian model (see Walsh 2003, Chapter 11). Hence, external sector shocks are analyzed as either aggregate demand or aggregate supply shocks

– In the monetary policy literature is common to distinguish between *Targeting Regimes* (i.e. Inflation Targeting), and *Instrument Rules* (i.e. the Taylor rule). Ball (1997) and Olivo (2003), however, show that there is a close relationship between the two schemes, thus we refer to them interchangeably along the analysis.

The paper is organized as follows: after this introduction, in section 2, the theoretical models are set up and solved to derive their inflation and output gap variances; section 3 builds upon these models to analyze empirically with data for Venezuela, the possible effects of different shocks under an interest-rate rule and two types of money-growth rules. Finally, some conclusions from the analysis are presented.

2. BASIC THEORETICAL MODELS

In this section, three simple AD-AS (Philips-curve) forward-looking models are set up: one with an interest rate rule; one with a Friedman constant money growth rate; and the last with a flexible or feedback money growth rule. The models are solved for the inflation rate and the output gap using the method of undetermined coefficients, and the variances of inflation and the output gap derived assuming that the various shocks are uncorrelated.

[1] Woodford (2003), Chapter 6, derives a loss function similar to the ad-hoc one from the expected utility of the representative agent.

2.1 Model with an interest rate rule

The interest rate rule model has a structure very similar to the log-linear approximation of the basic model developed by Woodford (2003, Chapter 4). The interest rate rule model includes a forward – looking aggregate demand equation in which the output gap $(\tilde{y}_t)$ responds to the long-run real interest rate. Thus in equation (2), i is defined as a long-run nominal interest rate that according to the expectations theory of the interest rate can be expressed as:

$$i_t = (1/n)\sum_{i=0}^{n-1} E_t i^s_{t+i} \ , \qquad (1)$$

where i^s is the short-run interest rate.

Given this, the interest rate equation (3) has two components: a) the interest rate rule that the central bank follows which determines the short-term interest rate; b) a random error υ_t that reflects the imperfect control of the monetary authority over the long-run interest rate. The introduction of a random shock in the interest rate equation attempts to capture the imperfect relationship between the short – term interest rate that is adjusted by a rule and the long – term interest rate relevant in the aggregate demand equation. This shock may be as important as the shock considered in the model with a rule for a monetary aggregate that is fundamentally linked to money demand shocks. Note that the interest rate equation specification further differs from the original Taylor rule in that it includes $E_t\pi_{t+1}$ instead of π_t in the inflation gap term.

The model also includes an aggregate demand shock (ε_t), and an aggregate supply shock (η_t). In the interest rate equation, r stands for the long-run equilibrium real interest rate and π^* for the target inflation rate.

The third equation in the model (4) is a forward – looking Phillips curve based on staggered price adjustment of the type suggested by Calvo (1983).

The complete model is specified as follows:

$$\tilde{y}_t = E_t \tilde{y}_{t+1} - \beta(i_t - E_t\pi_{t+1}) + \varepsilon_t \text{; AD equation,}^2 \qquad (2)$$

$$i_t = g_0 \tilde{y}_t + g_1(E_t\pi_{t+1} - \pi^*) + \bar{r} + \pi^* + \upsilon_t \text{; Interest rate equation;} \qquad (3)$$

[2] Note that defining $r_t \equiv i_t - E_t\pi_{t+1}$, the forward solution of this equation is: $\tilde{y}_t = -\beta\sum_{i=0}^{\infty} r_{t+i} + \sum_{i=0}^{\infty} \varepsilon_{t+i}$.

$$\pi_t = E_t\pi_{t+1} + \gamma \tilde{y}_t + \eta_t \text{; AS equation.} \quad (4)$$

Solving the model for the output gap and the rate of inflation in terms of the forward – looking variables, exogenous variables, and the random shocks yields:

$$\tilde{y}_t = \frac{1}{(1+\beta g_0)} E_t \tilde{y}_{t+1} + \frac{\beta(1-g_1)}{(1+\beta g_0)} E_t\pi_{t+1} - \frac{\beta(1-g_1)}{(1+\beta g_0)}\pi^* - \frac{\beta}{(1+\beta g_0)}\bar{r} - \frac{\beta}{(1+\beta g_0)}\nu_t + \frac{1}{(1+\beta g_0)}\varepsilon_t \quad ; \quad (5)$$

$$\pi_t = \frac{\gamma}{(1+\beta g_0)} E_t \tilde{y}_{t+1} +_t [1 + \frac{\gamma\beta(1-g_1)}{(1+\beta g_0)}] E_t\pi_{t+1} - \frac{\gamma\beta(1-g_1)}{(1+\beta g_0)}\pi^* - \frac{\gamma\beta}{(1+\beta g_0)}\bar{r} - \frac{\gamma\beta}{(1+\beta g_0)}\nu_t + \frac{\gamma}{(1+\beta g_0)}\varepsilon_t + \eta_t \quad . \quad (6)$$

The following trial solutions to apply the method of undetermined coefficients are used:

$$\tilde{y}_t = \delta_0 + \delta_1\varepsilon_t + \delta_2\eta_t + \delta_3\upsilon_t \, ; \quad (7)$$

$$\pi_t = \lambda_0 + \lambda_1\varepsilon_t + \lambda_2\eta_t + \lambda_3\upsilon_t \, . \quad (8)$$

Next, the solutions for the output gap and inflation, and their respective variances assuming that the shocks are uncorrelated are obtained :

Solution for $\tilde{y}_t$

$$\tilde{y}_t = \frac{1}{(1+\beta g_0)}\varepsilon_t - \frac{\beta}{(1+\beta g_0)}\upsilon_t \, . \quad (9)$$

Variance of $\tilde{y}_t$

$$\sigma^2_{\tilde{y}} = \frac{1}{(1+\beta g_0)^2}\sigma^2_\varepsilon + \frac{\beta^2}{(1+\beta g_0)^2}\sigma^2_\upsilon \, . \quad (10)$$

Solution for π_t

$$\pi_t = \pi^* + \frac{1}{1-g_1} r + \frac{\gamma}{(1+\beta g_0)}\varepsilon_t + \eta_t - \frac{\gamma\beta}{(1+\beta g_0)}\upsilon_t \, . \quad (11)$$

Variance of π_t

$$\sigma^2_\pi = \frac{\gamma^2}{(1+\beta g_0)^2}\sigma^2_\varepsilon + \sigma^2_\eta + \frac{(\gamma\beta)^2}{(1+\beta g_0)^2}\sigma^2_\upsilon \, . \quad (12)$$

2.2 Models with money growth rules

In this section, two models with money growth rules are specified. In these models the forward – looking aggregate demand function responds to the growth rate of real money defined as the monetary base. Nelson (2002) discusses the empirical and theoretical evidence that supports the inclusion of the real monetary base in the aggregate demand equation. Based in arguments developed by Meltzer (2001) and Friedman - Schwartz (1963), Nelson contends that the inclusion of the long – term nominal interest rate in the money demand function in the optimizing IS – LM model allows to derive an aggregate demand function that incorporates the real monetary base. Meltzer and Friedman and Schwartz argument is based on the monetarist transmission mechanism in which money exerts its influence over a wide array of relative prices and not only on a short-run interest rate. In this sense, money is a good indicator of the long-run interest rate that is more relevant that a short-run interest rate in the aggregate demand equation (equation 13).

The first money growth rule model is based on Friedman proposal of a fixed rate of growth for the money supply. In equation (14), the rate of growth of the monetary base (μ) follows the growth rate of potential output *(k)*, the inflation rate targeted by the monetary authority (π^*), plus a shock that captures both money demand and supply random variations *(z)*. In the second model, the money growth rule is a feedback rule where money growth follows the Friedman rule, but in addition responds to the output gap $(\tilde{y}_t)$ and the inflation gap $(E_t\pi_{t+1} - \pi^*)$, plus a shock that captures money demand and supply random variations *(z)* -equation 25-.

As in the interest rare rule model, the third equation in each model (15 and 26) is a forward – looking Phillips curve based on staggered price adjustment of the type suggested by Calvo (1983).

The detail specification of the money growth rule models and their rational expectations solutions are presented bellow.

2.2.1 Model with a fixed money growth rule

The basic equations of the model are the following:

$$\tilde{y}_t = E_t \tilde{y}_{t+1} + \psi(\mu_t - \pi_t) + \varepsilon_t \text{; AD equation;} \quad (13)$$

$$\mu_t = k + \pi^* + z_t \text{ ; Money growth rule;} \quad (14)$$

$$\pi_t = \gamma \tilde{y}_t + E_t\pi_{t+1} + \eta_t \text{; AS equation;} \quad (15)$$

Solving the model for the output gap and the rate of inflation in terms of the forward – looking variables, exogenous variables, and the random shocks, yields:

$$\begin{aligned}\tilde{y}_t = & \frac{1}{(1+\gamma\psi)}E_t\tilde{y}_{t+1} - \frac{\psi}{(1+\gamma\psi)}E_t\pi_{t+1} + \frac{\psi}{(1+\gamma\psi)}(\pi^* + k) \\ & + \frac{\psi}{(1+\gamma\psi)}z_t - \frac{\psi}{(1+\gamma\psi)}\eta_t + \frac{1}{(1+\gamma\psi)}\varepsilon_t\end{aligned} \text{;} \quad (16)$$

$$\begin{aligned}\pi_t = & \frac{1}{(1+\gamma\psi)}E_t\tilde{y}_{t+1} + \frac{(1-\gamma\psi)}{(1+\gamma\psi)}E_t\pi_{t+1} + \frac{\gamma\psi}{(1+\gamma\psi)}(\pi^* + k) \\ & + \frac{\gamma\psi}{(1+\gamma\psi)}z_t + \frac{\gamma}{(1+\gamma\psi)}\varepsilon_t + \frac{(1-\gamma\psi)}{(1+\gamma\psi)}\eta_t\end{aligned} . \quad (17)$$

Using the trial solutions:

$$\tilde{y}_t = \delta_0 + \delta_1\varepsilon_t + \delta_2\eta_t + \delta_3 z_t; \quad (18)$$

$$\pi_t = \lambda_0 + \lambda_1\varepsilon_t + \lambda_2\eta_t + \lambda_3 z_t; \quad (19)$$

the following results for the output gap and the inflation rate and their respective variances, assuming that the shocks are uncorrelated, emerge:

Solution for $\tilde{y}_t$:

$$\tilde{y}_t = \frac{1}{1+\gamma\psi}\varepsilon_t - \frac{\psi}{1+\gamma\psi}\eta_t + \frac{\psi}{1+\gamma\psi}z_t. \quad (20)$$

Variance of $\tilde{y}_t$:

$$\sigma^2_{\tilde{y}} = (\frac{1}{1+\gamma\psi})^2\sigma^2_\varepsilon + (\frac{\psi}{1+\gamma\psi})^2\sigma^2_\eta + (\frac{\psi}{1+\gamma\psi})^2\sigma^2_z \ . \quad (21)$$

Solution for π_t:

$$\pi_t = \pi^* + k + \frac{\gamma}{1+\gamma\psi}\varepsilon_t + \frac{(1-\gamma\psi)}{1+\gamma\psi}\eta_t + \frac{\gamma\psi}{1+\gamma\psi}z_t. \quad (22)$$

Variance of π_t:

$$\sigma^2_\pi = (\frac{\gamma}{1+\gamma\psi})^2\sigma^2_\varepsilon + (\frac{1-\gamma\psi}{1+\gamma\psi})^2\sigma^2_\eta + (\frac{\gamma\psi}{1+\gamma\psi})^2\sigma^2_z \ . \quad (23)$$

2.2.2 Model with a flexible money growth rule

The model is specified as follows:

$$\tilde{y}_t = E_t \tilde{y}_{t+1} + \psi(\mu_t - \pi\ ) + \varepsilon_t \text{ ; AD equation ;} \quad (24)$$

$$\mu_t = k + \pi^* - h_0 \tilde{y}_t - h_1(E_t\pi_{t+1} - \pi^*) + z_t \text{ ; Money growth rule;} \quad (25)$$

$$\pi_t = \gamma \tilde{y}_t + E_t\pi_{t+1} + \eta_t \text{ ; AS equation;} \quad (26)$$

Solving the model for the output gap and the rate of inflation in terms of the forward – looking variables, exogenous variables, and the random shocks, yields:

$$\tilde{y}_t = \frac{1}{1+\psi(h_0+\gamma)} E_t \tilde{y}_{t+1} - \frac{\psi(1+h_1)}{1+\psi(h_0+\gamma)} E_t\pi_{t+1} + \frac{\psi}{1+\psi(h_0+\gamma)} k + \frac{\psi(1+h_1)}{1+\psi(h_0+\gamma)} \pi^* + \frac{\psi}{1+\psi(h_0+\gamma)} z_t - \frac{\psi}{1+\psi(h_0+\gamma)} \eta_t + \frac{1}{1+\psi(h_0+\gamma)} \varepsilon_t \quad ;(27)$$

$$\pi_t = \frac{\gamma}{1+\psi(h_0+\gamma)} E_t \tilde{y}_{t+1} + [1 - \frac{\gamma\psi(1+h_1)}{1+\psi(h_0+\gamma)}] E_t\pi_{t+1} + \frac{\gamma\psi}{1+\psi(h_0+\gamma)} k + \frac{\gamma\psi(1+h_1)}{1+\psi(h_0+\gamma)} \pi^* + \frac{\gamma\psi}{1+\psi(h_0+\gamma)} z_t + \frac{\gamma}{1+\psi(h_0+\gamma)} \varepsilon_t + \frac{1+\psi h_0}{1+\psi(h_0+\gamma)} \eta_t \quad .(28)$$

Using the trial solutions:

$$\tilde{y}_t = \delta_0 + \delta_1\varepsilon_t + \delta_2\eta_t + \delta_3 z_t ; \quad (29)$$

$$\pi_t = \lambda_0 + \lambda_1\varepsilon_t + \lambda_2\eta_t + \lambda_3 z_t ; \quad (30)$$

the following results for the output gap and the inflation rate and their respective variances (assuming that the shocks are uncorrelated) are obtained,

Solution for $\tilde{y}_t$:

$$\tilde{y}_t = \frac{1}{1+\psi(h_0+\gamma)} \varepsilon_t - \frac{\psi}{1+\psi(h_0+\gamma)} \eta_t + \frac{\psi}{1+\psi(h_0+\gamma)} z_t . \quad (31)$$

Variance of $\tilde{y}_t$:

$$\sigma^2_{\tilde{y}} = (\frac{1}{1+\psi(h_0+\gamma)})^2 \sigma^2_\varepsilon + (\frac{\psi}{1+\psi(h_0+\gamma)})^2 \sigma^2_\eta + (\frac{\psi}{1+\psi(h_0+\gamma)})^2 \sigma^2_z \ . \quad (32)$$

Solution for π_t :

$$\pi_t = \frac{1}{1+h_1}k + \pi^* + \frac{\gamma}{1+\psi(h_0+\gamma)}\varepsilon_t + \frac{1+\psi h_0}{1+\psi(h_0+\gamma)}\eta_t + \frac{\gamma\psi}{1+\psi(h_0+\gamma)}z_t\,. \quad (33)$$

Variance of π_t:

$$\sigma_\pi^2 = [\frac{\gamma}{1+\psi(h_0+\gamma)}]^2\sigma_\varepsilon^2 + [\frac{1+\psi h_0}{1+\psi(h_0+\gamma)}]^2\sigma_\eta^2 + [\frac{\gamma\psi}{1+\psi(h_0+\gamma)}]^2\sigma_z^2\ . \quad (34)$$

2.3 Comparing variances

In this section the output and inflation variances obtained for the three models under analysis are compared.

In terms of the variance of output, the interest rate rule has the advantage over the money growth rules as it isolates output fluctuations from supply shocks (see equations 10, 21, 32). Further, compared with the Friedman rule, the interest rate rule may reduce the volatility of the output gap to aggregate demand and monetary sector shocks through the adjustment of the policy parameter g_0 that captures the response of the interest rate to the output gap. The feedback money growth rule may be also superior to the Friedman rule in terms of output gap volatility, as it also allows to reduce fluctuations by adjusting the policy parameter h_0 that captures the response of the rate of growth of money to the output gap.

In terms of the variance of inflation the money growth rules are superior to the interest rate rule in the presence of aggregate supply shocks. These shocks have a smaller effect (coefficient less than one) on inflation variations in the models with money growth rules, in contrast to a coefficient equal to one in the model with an interest rate rule (see equations 12, 23, 34).

Compared to the Friedman rule the interest rate rule and the feedback money growth rule may reduce inflation volatility due to aggregate demand shocks through the adjustment of the policy parameters g_0, h_0.

But the main issue highlighted in this analysis is that output and inflation volatility will be affected by monetary sector shocks under both kinds of monetary policy rules. The effects of this kind of shocks on the volatility of the output gap and the inflation rate, however, can be reduced under the interest rate rule and the flexible money growth rule through adjustments in the parameters g_0, h_0.

3. EVALUATION OF SHOCKS IN CALIBRATED MODELS FOR VENEZUELA

In this section calibrated versions of the models discussed previously using data from the Venezuelan economy are set up and solved. In contrast to the theoretical models that include only forward – looking and contemporaneous variables, the calibrated models have a more complex dynamic that is captured by adding several backward - looking variables. In this case, trying to get analytical solutions is more complicated hence, calibration is an attractive option. The parameters for the calibration of the aggregate demand and supply equations are obtained from econometric estimates based on quarterly data for the period 1990 – 2002. The policy equations parameters are chosen from values used in the literature in the case of the interest rate rule, and from evaluating successive specifications in the case of the feedback money growth rule. I then solve the models in the Eviews solver using the Gauss-Seidel iterative method.[3] I present the calibrated models below.

i) Interest rate rule model

y = 0.25 * y(-1) + 0.20 * y(+1) - 0.09 * (i(-4) - dp(-4)) + e
i = 0.5 * y + 1.5 * (dp(+1) - dpm) + r + dpm + v
dp = 0.41 * dp(-1) + 0.59 * dp(+1) + 0.085 * y(-4) + n

ii) Fixed money growth rule model

y = .31 * y(-1) + .38 * y(+1) + 0.19 * (dm(-4) - dp(-4)) + e
dm = dpm + k + z
dp = .41 * dp(-1) + .59 * dp(+1) + 0.085 * y(-4) + n

iii) Flexible money growth rule model

y = 0.31 * y(-1) + 0.38 * y(+1) + 0.19 * (dm(-4) - dp(-4)) + e
dm = - 0.40 * y - 1.0 * (dp(+1) - dpm) + k + dpm + z
dp = 0.41 * dp(-1) + 0.59 * dp(+1) + 0.085 * y(-4) + n

Where y is the output gap; i the long-run nominal interest rate; dp the rate of inflation; dpm the target inflation rate; e is

[3] In the Gauss-Seidel algorithm, in each iteration, each equation of the model is solved for the value of its associated endogenous variable, treating all other endogenous variables as fixed. The iterative process is repeated until changes in the values of the endogenous variables between successive iterations become less than a specified tolerance (see Eviews 4 User's Guide).

the aggregate demand shock; r is a long-run equilibrium real interest rate; v is a money market shock that affects the relationship between the short-run and the long-run nominal interest rate; n is a supply or cost shock; dm is the rate of growth of the monetary base; k is the rate of growth of natural output; z is a money market shock that captures random fluctuations in both money demand and money supply.

The aggregate demand equation in the interest rate rule is derived from the following OLS regression using quarterly data from 1990 to 2002:

Dependent Variable: LYG
Method: Least Squares
Date: 11/25/03 Time: 11:56
Sample(adjusted): 1992:2 2002:4
Included observations: 43 after adjusting endpoints

Variable	*Coefficient*	*Std. Error*	*t-Statistic*	*Prob.*
C	0.033668	0.018931	1.778427	0.0831
LYG(-1)	0.248048	0.153977	1.610945	0.1153
LYG(1)	0.202487	0.100435	2.016099	0.0507
I(-4)-DLP(-4)	-0.091821	0.056834	-1.615608	0.1142

R-squared	0.309635	Mean dependent var	0.008362
Adjusted R-squared	0.256530	S.D. dependent var	0.044426
S.E. of regression	0.038306	Akaike info criterion	-3.598001
Sum squared resid	0.057227	Schwarz criterion	-3.434168
Log likelihood	81.35701	F-statistic	5.830617
Durbin-Watson stat	2.459533	Prob(F-statistic)	0.002163

Where LYG is the output gap measure as the difference between the logarithm of GDP and its Hodrick – Prescott trend; I is a short-run nominal lending rate that proxies for the long-run nominal rate due to the lack of information of the latter; DLP is the inflation rate measured as the first difference of the logarithm of the CPI.

Given the presence of the output gap one period ahead [LYG(1)], we tried to estimate the equation using the GMM method. The results were highly sensitive to the instruments used, and in all cases none of the coefficients of the variables were significantly different from zero at standard critical levels.

The aggregate demand equation in the money growth models is derived from the following OLS regression using quarterly data from 1990 to 2003 (first quarter):

Dependent Variable: LYG
Meted: Least Squares
Date: 11/20/03 Time: 17:02
Sample(adjusted): 1990:2 2003:1
Included observations: 52 after adjusting endpoints

Variable	*Coefficient*	*Std. Error*	*t-Statistic*	*Prob.*
C	-0.000697	0.006069	-0.114863	0.9090
LYG(-1)	0.307908	0.137911	2.232656	0.0303
LYG(1)	0.380361	0.107145	3.549952	0.0009
DLBM(-4)-DLP(-4)	0.194744	0.039559	4.922835	0.0000
R-squared	0.472962	Mean dependent var		0.002116
Adjusted R-squared	0.440022	S.D. dependent var		0.057918
S.E. of regresión	0.043341	Akaike info criterion		-3.365621
Sum squared resid	0.090166	Schwarz criterion		-3.215525
Log likelihood	91.50614	F-statistic		14.35835
Durbin-Watson stat	1.989039	Prob(F-statistic)		0.000001

Where DLBM is the rate of growth of the monetary base measured as the first difference of the logarithm of the monetary base.

The GMM estimation was again very sensitive to the instruments used, particularly for the coefficients of LGY(-1) and LYG(1). The results indicate, however, that the coefficients of LYG(1) and (DLBM(-4)-DLP(-4)) are significantly different from zero (p values close to zero), with the latter taking values around 0.28.

The aggregate supply equation, that is shared by all the models, is taken from a paper by Arreaza et al. (2003) who estimated it using the GMM method. A similar estimation is derived , however, by using OLS with the White heteroskedacity correction:

Dependent Variable: DLP
Meted: Least Squares
Date: 02/17/04 Time: 15:45
Sample(adjusted): 1991:1 2002:3
Included observations: 47 after adjusting endpoints
White Heteroskedasticity-Consistent Standard Errors & Covariance

Variable	*Coefficient*	*Std. Error*	*t-Statistic*	*Prob.*
C	-0.003112	0.006767	-0.459854	0.6479
DLP(-1)	0.515437	0.093172	5.532084	0.0000
DLP(1)	0.522560	0.069900	7.475813	0.0000
LYG(-4)	0.098432	0.059642	1.650391	0.1061
R-squared	0.854392	Mean dependent var		0.081710
Adjusted R-squared	0.844233	S.D. dependent var		0.047191
S.E. of regression	0.018625	Akaike info criterion		-5.047370
Sum squared resid	0.014916	Schwarz criterion		-4.889910
Log likelihood	122.6132	F-statistic		84.10431
Durbin-Watson stat	2.583677	Prob(F-statistic)		0.000000

The models are solved for the different shocks: aggregate demand, aggregate supply, and monetary sector shocks. In each case, a one-unit shock that last for eight quarters is introduced and the impulse – response functions for the output gap and the inflation rate are studied.

3.1 Interest rate rule *vs.* Friedman money growth rule

Figure 1 shows the effect of an aggregate demand shock of one unit that last eight quarters on the output gap. The continuous line (y–i) represents the evolution of the output gap under the interest rate rule and the dotted line (y-m) under the Friedman rule. There is evidently a sharper initial reaction and a more volatile response of the output gap to an aggregate demand shock under the Friedman rule. The variance of the output gap with the Friedman rule is 0.29 *vs.* 0.11 with the interest rate rule.

The effect of an aggregate demand shock on the inflation rate is shown in Figure 2. The continuous line (dp-i) traces the effect on inflation under the interest rate rule and the dotted line (dp-m) under the Friedman rule. In this case, the interest rate rule and the Friedman rule produce similar results, both in terms of the initial reaction of the inflation rate and its volatility. The variance of the interest rate rule is 0.23 against 0.25 for the Friedman rule.

The effect of an aggregate supply shock of one unit for eight quarters on the output gap is illustrated in Figure 3. The Friedman rule (y-m) generates much more volatility with a variance of 0.53 than the interest rule (y-i) with a variance of 0.017.

Figure 4 shows that in contrast to what is observed in the case of the output gap, the Friedman rule (dp-m) produces a smaller initial response and reduces the overall volatility of the inflation rate when supply shocks occur, relative to the interest rate rule (dp-i). The variance of inflation under the Friedman rule is 2.14 compared to 3.36 under the interest rate rule.

The next two figures (5 and 6) show the effects of random shocks in the money market. For the interest rate model, a one unit reduction in the interest rate (y-i and dp-i) that lasts for eight quarters is introduced ; for the Friedman model a one unit increase in money growth (y-m and dp-m) for the same period is introduced. Here the interest rate rule gives rise to less volatility in both cases, particularly for the output gap. In this latter case the variance is 0.0007 for the interest rate rule compare to 0.07 for the Friedman rule; for the inflation rate, the variance is 0.003 for the interest rate rule against 0.01 for the Friedman rule.

FIGURE 1. AGGREGATE DEMAND SHOCK

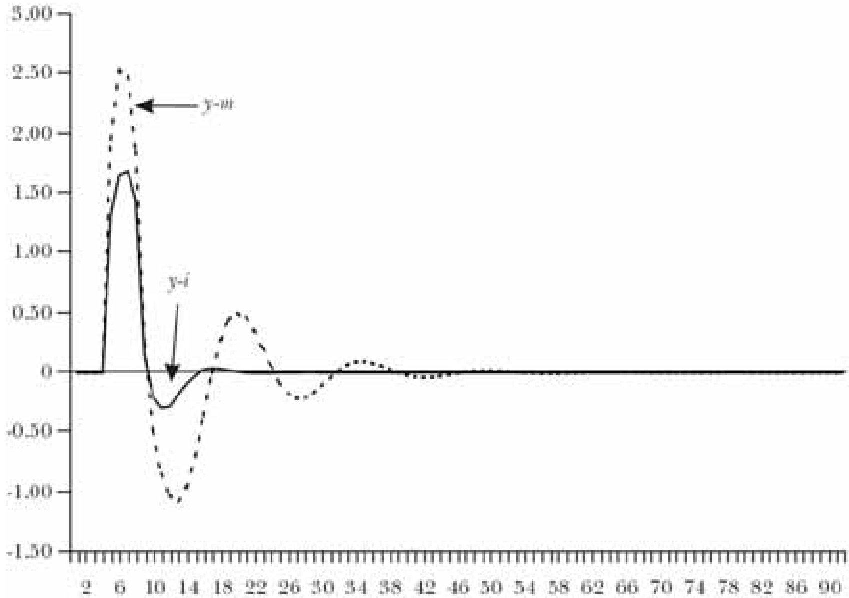


FIGURE 2. AGGREGATE DEMAND SHOCK

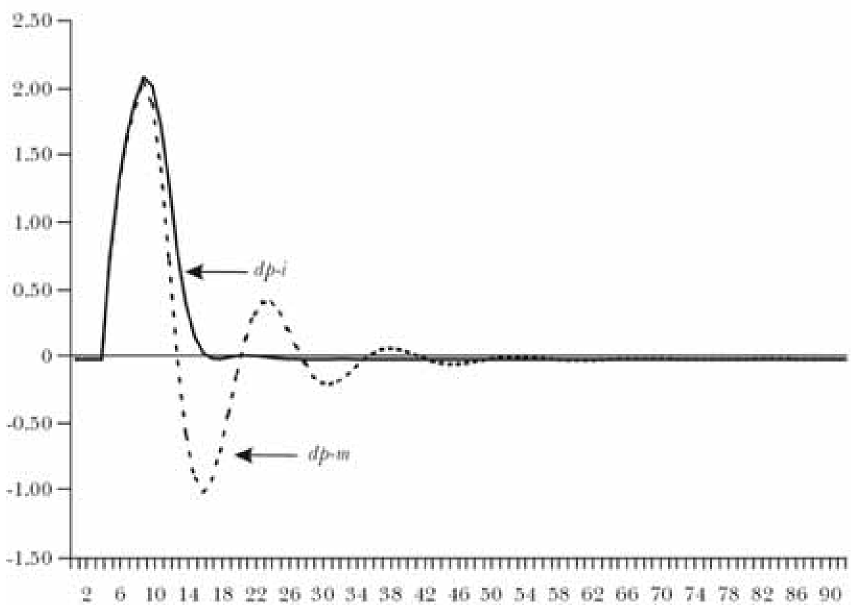


FIGURE 3. AGGREGATE SUPPLY SHOCK

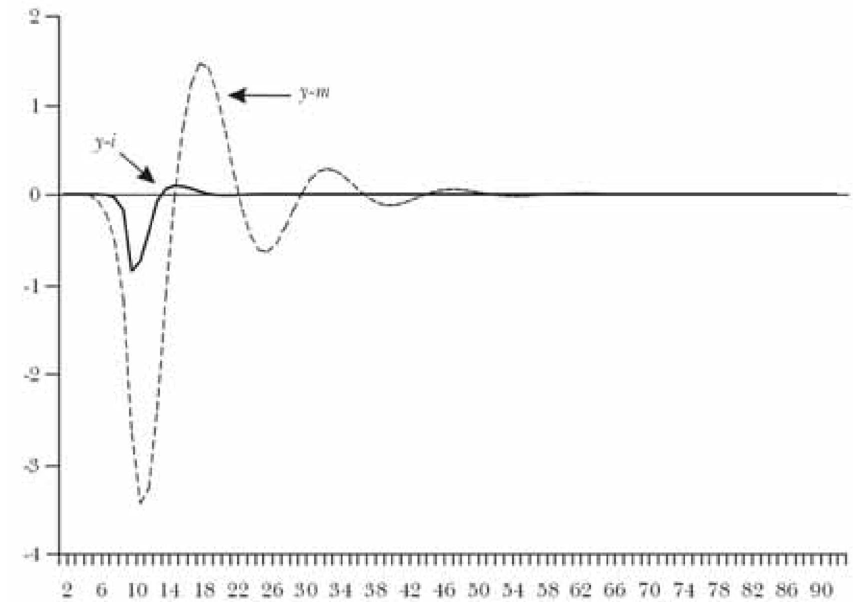


FIGURE 4. AGGREGATE SUPPLY SHOCK

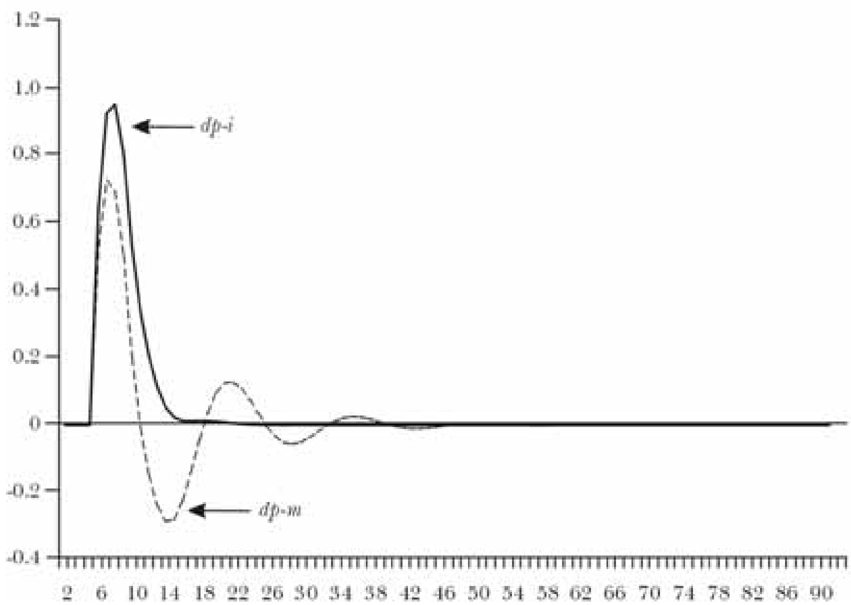


FIGURE 5. MONETARY INSTRUMENT SHOCK

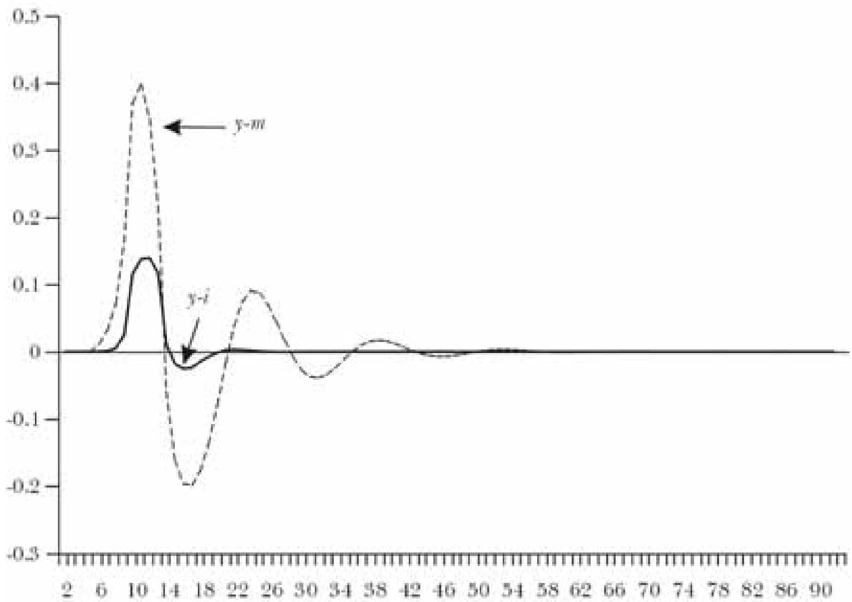


FIGURE 6. MONETARY INSTRUMENT SHOCK

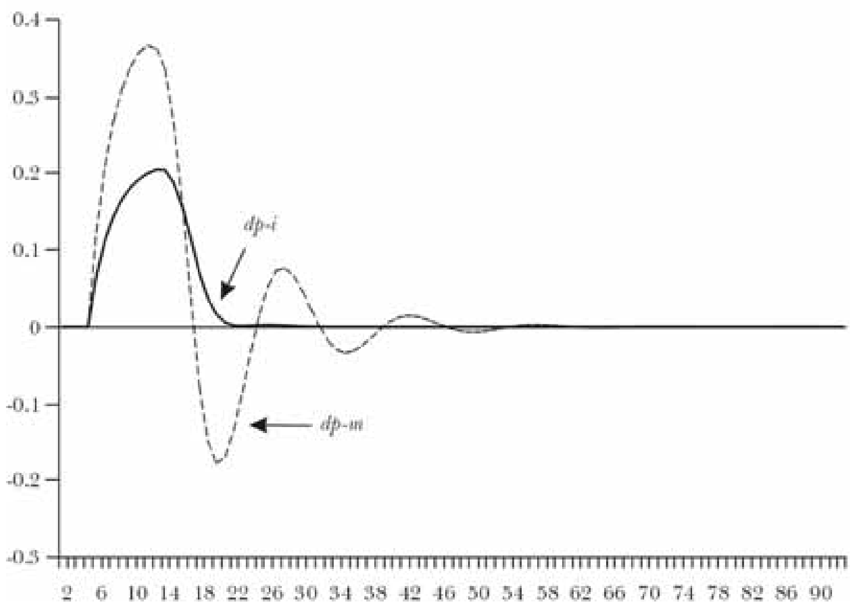


FIGURE 7. AGGREGATE DEMAND SHOCK

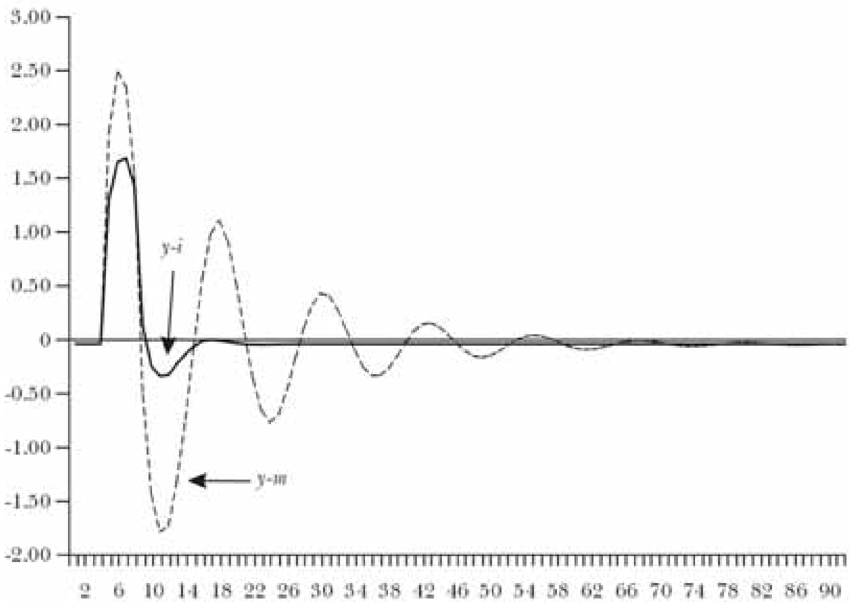
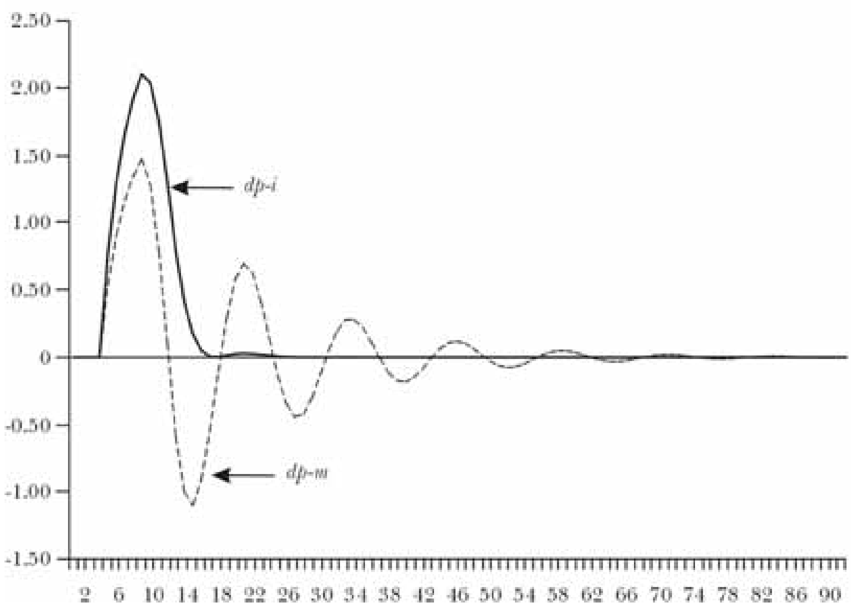


FIGURE 8. AGGREGATE DEMAND SHOCK



In general, the interest rate rule tends to generate a smaller initial reaction and less overall volatility in the output gap under any kind of shock. It also causes less volatility of the inflation rate when a money market shock occur. The Friedman rule outperforms the interest rate rule only in terms of fluctuations in inflation due to an aggregate supply shock. The reaction of inflation to an aggregate demand shock is similar under the two rules.

3.2 Interest rate rule *vs.* Flexible money growth rule

In this section, the introduction of some flexibility in the Friedman rule is evaluated against the interest rate rule. After several trials a money growth rule that adds to the Friedman rule the output gap with a policy parameter of –0.40 and the inflation gap with a policy parameter equal to –1 was chosen.

Figure 7 presents the effect of an aggregate demand shock on the output gap under each type of rule. The output gap exhibits less fluctuations under the interest rate rule (y-i) than under the flexible money growth rule (y-m). The variance of the output gap under the interest rate rule is 0.11 against 0.35 with the flexible rule. Notice that the flexible rule introduces even more volatility on the output gap than the Friedman rule.

In Figure 8, we observe that under a flexible money growth rule (dp-m) an aggregate demand shock produces less volatility of the inflation rate than under the interest rate rule (dp-i). The variance of inflation under the flexible rule is 0.16 *vs.* 0.23 under the interest rate rule. Also in this case, the flexible money growth rule outperforms the Friedman rule (variance of 0.25).

The effect of an aggregate supply shock on the output gap under each rule is displayed in Figure 9. It shows that the interest rate rule (y-i) generates less fluctuations of the output gap than the flexible money growth rule (y-m). The variance of the output gap with the interest rate rule is 0.017 compare to 1.77 with the flexible money growth rule. Note that the flexible money growth rule introduces more volatility of the output gap than the Friedman rule (variance 0.53).

Figure 10 shows that the flexible money growth rule (dp-m) performs better in terms of inflation volatility than the interest rate rule (dp-i) in the face of a supply shock. The variance of the inflation rate with the flexible money growth rule is 2.15 *vs.* 3.36 with the interest rate rule. The flexible rule performance in this case is quite similar to the Friedman rule (variance 2.14).

In the case of a shock in the money market, the interest rate rule

FIGURE 9. AGGREGATE SUPPLY SHOCK

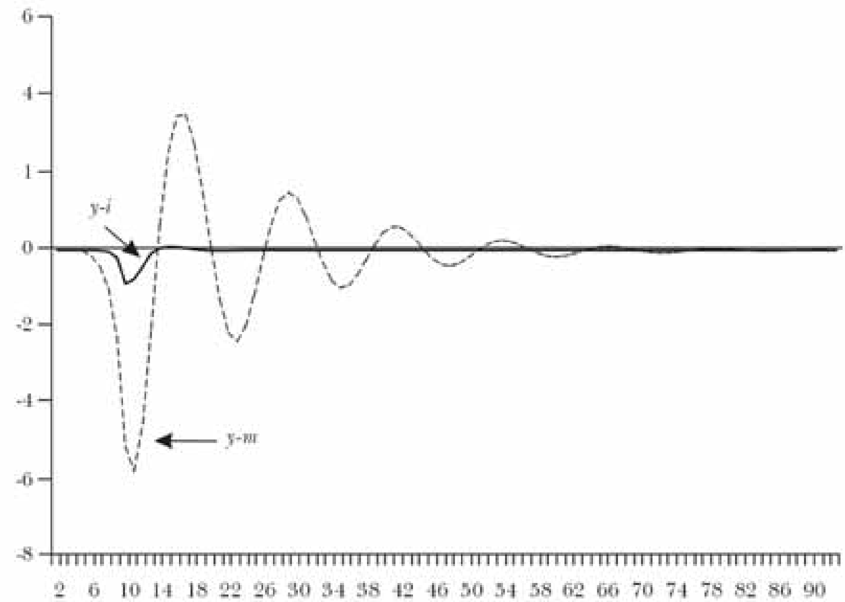
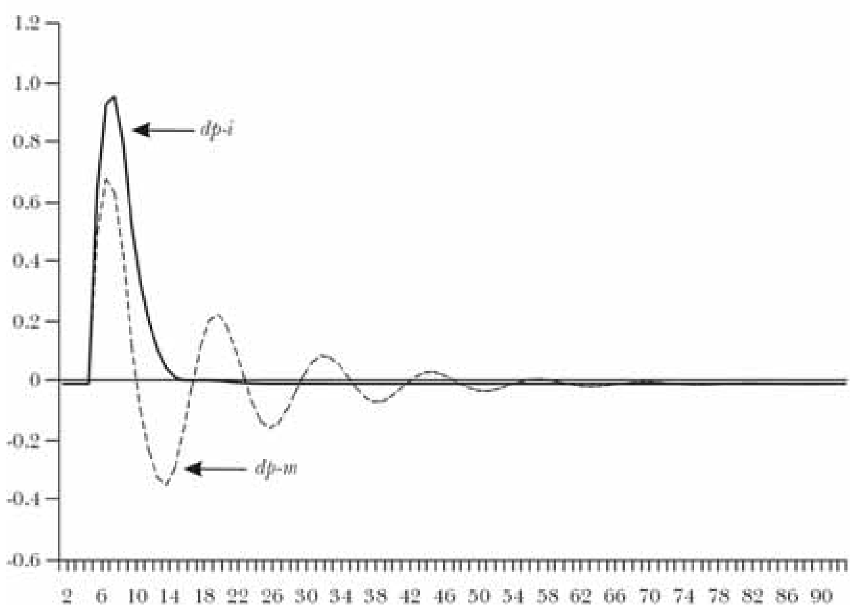


FIGURE 10. AGGREGATE SUPPLY SHOCK



(y-i) performs better than flexible money growth rule (y-m) in terms of the output gap (Figure 11). The variance of the output gap with the interest rate rule is 0.0007 compare to 0.008 with the flexible money growth rule. The effects of a shock in the money market on the inflation rate are very similar (Figure 12). The variance of the inflation rate under the interest rate rule (dp-i) is 0.003 *vs.* 0.005 under the flexible money growth rule (dp-m). In terms of the inflation rate, the flexible money rule outperforms the Friedman rule (variance 0.01 for the inflation rate).

In general, the flexible or feedback money growth rule specified outperforms the interest rate rule and the Friedman constant money growth rule in terms of the reaction of inflation to the different shocks, but introduces more volatility on the output gap. Thus, interestingly, the introduction of some flexibility in the money growth rate in the form of parameters adjusting to the output gap and the inflation gap improves its performance in terms of inflation fluctuations, but not in terms of the output gap behavior. These results held for a variety of combinations of the output gap and inflation gap parameters.

The disadvantage of the money growth rules in terms of output gap volatility is related to the larger effect of the real growth of the monetary base on the output gap relative to the real interest rate effect. In addition, the coefficients of the one-lagged period and the one period ahead output gap are larger in the aggregate demand equation specified with the real monetary base than in the one that includes the real interest rate.

The following table summarizes the results discussed above about the effects of the different kinds of shocks on the output gap and the inflation rate.

SUMMARY TABLE: OUTPUT GAP AND INFLATION VARIANCE UNDER DIFFERENT KINDS OF SHOCKS

	Interest rate rule	*Friedman rule*	*Flexible money rule*
AD shocks			
Output gap	0.11	0.29	0.35
Inflation	0.23	0.25	0.16
AS shocks			
Output gap	0.017	0.53	1.77
Inflation	3.36	2.14	2.15
Money market shocks			
Output gap	0.0007	0.07	0.008
Inflation	0.003	0.01	0.005

FIGURE 11. MONETARY INSTRUMENT SHOCK

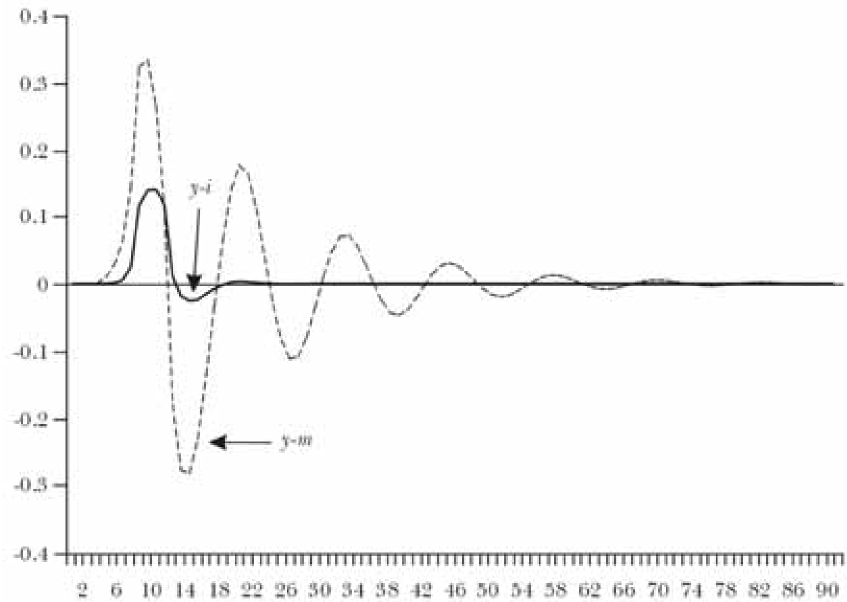
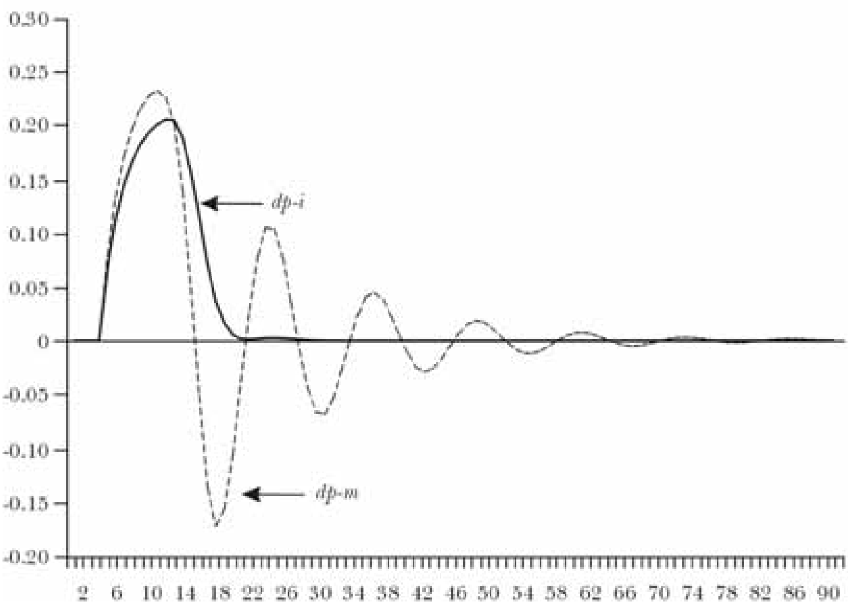


FIGURE 12. MONETARY INSTRUMENT SHOCK



4. CONCLUSIONS

This paper attempts to compare the performance of an interest rate rule, similar to the widely popular Taylor rule, with that of the money growth rule proposed by Friedman (1962, A Program for Monetary Stability), and a flexible version of this rule that responds to the output gap and the inflation gap (given a target inflation rate set by the monetary authority). Performance in this paper is measured in terms of the volatility (unconditional variance) exhibit by the output gap and the inflation rate when different shocks affect the economy under each type of rule.

The issue is first explored by specifying a theoretical stochastic aggregate demand – aggregate supply model of a forward-looking nature. To this basic model the monetary policy rules under analysis are added. The three resulting models are solved with a rational expectations approach through the method of undetermined coefficients. The variance of the output gap and the inflation rate solutions obtained are compared, assuming that the random errors included in each equation of the model are uncorrelated. The second approach to the problem is based on setting up models with dynamics a little more complex than that of the theoretical ones, and calibrated with information of the Venezuelan economy.

The main conclusion derived from both analyses is that the economy (characterized by the output gap and the rate of inflation) should behave in a fairly similar way when different shocks are introduced under the diverse rules considered. From a theoretical point of view, this conclusion is warranted by the fact that the introduction of different monetary policy rules should not change radically the basic structure of the economy. Hence, the variances for the output gap and the inflation rate under the different rules share a similar specification. In the empirical analysis, this conclusion derives from the fact that the impulse-response functions present similar patterns for each rule under the different random shocks. In general, the interest rate rule is better in terms of stabilization of the output gap, while the money growth rules, particularly the flexible version, is better in terms of inflation stabilization.

These results accord with those in the empirical studies conducted by Neumann and von Hagen (2002) and Ball and Sheridan (2003), which find that different monetary strategies achieved similar disinflationary results during the 90s.

What is then the reason for the substantial improvement in the effectiveness of monetary policy and central banks in the 90s in terms of macroeconomic stabilization? Here we contend that the answer is the widespread change in focus toward price stability in the long-run. Although in the short-run there is still room for output stabilization in terms of reducing its volatility, this objective has been kept in check by the pursuance of the overriding goal of price stability in the long-run. In contrast to previous decades, central banks in the 90s seriously committed to the attainment of price stability. Although, this emphasis of monetary policy on price stability only gained popularity in the 90s, it has been intensively promoted by Milton Friedman since the early 60s.

Both, the academia and the central banks, however, had put up some resistance to the idea that central banks should abandon monetary policy excessively biased towards output stabilization. Since Milton Friedman stated his first seminal ideas about this issue, several researchers (i.e. Robert Lucas, Robert Barro, David Gordon) have been introducing new elements that gradually reinforced the change in focus we see today.

Some countries have accompanied the focus on price stability with profound institutional changes that comprise the granting of legal instrument independence of their central banks to pursue this goal, and the introduction of mechanisms of transparency to make them accountable to society for its achievement. In other countries, notably the United States, the institutional arrangement has not been altered, but the central banks have adapted their practices to the new paradigm.[4]

Despite the crucial role assigned to the commitment to price stability in the long-run to the recent success of monetary policy in terms of macroeconomic stabilization, this paper shares Friedman's concern with the granting of too much discretion to monetary authorities. "The granting of wide and important responsibilities that are neither limited by clearly defined rules for guiding policy nor subject to test by external criteria of performance is a serious defect of our present monetary arrangements. It renders monetary policy a potential source of uncertainty and instability" (Friedman, 1960, p. 86).

A modern restatement of this argument is presented by Woodford (2003), who considers that to lock in this success, it is necessary to accompany it with a policy commitment. "A systematic ap-

[4] Mishkin (2000) considers that the FED should "advocate a change in its mandate to put price stability as the overriding, long-run goal of monetary policy."

proach to policy provides an explicit framework for decision making within the bank, but that is also used to explain the bank´s decisions to the public." (Woodford 2003, p. 14). This proposal is based on Woodford's view that when the private sector behavior is forward - looking as implied by optimizing models, central banking is basically about shaping market expectations.

But this study departs from Woodford in his assessment that interest rate rules are the first option to establish a policy commitment. This paper holds that the widespread rejection of monetary policy rules based on the management of a monetary aggregate is not well supported neither theoretically nor empirically. The evidence provided by Meltzer (2001) and Nelson (2002) makes a strong case in favor of a monetary policy based on monetary aggregates in advanced economies. Meltzer (2001) goes beyond econometrics and present some interesting historic data from periods of deflation "to show than changes in real interest rates cannot explain some major episodes in monetary history."

The case in favor of a monetary policy based on the control of a monetary aggregate is stronger in less advanced countries with their shallow financial markets and weak fiscal institutions. In this environment, the link between the short-run interest rate managed by the central bank and the long-run rate relevant for aggregate demand decisions may be quite weak and unstable.

A weaker conclusion from this study would be that even if monetary policy is conducted based on an interest rate, monetary aggregates should still play a major role in policy decisions. This view is strongly supported by Poole (1994) and Meltzer (2001), and is embedded in the European Central Bank "two pillars" strategy.

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The bases of the 2006 call for papers are as follows:

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