

# **A signal of imperfect portfolio capital adjustments from the relationship between yields of domestic and foreign Colombian debt**

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## **Summary**

*In this paper we check the relationship between the yields of the Colombian bonds traded in the (secondary) internal market and the yields of the sovereign global securities for the sample period 1999-2001. The hypothesis we maintain is that, under the assumption of capital mobility, it should exist a comovement between the two yields that we effectively find. However, the results suggest that capital mobility is much less than perfect. By invoking concepts of immunization and duration we find evidence of a TAR adjustment cointegration between the two yields plus a constant risk premium for bonds with maturity in 2003 and a symmetric adjustment cointegration between the yields of securities with maturity in 2004. Since the assets are issued by the same issuer (the Colombian Government) the country or credit risks is the same for the bonds we consider that the risk premium is purely connected to currency risks: exchange-rate and inflation risks.*

*JEL classification: C32, G14, G15.*

*Key words: yield, interest parity, capital mobility, nonlinearities, cointegration, threshold adjustment.*

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## 1. Introduction

In this work, following an extremely simple approach, we check the hypothesis of portfolio capital mobility for Colombia. The approach consists of comparing the yields of domestic debt instruments with those of external Colombian debt instruments. We test for long-run equilibrium between yields of sovereign debt securities, with the same year of maturity, both expressed in a common currency for the sample period 1999-2001.

Two strands of literature could eventually be related to the topic considered in this paper: long term capital mobility and the exchange rate-interest rate differential. In the first case, the tests have focused, on the one hand, on the relationship between saving and investment. Regardless of the multiple tests to the hypothesis of (long-term) capital mobility the evidence is not conclusive (for a survey see Obstfeld, 1995). Sometimes the estimations suggest low long-term capital mobility<sup>1</sup> (e.g. Feldstein and Horioka, 1980; Dooley et al., 1987; Tesar, 1991; and Bayoumi, 1990, among many others), whereas other results suggest high mobility (e.g. Taylor, 1994; and Ghosh, 1995). On the other hand, the tests have analyzed the relationship between domestic investment and the current account to determine the extent to which the capital account has been financed by the foreign saving (Sachs, 1983). Regarding short and possibly medium term capital mobility, instead of the saving-investment relationships as in the previous cases, the exchange rate-interest rate differential has been the building block of the tests. Literature is also vast on this issue (see for example, Campbell and Clarida, 1987; Meese and Rogoff, 1988; Baxter, 1994). In particular, Campbell and Clarida and Meese and Rogoff, by using different econometric methodologies, rejected the hypothesis that there is a statistically significant link between real exchange rates and real interest rate differentials. Then, the assumptions of sticky-price theories of exchange rate determination (Dornbusch, 1976; Frankel, 1979) underlying the models become inadequate empirically. Baxter (1994), on the other hand, by using a filtering-correlation approach, found that the higher correlations between exchange rates and real interest rate differentials are found at trend and business cycle-frequencies while there was no relationship between them at high frequencies (cycle of 2-5 quarters). Given these results, Baxter concluded that the link between the real exchange rates and real interest rate differentials is very

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<sup>1</sup> This finding might be evidence of incomplete, segmented and imperfect financial markets. These are incomplete when some securities are not available for trading. The markets are segmented when some investors are not allowed to trade some securities. Finally, financial markets are imperfect when frictions such as transaction costs cause the equilibrium to differ from the perfect market outcome (This concepts are borrowed from Dumas, 1994).

weak since most of the movement in the temporary component of real exchange rates is explained by the unobserved risk premium.

As for the case of Colombia, previous attempts for checking the relationship between returns of investment at home and abroad have been done (see, for example, Edwards and Khan, 1985; Correa, 1992; Herrera, 1993; Steiner et al., 1993; and Posada, 1998). With the exception of Posada, who used different series for the domestic and foreign interest rates described in the statistical annex of his paper, the rates used in these studies to revise the interest rate differential are Libor (as a proxy for the external rate) and the Fixed Term Deposit Rates (as a proxy for the domestic rate) aggregating country and currency risks (see also the works in Cárdenas and Garay, 1993).

However, the financing vehicles of governments (and firms) have changed: nowadays they also go to the capital markets rather than only for commercial loans to the private banks. Hence, we consider that it is also important to look at the returns of the securities used to obtain revenues by the government both domestically and from abroad. By picking up securities issued by the same issuer (the Colombian government) we assume that the default (credit) risk is the same for the two securities. In this case, assuming perfect capital mobility and that no currency risks (exchange-rate and inflation risks) are present, then uncovered interest parity condition holds. It would imply that securities are perfect substitutes. As a result, an investor should be indifferent in investing his money in bonds denominated in Colombian pesos or dollars issued by the Colombian government. However, when investors are risk averse and there are currency risks a premium is required to equalize the yields.

The hypothesis we maintain is that under capital mobility it should exist a comovement between the yields of internal and external sovereign debt securities. Given that we do not have any priors on the risk attitudes of investors interested in such securities and the presence and magnitude of currency risks, we check a general version of the interest rate differentials with a constant risk premium for each year of maturity of debt instruments. Further, given the coexistence of different type of traders that might give raise returns being influenced by changing perception of country risks by agents, we allow for nonlinearities (threshold adjustments) in the long run equilibrium relationship between yields.

Since we deal with securities of different term to maturities and coupon rates, to make feasible the comparison of the securities' yields, we invoke the concept of portfolio

*immunization*<sup>2</sup>. Securities with the same Macaulay duration and the same maturity are comparable for our purposes. To find any possible relationship under this approach, where we allow for nonlinearities, we apply the cointegration and threshold adjustment procedure put forth by Enders and Siklos (2001). If there is evidence of capital mobility (a long run comovement between the yields) a crowding out effect might exist independently of the place where public debt is allocated as far as a country risk spread is a component of the yields both internal and external.

From the approach we use, however, we cannot distinguish movements caused by short-term securities arbitrage opportunities from those caused by long-term portfolio movements. This is because any short term movement caused firstly by arbitrage opportunities can evolve towards a long term portfolio decision<sup>3</sup>.

This paper evolves as follows. Section one is this introduction. Section two is devoted to explain the theoretical object we are looking at. Section three explains the concepts of Macaulay duration and portfolio immunization and describes the data. Section four presents the cointegration and threshold adjustment technique. Section five shows and discusses the results. Finally, section six makes some concluding remarks.

## 2. The theoretical object

Uncovered interest parity condition is a building block of exchange rate theories. To hold it needs not only perfect capital mobility but also the same risk and taxation treatment for the securities under comparison. It takes the familiar form:

$$y_{k,t}^{e*} + E_t \Delta e_{t+1} = y_{k,t}^i \quad (1)$$

where  $y_{k,t}^{e*}$  is the yield to maturity, expressed annually, of the securities marketed abroad in foreign currency,  $e_t$  is the exchange rate at the end of period  $t$ ,  $y_{k,t}^i$  is the yield to maturity annual compounded of securities marketed locally expressed in domestic currency,  $E_t$  is the expectations operator,  $\Delta$  is the difference operator, and  $k$  is an index that identifies the year of maturity of the asset we are dealing with. Specifically  $k$  will take the values 03 and 04 for maturities in 2003 and 2004. We shall identify securities marketed in foreign markets as “global

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<sup>2</sup> The data were kindly provided by Catalina Villa from the Ministry of Finance.

securities” whereas those marketed in the local market as “TES” both being representative of sovereign debt. If uncovered interest parity condition holds an investor should be indifferent in investing in assets denominated in Colombian pesos or dollars as long as taxes on returns are the same.

If investors no longer consider global and TES as perfect substitutes then the expected returns on the two assets no longer have to be equal. As a result, uncovered interest parity will not hold. This might be the case when, for example, investors are risk averse and consider TES as more risky as compared to global not because of country risk (they are issued by the same issuer: the Colombian government) but because of currency risks. In this case, they will require a higher expected yield on TES than global securities in a magnitude equal to the risk premium<sup>4</sup>. This being the case, expression (1) may be replaced by:

$$y_{k,t}^{e*} + E_t \Delta e_{t+1} = a + y_{k,t}^i \quad (2)$$

where  $a$  represents the risk premium for each year of maturity. This is the version that we shall check for in what follows. However, as we mentioned above, the coexistence of different type of traders might give raise to the presence of nonlinearities, that we also check for, identified with returns being influenced by changing perception of currency and country risks by agents.

Obviously, since we deal with securities of different term to maturity and coupon rates, the comparison between yields cannot be done directly. To make feasible the comparison of the yields, we invoke the concept of portfolio *immunization* assuming that securities with the same Macaulay duration and the same year of maturity are comparable for our purposes. We do all this treatment next.

### 3. Dealing with data

The price of a (fixed income) coupon bond  $j$ ,  $p^j$  can be calculated as (see details in Fabozzi, 2000):

$$p^j = \sum_{t=1}^n \frac{c^j}{(1+y^j)^t} + \frac{fv^j}{(1+y^j)^n} \quad (3)$$

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<sup>3</sup> We do not have priors on the preferred habitats of investors.

where  $c^j$  stands for the coupon,  $y^j$  for the yield to maturity,  $fv^j$  for the face value,  $n$  for the maturity of the asset and  $j$  for either TES or global bonds. The approximate change of the price for a small change in yield can be calculated as:

$$\frac{\partial p^j}{\partial y^j} = -\frac{1}{(1+y^j)} \left[ \frac{c}{(1+y^j)} + \frac{2c}{(1+y^j)^2} + \dots + \frac{nc}{(1+y^j)^n} + \frac{nfv}{(1+y^j)^n} \right] \quad (4)$$

where the term in brackets is a weighted average term to maturity of the coupons and the bond face value, with the present value of the cash flow used as weights. Dividing both sides of (4) by  $p^j$  gives the approximate percentage price change, known as modified duration:

$$\frac{\partial p^j}{\partial y^j} \frac{1}{p^j} = -\frac{1}{(1+y^j)} \frac{1}{p^j} \left[ \frac{c}{(1+y^j)} + \frac{2c}{(1+y^j)^2} + \dots + \frac{nc}{(1+y^j)^n} + \frac{nfv}{(1+y^j)^n} \right] \quad (5)$$

where the term in brackets divided by the price is referred to as Macaulay duration:

$$\frac{1}{p^j} \left[ \frac{c}{(1+y^j)} + \frac{2c}{(1+y^j)^2} + \dots + \frac{nc}{(1+y^j)^n} + \frac{nfv}{(1+y^j)^n} \right] \quad (6)$$

These two concepts of duration are related to each other in this way:

$$-\text{Modified duration} = \frac{\partial p^j}{\partial y^j} \frac{1}{p^j} = -\frac{1}{(1+y^j)} \times \text{Macaulay duration}$$

In this work, we computed the Macaulay duration for the yields series of TES and global securities with the same year of maturity that takes into account settlement date, coupon rate and frequency of payments.

To build the series we invoke the concept of immunization of portfolios. This is a technique used to protect portfolios against changes of rates when dealing with coupon bonds. Immunization ensures that the design of a portfolio is such that any capital losses due to movements in interest rates are compensated by gains on reinvested return. A coupon bond with a given duration is similar mathematically to a zero-coupon having maturity equal to that duration (Farrell, 1997)<sup>5</sup>. Thus, a portfolio composed fixed-income coupon securities with the same duration is said to be immunized against movements in interest rates.

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<sup>4</sup> Risk aversion on the part of economic agents, perceived differences in risks between TES and global, and a difference between risk minimizing portfolio and the actual portfolio are the three conditions for a risk premium to exist (see Isard, 1983).

<sup>5</sup> We do not make any consideration on convexity of the securities.

The new series are built by matching yields of bonds with the same year of maturity and (almost) with the same duration<sup>6</sup>. These new series are sorted by duration in ascending order. All things being equal, the theoretical relationship between yields and durations should be negative (the higher the yield the smaller the duration). However, in Figure 1 we observe that this relationship is positive. The reason is that in this case, not only the yields are changing but also the coupon rates and the time between the operation date and the date of maturity. After this treatment the resulting series allows us to work with securities of maturity in 2003 and 2004 where the sample sizes are 108 and 64 observations, respectively (Figure 2).

#### 4. Testing for cointegration

By observing expression (2) we could expect that for each year of maturity the returns of securities with the same duration tend to be equal once we allow for a risk term. Consistently, the analysis is focused on equation (2) where we check for any long run equilibrium relationship between the yield of global securities in domestic currency<sup>7</sup>,  $y_{k,t}^e$ , and the yield of the internal securities (TES),  $y_{k,t}^i$ , with maturities in 2003 and 2004, respectively<sup>8,9</sup>. To this end we follow the procedure put forth by Enders and Siklos (2001) closely related to the two-step methodology of Engle and Granger (1987), which consists of running the regressions in equations (7) and (8):

$$y_t^e = \beta_0 + \beta_1 y_t^i + \varepsilon_t^e \quad (7)$$

where  $\varepsilon_t^e$  is disturbance term that may be serially correlated. Notice that  $\beta_0$  represents the risk premium derived from currency risks. The second step of the Engle-Granger approach of testing for cointegration involves running the regression:

<sup>6</sup> The width of the interval of class we choose is 0.01 starting from 0.900 and ending at 5.673. When more than one yield was into the interval we compute the average.

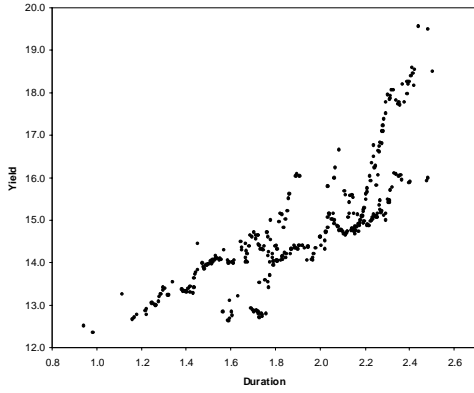
<sup>7</sup> We use the previous depreciation rate of domestic currency as a proxy for the expected rate of depreciation. The value corresponds to a centered moving average of order three. Accordingly, the rate of return in Colombian pesos was calculated as  $y^e = [(1 + y^{e*}) \times (1 + \text{depreciation rate}) - 1] \times 100$ .

<sup>8</sup> According to the results of unit root tests the yields series are nonstationary. These results are available from the authors upon request.

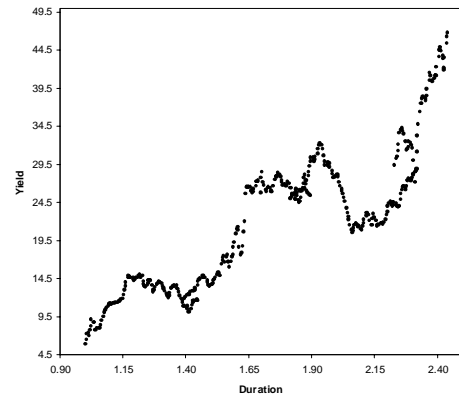
<sup>9</sup> For the rest of this section we drop index  $k$ .

**Figure 1. Yield of TES and Global securities sorted by duration**

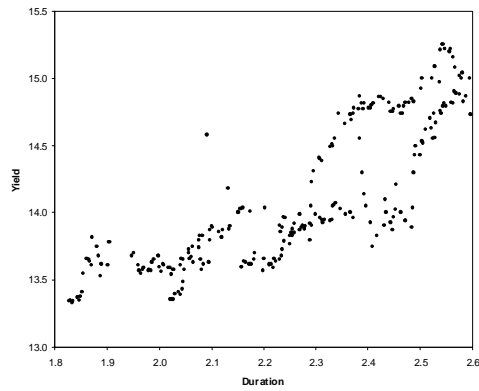
**TES maturity 2003**



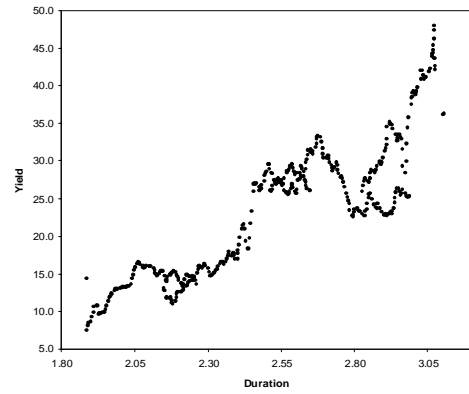
**Global maturity 2003**



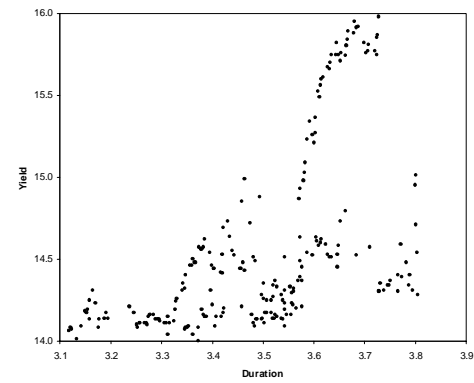
**TES maturity 2004**



**Global maturity 2004**



**TES maturity 2006**



**Global maturity 2006**

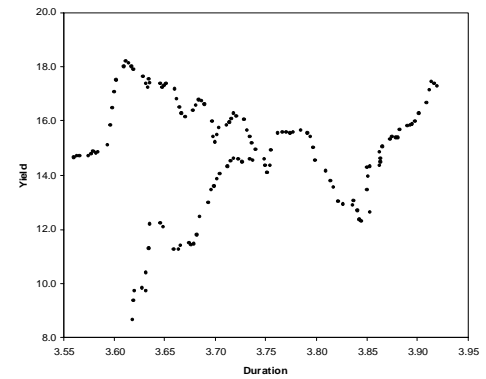




Figure 1 (continued). Yield of TES and Global securities sorted by duration

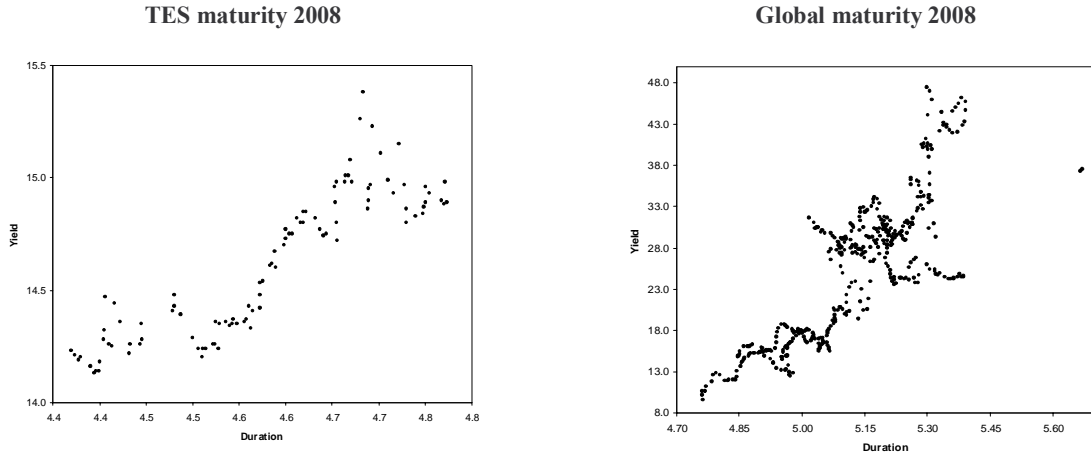
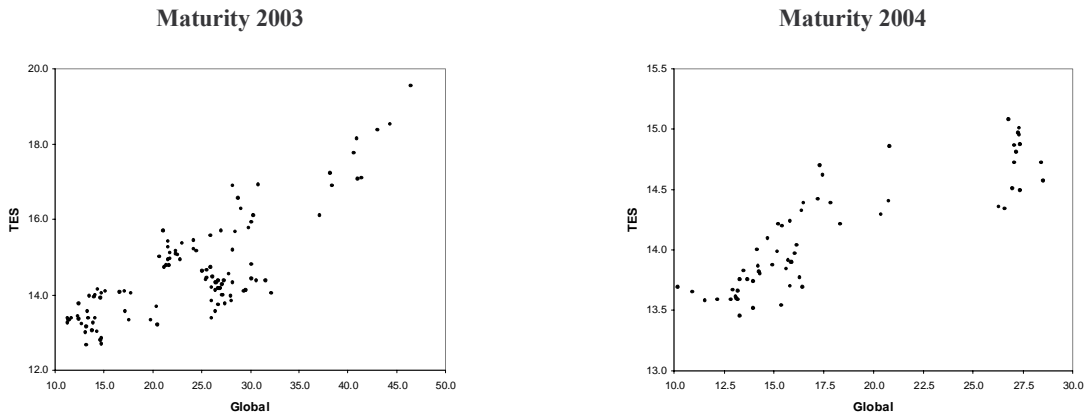


Figure 2. Yield of TES and Global securities (sorted by duration) per year of maturity



$$\Delta \varepsilon_t^e = \rho \varepsilon_{t-1}^e + \sum_j \delta_j \Delta \varepsilon_{t-j}^e + \mu_t^e \quad (8)$$

where the disturbance term  $\mu_t^e$  is white-noise while  $\rho$  should be  $-2 < \rho < 0$  for rejecting the null of no cointegration ( $H_0 : \rho = 0$ ) between the yields of each maturity. If  $H_0$  is rejected, the Granger representation theorem implies the following error correction model:

$$\Delta y_t^e = \lambda_e + \alpha_e \varepsilon_{t-1}^e + \sum_j \delta_{e,j}^e \Delta y_{t-j}^e + \sum_j \delta_{e,j}^i \Delta y_{t-j}^i + w_t^e \quad (9a)$$

$$\Delta y_t^i = \lambda_i + \alpha_i \varepsilon_{t-1}^e + \sum_j \delta_{i,j}^e \Delta y_{t-j}^e + \sum_j \delta_{i,j}^i \Delta y_{t-j}^i + w_t^i \quad (9b)$$

where  $w_t^e$  and  $w_t^i$  are disturbance terms  $iid \sim (0, \sigma_e^2)$  and  $iid \sim (0, \sigma_i^2)$ , respectively. Equations (9a) and (9b) suggest that adjustment of any deviation from the long run equilibrium is symmetric in the sense that it is proportional to the absolute value of the error correction term. However, if the adjustment process happens to be asymmetric as evidence of the influence on returns of changing perception of country risks by different type of traders coexisting in the market, then the above test is misspecified. Enders and Granger (1998) and Enders and Siklos (2001) have shown that if the adjustment process is asymmetric an alternative specification for equation (8) is the threshold autoregressive (TAR) model which takes the form:

$$\Delta \varepsilon_t^e = I_t \rho_1 \varepsilon_{t-1}^e + (1 - I_t) \rho_2 \varepsilon_{t-1}^e + \sum_j \delta_j \Delta \varepsilon_{t-j}^e + \mu_t^e \quad (10)$$

where  $I_t$  is a Heaviside indicator function based on the threshold value  $\tau$  :

$$I_t = \begin{cases} 1 & \text{si } \varepsilon_{t-1}^e \geq \tau \\ 0 & \text{si } \varepsilon_{t-1}^e < \tau \end{cases} \quad (11)$$

This asymmetric model then replaces the symmetric version of equation (8). The necessary and sufficient conditions for the stationarity of  $\{\varepsilon_t^e\}$  is  $\rho_1 < 0$  and  $\rho_2 < 0$  and  $(1 + \rho_1)(1 + \rho_2) < 1$  for any value of  $\tau$ . When  $\rho_1 = \rho_2$ , then the Engle-Granger symmetric test arises as a special case.

If the threshold value  $\tau$  is equal to zero, we estimate a regression equation in the form (10) and compare the larger of the  $t$ -statistics corresponding to the null hypothesis  $H_0 : \rho_i = 0$  for  $i = 1, 2$  as well as the  $F$ -statistic for the null  $H_0 : \rho_1 = \rho_2 = 0$  with the appropriate critical values associated to  $t$ -Max and  $F$ -test  $\Phi$ , respectively (Enders and Siklos, 2001).

If the threshold value  $\tau$  is unknown, the procedure requires the estimation of the residuals from equation (7), then these residuals are sorted in ascending order,  $\hat{\varepsilon}_{1,\tau}^e < \hat{\varepsilon}_{2,\tau}^e < \dots < \hat{\varepsilon}_{T,\tau}^e$ , where  $T$  is the number of usable observations. The largest and smallest 15% of the  $\{\hat{\varepsilon}_{t,\tau}^e\}$  are discarded while the remainder 70% are considered as potential

thresholds. For each of these possible thresholds we estimate an equation of the form (10) and (11). The model with the smallest residual sum of squares is chosen to obtain the value of  $\tau$ . If this is the case, the critical values are  $t$ -Max\* and  $F$ -test  $\Phi^*$ , respectively (see Enders and Siklos, 2001).

As an alternative adjustment specification Enders and Granger (1998) have suggested the momentum-threshold autoregressive (M-TAR) models where the sequence of disturbances  $\{\varepsilon_t^e\}$  exhibits more momentum in one direction than the other. In this case, the Heaviside indicator function  $M_t$  is based on the threshold value  $\tau$  and the previous period's change in  $\varepsilon_{t-1}^e$  as follows<sup>10</sup>:

$$M_t = \begin{cases} 1 & \text{if } \Delta\varepsilon_{t-1}^e \geq \tau \\ 0 & \text{if } \Delta\varepsilon_{t-1}^e < \tau \end{cases} \quad (12)$$

In this case, the series of the first difference of the residuals is sorted in ascending order,  $\Delta\varepsilon_{1,\tau}^e < \Delta\varepsilon_{2,\tau}^e < \dots < \Delta\varepsilon_{T,\tau}^e$  while the rest of the procedure continues the same. The critical values of  $t$  and  $F$  statistics change consequently.

## 5. Results

Here we present the results of the testing procedure just described in the previous section. At first we show the estimates of equation (7) for yields of bonds with maturity in 2003:

$$y_{03,t}^e = -48.292 + 4.923 y_{03,t}^i + \varepsilon_{03,t}^e \quad (13)$$

(-9.378) (14.036)

The estimates of equation (13) have the right sign but [together with the estimates of equation (14) below] according to the significance they show that capital mobility is less than perfect. Since the sign is negative, the coefficient  $\beta_0$  suggests that global bonds are less risky

than TES bonds in terms of currency risks while the sign and magnitude of the coefficient  $\hat{\beta}_1$ , suggest that in terms of country risk, global bonds are riskier than the TES ones. In addition, notice that the coefficient  $\hat{\beta}_1$  of equation (7) is far from unity as predicted by expression (2) and the restriction that  $\hat{\beta}_1 = 1$  is clearly rejected. A number of reasons could help us to explain this result. First, it might be suggesting a different perception of country risk depending on the place where Colombia allocates the securities: home or abroad. That is, the investors might think that the Colombian government would be more tempted to honor first its compromises with TES than with global (yankees and so on) bonds. The consequence of this believe is that –contrary to what we assumed from the beginning- there are differences in the country risk associated to each type of instrument revealed by the estimate of  $\hat{\beta}_1$  which might be a symptom of incomplete, segmented<sup>11</sup> and imperfect financial markets (see footnote 1). A second reason could be found in the way in which we have set the exchange rate expectations. Past depreciation of the exchange rate might not be a good *proxy* for the expected value since the information set could be very different. In addition, we did not consider any possible deviations of the past depreciation rate with respect to any equilibrium level of the exchange rate. Third, transaction costs, information costs and differences of taxes on returns, might also appear as a reason to explain the above estimate of  $\hat{\beta}_1$ . Finally, to make the comparison of the yields of the coupon bonds we focused only on duration disregarding the convexity of the instruments, a concept associated to the second derivative of the bond price with respect to the yield. It is possible that we are losing important information with this simplification.

For the yields of securities with maturity in 2003 ( $\mathcal{Y}_{03,t}^e, \mathcal{Y}_{03,t}^i$ ), equations (13) and (14) show that one can reject the null hypothesis of no cointegration according to the *t*-Max\* criterion. This evidence of cointegration is found when a *M*-TAR model with a threshold of -2.456 is fitted, after following all the steps suggested by Enders and Siklos (2001):

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<sup>10</sup> The *M*-TAR adjustment specification can be helpful when the monetary authority is willing to take measures if expectations of future inflation are increasing.

<sup>11</sup> This obviously includes regulations such as bounds in holding some securities and so on.

$$\begin{aligned}
\Delta \hat{\varepsilon}_{03,t}^e = & -0.034 - 0.091 M_t \hat{\varepsilon}_{03,t-1}^e - 0.280 (1 - M_t) \varepsilon_{03,t-1}^e - 0.154 \Delta \hat{\varepsilon}_{03,t-1}^e \\
& (-0.158)(-1.791) \quad (-2.161) \quad (-1.493) \\
& - 0.072 \Delta \hat{\varepsilon}_{03,t-2}^e + 0.293 \Delta \hat{\varepsilon}_{03,t-3}^e + 0.131 \Delta \hat{\varepsilon}_{03,t-4}^e + \mu_{03,t}^e \\
& (-0.693) \quad (-2.881) \quad (1.274)
\end{aligned} \tag{14}$$

$$\begin{aligned}
t - Max_{CV,90\%}^* &= -1.660; \Phi_{est}^* = 3.695; \Phi_{CV,90\%}^* = 5.520 \\
D - W &= 2.016; p - value[Q(20)] = 0.998; R^2 = 0.207.
\end{aligned}$$

The finding of cointegration with *M*-TAR adjustment allows the estimation of the following error correction models<sup>12</sup> (equations 15 and 16).

$$\begin{aligned}
\Delta \hat{y}_{03,t}^e = & 0.224 - 0.024 M_t \hat{\varepsilon}_{03,t-1}^e - 0.175 (1 - M_t) \hat{\varepsilon}_{03,t-1}^e + 0.204 \Delta y_{03,t-1}^e \\
& (1.518)(-0.827) \quad (-2.285) \quad (2.066) \\
& - 0.466 \Delta y_{03,t-1}^i + 0.380 \Delta y_{03,t-2}^i - 0.086 \Delta y_{03,t-3}^i + 0.666 \Delta y_{03,t-4}^i + w_{03,t}^e \\
& (-1.172) \quad (0.855) \quad (-0.191) \quad (1.642)
\end{aligned} \tag{15}$$

$$D - W = 1.971; p - value[Q(20)] = 0.241; R^2 = 0.140.$$

$$\begin{aligned}
\Delta \hat{y}_{03,t}^i = & 0.077 + 0.008 M_t \hat{\varepsilon}_{03,t-1}^e + 0.014 (1 - M_t) \hat{\varepsilon}_{03,t-1}^e + 0.011 \Delta y_{03,t-1}^e \\
& (1.853) (1.002) \quad (0.665) \quad (0.397) \\
& - 0.489 \Delta y_{03,t-1}^i - 0.166 \Delta y_{03,t-2}^i + 0.146 \Delta y_{03,t-3}^i + 0.288 \Delta y_{03,t-4}^i + w_{03,t}^i \\
& (-4.352) \quad (-1.324) \quad (1.150) \quad (2.517)
\end{aligned} \tag{16}$$

$$D - W = 1.850; p - value[Q(20)] = 0.680; R^2 = 0.273.$$

<sup>12</sup> Previously, the error correction approach was used by Correa (1992) and Herrera (1993). Correa, reported an error correction model where the cointegrating relationship between the inflation differential and the interest rate differential, assuming neither risks nor partial adjustment of prices, was not completely satisfactory according to her interpretation. Herrera, used the same technique to find evidence of a long run cointegration between the domestic and the external rates (and a supposed time-varying risk premium derived from an error correction model not well specified). Finally, Steiner et al (1993), by using a different approach, found that the hypothesis of an almost zero (constant) risk premium could not be rejected.

The  $t$ -statistics for the error-correction terms indicate that the yield of assets traded domestically (TES),  $y_{03,t}^i$ , is weakly exogenous and the yield of assets traded abroad (global),  $y_{03,t}^e$ , adjusts to deviations from the long-run equilibrium if  $\Delta \hat{\epsilon}_{03,t}^e < -2.456$ .

The dynamics can be analyzed by observing that from equation (13) we have:

$$\Delta y_{03,t}^e = 4.923 \Delta y_{03,t}^i + \Delta \hat{\epsilon}_{03,t}^e$$

and considering the threshold we have:

$$\Delta \hat{\epsilon}_{03,t-1}^e = \Delta y_{03,t-1}^e - 4.923 \Delta y_{03,t-1}^i < -2.456$$

which means that:

$$\frac{1}{4.923} \Delta y_{03,t}^e + \frac{1}{4.923} 2.456 < \Delta y_{03,t}^i$$

Hence any previous change in the local debt yield higher than the left-hand side of this condition produces a quick adjustment towards the long run equilibrium: Changes in  $y_{03,t-1}^i$  greater than the left-hand side of the condition are reflected in  $\Delta y_{03,t}^e$  [equation (15)]. The explanation might be linked to the perception of the country risks first captured in the local market by heterogeneous agents.

As for securities with maturity in 2004, equations (17)<sup>13</sup> and (18) allow us to reject the null hypothesis of no cointegration. In addition, when we estimate the residuals of (17) in the form of a TAR using a threshold  $\tau = 0$  we find evidence of cointegration between the two yields under  $t$ -Max criterion<sup>14</sup> (equation 19). However, since the restriction  $\rho_1 = \rho_2$  cannot be rejected at 95% significance level we continue with the linear adjustment.

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<sup>13</sup> The analysis of equation (13) above also serves here to explain the estimate of  $\beta_1$ .

<sup>14</sup> The critical values were generated with a 10.000 replications Monte Carlo experiment where the coefficients are equal to those of equation (19). The reason for the simulation is that the tables of Enders and Siklos consider up to four lagged changes.

$$y_{04,t}^e = -131.518 + 10.573 y_{04,t}^i + \hat{\varepsilon}_{04,t}^e \quad (17)$$

(-11.578) (13.141)

$$\Delta \hat{\varepsilon}_{04,t}^e = -0.286 \hat{\varepsilon}_{04,t-1}^e - 0.023 \Delta \hat{\varepsilon}_{04,t-1}^e + 0.297 \Delta \hat{\varepsilon}_{04,t-2}^e - 0.051 \Delta \hat{\varepsilon}_{04,t-3}^e$$

(-3.164) (-0.174) (2.262) (-0.402)

$$+ 0.247 \Delta \hat{\varepsilon}_{04,t-4}^e + 0.357 \Delta \hat{\varepsilon}_{04,t-5}^e + \mu_{04,t}^e \quad (18)$$

(1.925) (2.804)

$$CV = -3.114; D - W = 1.936; p - value[Q(14)] = 0.892; R^2 = 0.309$$

$$\Delta \hat{\varepsilon}_{04,t}^e = -0.271 I_t \hat{\varepsilon}_{04,t-1}^e - 0.303 (1 - I_t) \hat{\varepsilon}_{04,t-1}^e - 0.022 \Delta \hat{\varepsilon}_{04,t-1}^e + 0.296 \Delta \hat{\varepsilon}_{04,t-2}^e$$

(-2.311) (-2.491) (-0.165) (2.221)

$$- 0.057 \Delta \hat{\varepsilon}_{04,t-3}^e + 0.245 \Delta \hat{\varepsilon}_{04,t-4}^e + 0.356 \Delta \hat{\varepsilon}_{04,t-5}^e + \mu_{04,t}^e \quad (19)$$

(-0.434) (1.882) (2.761)

$$t - Max_{CV, 90\%} = -1.95; \Phi_{est} = 4.933; \Phi_{CV, 90\%} = 5.20$$

$$D - W = 1.929; p - value[Q(14)] = 0.894; R^2 = 0.309$$

Looking at equation (17), both the level and signs of the coefficients suggest that capital mobility is much less than perfect. As in the case of bonds with maturity at 2003, the negative sign of the coefficient  $\beta_0$  suggests that global bonds are less risky than TES bonds in terms of currency risks while the sign and magnitude of the coefficient  $\hat{\beta}_1$ , suggest that TES bonds are less risky than global regarding country risk. The level of the estimates deserve some attention. Firstly, intuitively, equations (13) and (17) show that risk premium is greater for TES bonds with maturity in 2004 than for those with maturity in 2003 [ $131.518 \div 10.573 (=12.43) > (9.81=) 48.292 \div 4.923$ ], which means that the risk premium is time varying and increasing overtime. Secondly, the estimate linked to the country risk is greater for global securities with maturity in 2004 than for bonds with maturity in 2003. However, the changes in the estimates are not unambiguously nonlinear [ $131.518 \div 48.292 (=2.72) > (2.14=) 10.573 \div 4.923$ ].

The finding of cointegration permits the estimation of the error correction models of equations (20) and (21). As in the case of assets with maturity 2003, the  $t$ -statistics for the error-correction terms indicate that the yield of assets traded domestically (TES),  $y_{04,t}^i$ , is weak exogenous (see equation 22). In equation (21) we see that the changes in the yield of assets traded abroad (global bonds),  $\Delta y_{04,t}^e$ , adjust to deviations from the long-run equilibrium symmetrically, for that reason the dynamics of this adjustment process is much simpler than the former.

$$\begin{aligned} \Delta \hat{y}_{04,t}^e = & 0.263 - 0.156 \hat{\varepsilon}_{04,t-1}^e + 0.433 \Delta y_{04,t-1}^e + 0.358 \Delta y_{04,t-5}^e + 0.367 \Delta y_{04,t-8}^e - 3.909 \Delta y_{04,t-1}^i \\ & (2.277) (-3.807) \quad (3.743) \quad (3.285) \quad (3.914) \quad (-5.790) \\ & - 2.112 \Delta y_{04,t-2}^i + 1.290 \Delta y_{04,t-3}^i - 2.518 \Delta y_{04,t-4}^i - 4.930 \Delta y_{04,t-5}^i - 1.012 \Delta y_{04,t-8}^i + w_{04,t}^e \\ & (-2.475) \quad (1.602) \quad (-3.047) \quad (-6.545) \quad (-1.417) \end{aligned}$$

$$D - W = 2.121; p - \text{value}[Q(13)] = 0.315; R^2 = 0.728.$$

(21)

$$\begin{aligned} \Delta \hat{y}_{04,t}^i = & 0.005 + 0.008 \hat{\varepsilon}_{04,t-1}^e + 0.028 \Delta y_{04,t-1}^e + 0.010 \Delta y_{04,t-5}^e + 0.027 \Delta y_{04,t-8}^e - 0.454 \Delta y_{04,t-1}^i \\ & (0.210) (0.991) \quad (1.099) \quad (0.439) \quad (1.336) \quad (-3.045) \\ & - 0.038 \Delta y_{04,t-2}^i + 0.143 \Delta y_{04,t-3}^i + 0.256 \Delta y_{04,t-4}^i + 0.0231 \Delta y_{04,t-5}^i - 0.231 \Delta y_{04,t-8}^i + w_{04,t}^i \\ & (-0.206) \quad (0.808) \quad (1.405) \quad (0.138) \quad (-1.468) \end{aligned}$$

$$D - W = 2.046; p - \text{value}[Q(13)] = 0.984; R^2 = 0.401.$$

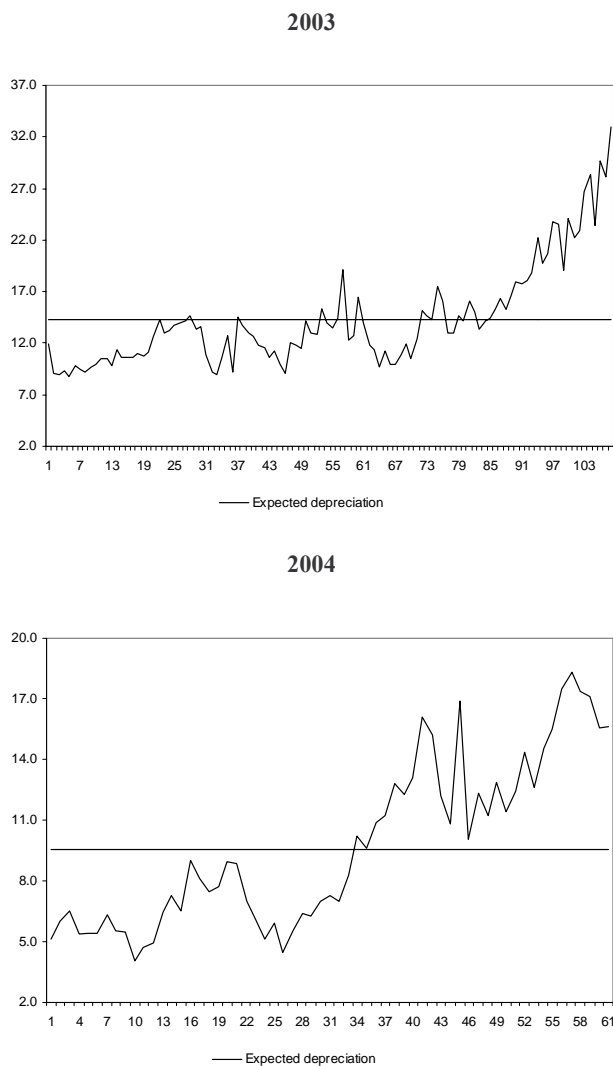
(22)

According to our results of the presence of long-run relationships between the two yields and a constant risk premium for each year of maturity, based upon equations (13) and (17), we can compute the difference between the TES yield (plus the risk premium) and the global yield to form an expectation for the depreciation rate ex-currency risk and (almost) ex-country risk. The expected average exchange rate depreciations are about 14.3%, and 9.6% for



2003 according to the securities with maturities in 2003 and 2004, respectively<sup>15</sup>. In Figure 3, where we present these results with the operation based on equations (13) and (17) sorted by duration, it is noticeable that as duration is increasing the expected rate of depreciation is also increasing.

**Figure 3. Expected depreciation rate in 2003 according to the maturity of bonds**



<sup>15</sup> It might appear wrong to compute the mean of series such as those presented in Figure 3 given the trend exhibited by the series. However, we have to remember that the variable has been ordered by duration. Otherwise it would look much erratic.

## 6. Concluding remarks

In this work we compare the yields of coupon bonds issued by the Colombian government in order to check the hypothesis of (short and medium term) capital mobility for Colombia. For making the comparison between yields of coupon bonds (with different coupon rates) for each year of maturity (2003 and 2004) we used the concepts of immunization and Macaulay duration.

We find evidence of long run comovements between the yields of TES and global securities issued by the Colombian government, which is also evidence of (short and medium term) capital mobility. However, according to the results, the mobility is much less than perfect. We also find evidence of weak exogeneity of the yields of TES bonds which allows us to represent the system by a single equation. We associate the estimates of risk premium with currency risks since the country risks are included into the yields.

After applying the Enders and Siklos (2001) approach we end up with two error correction models for the yields corresponding to assets with maturity in 2003 and 2004, respectively. For 2003 we found that the adjustment towards the long run equilibrium is nonlinear which might be evidence of the coexistence of different types of traders in the market with distinct perceptions of risks. For the M-TAR model that we have estimated, the asymmetric adjustment is triggered by a change in the error term when the threshold is equal to -2.456. The velocity of adjustment is different when deviations are above the long run equilibrium than when they are below. In particular, we find that only when  $\Delta \varepsilon_{03,t-1}^e < -2.456$  the adjustment is less persistent (note that  $|\rho_1| < |\rho_2|$ ). For bonds with maturity in 2004, the adjustment around the long run equilibrium is symmetric while the threshold is zero. In this case the differential interest rate plus a constant risk premium is a symmetric attractor. Finally, we obtain that the implied value of the expected depreciation rate for 2003 is about 14.3% and 9.6% according to securities with maturities in 2003 and 2004, respectively.

## References

- Bayoumi, T., 1990, Saving-investment correlations, *IMF Staff paper* 37, 360-387.
- Baxter, M., 1994, Real exchange rates and real interest differentials. Have we missed the business-cycle relationship?, *Journal of Monetary Economics*, 33, 5-37.
- Campbell, J.Y. and R.H. Clarida, 1987, The dollar and the real interest rates, *Carnegie-Rochester Conference Series on Public Policy*, 27, 103-140.
- Cárdenas, M., and Garay, L. J. (Eds), 1993, Macroeconomía de los flujos de capital en Colombia y América Latina. Tercer Mundo Editores, Fedesarrollo y Fescol, Colombia.
- Correa, P., 1992, Paridad entre la tasa de interés real interna y externa: notas sobre el caso colombiano. *Coyuntura Económica*, Vol. XXII, No., April. 139-50.
- Dooley, M., Frankel, J., Mathieson, D.J., 1987, International capital mobility: what do the saving investment correlation tells us? *IMF Staff papers*, 34, 503-530.
- Dornbusch, R., 1976. Expectations and exchange rate dynamics, *Journal of Political Economy*, Vol. 84, 1161-1176.
- Dumas, B., 1994. Partial Equilibrium versus General Equilibrium Models of the International Capital Market. In: *Van der Ploeg, F. (Ed.), The Handbook of International Macroeconomics*, Blackwell, Oxford and Cambridge, pp. 301-347.
- Edwards, S. and M. Khan, 1985, Interest rate determination in developing countries: A conceptual framework. *IMF Staff Papers*, Vol. 32, No. 3, 377-403.
- Enders, W., and Granger, C. W. J., 1998, Unit-Root Tests and Asymmetric Adjustment with an example using the Term Structure of Interest Rates, *Journal of Business & Economic Statistics*, 16, 304-311.
- Enders, W., Siklos, P., 2001, Cointegration and Threshold Adjustment, *Journal of Business & Economic Statistics*, Vol. 19, No 2, 166-176.
- Engle, R.F., and Granger, C.W.J., 1987, Cointegration and Error Correction: Representation, estimation and testing, *Econometrica*, 55, 251-276.
- Fabozzi, F., 2000. Bond Markets, Analysis and Strategies. Fourth Edition. Prentice Hall, New Jersey.
- Farrell, J., 1997. Portfolio Management: Theory and Application. Second Edition. MCGraw Hill International Editions.

Feldstein, M., and Bacchetta, P., 1991, National saving and international investment. In: *Bernheim, D., Shoven, J. (Eds.), National Saving and Economic Performance*. University of Chicago Press, pp. 201-220.

Feldstein, M., and Horioka, C., 1980, Domestic saving and international capital flows. *Economic Journal* N° 90, 314-329.

Frankel, J.A., 1979, On the mark: A theory of floating exchange rates based on real interest differentials, *American Economic Review*, Vol. 69, No.4, 610-622.

Ghosh, A.R., 1995, International capital mobility amongst the major industrialized countries: too little or too much?. *Economic Journal*, 105, 107-128.

Herrera, S., 1993, Movilidad de capitales en la economía colombiana, en: *M. Cárdenas, and L. J. Garay (Eds.), Macroeconomía de los flujos de capital en Colombia y América Latina*, Tercer Mundo Editores, pp. 167-188.

Isard, P., 1983. An Accounting Framework and Some Issues for Modeling How Exchange Rates Respond to the News. In: *Frenkel, J.A. (Ed.), Exchange Rates and International Macroeconomics*, University of Chicago Press, Chicago.

Meese, R. and K. Rogoff, 1988. Was it real? The exchange rate differential relation over the modern floating-rate period. *Journal of Finance*, XLIII, 4, 933-948.

Obstfeld, M., 1995, International capital mobility in the 1990s. In P.B. Kenen (ed.) *Understanding interdependence: the macroeconomics of the open economy*. Princeton University Press.

Pilbeam, K., 1992, *International Finance*. MacMillan, City University, London.

Posada, C.E., 1998, La tasa de interés: el caso colombiano del siglo XX (1905-97), *Ensayos sobre política económica*, No. 33. 5-60.

Steiner, R., Suescún, R., and Melo, F., 1993, Flujos de capital y expectativas de devaluación. En: *Cárdenas, M., y L. J. Garay (Eds), Macroeconomía de los flujos de capital en Colombia y América Latina*, Tercer Mundo Editores, pp. 189-220.

Sachs, J., 1983, Aspects of the current account behaviour of OECD economies”, in *Recent issues in the theory of flexible exchange rates*, E. Claassen and P. Salin (eds.), North Holland.

Taylor, Alan M., 1994, Domestic savings and international capital flows reconsidered. *Working paper 4892, National Bureau of Economic Research*, October.

Tesar, L.L., 1991, Saving, investment and international capital flows. *Journal of International Economics*, 31, 55-78.