Uncovered Interest Parity and the USD/COP Exchange Rate

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Abstract

This paper tests the uncovered interest parity (UIP) hypothesis for the USD/COP exchange rate, using weekly data for the period from January 1994, when Colombia introduced its crawling band exchange rate regime, to August 2002. The study yields several interesting results. For the period October 1996 to August 2002 the UIP hypothesis receives relatively strong support, even if this is weakened towards the end of the period. This is in stark contrast with the almost unanimous rejection of UIP shown by the literature. UIP is, furthermore, tested for a duration of time of 3, 6 and 12 months, and in line with other studies, the validity of the UIP relationship increases with the term of the investment. However, we suspect that the strong support for UIP might be a temporary occurrence due to the fact that Colombia during this period went through a considerable macroeconomic transition, where a high rate of inflation were brought down from double-digit to single-digit levels.

JEL classification:

Key Words:

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1 Introduction

Uncovered interest parity (UIP) is a classical theory in international economics, and it is a fundamental building block in most monetary models. However, empirically it has proven to be a miserable failure. According to UIP, the interest differential between two countries should on average equal the expected change in the exchange rate. Under rational expectations, the actual ex-post change in the exchange rate should be a good proxy for the expected change. Economies with high interest rates should, consequently, have depreciating currencies. Empirical studies have, however, shown that in the majority of such cases the currencies have appreciated, particularly in the short to medium term. A strong consensus has, therefore, developed in the literature that UIP works poorly.

In this paper, UIP is tested for the rate of exchange of the Colombian peso (COP) to the US dollar (USD) during recent years. Weekly data is used for the period January 1994 to August 2002. Relatively strong support for UIP is found from October 1996 and onwards, even if this seems to weaken towards the end of the period. These findings are in stark contradiction with the failure of UIP in other studies, and in that sense the results are somewhat surprising. The strong support for UIP might, however, be temporary and might have been caused by the fact that Colombia during this period went through a considerable macroeconomic transition where the inflation rate was brought down from double-digit to single-digit levels.

In section 2, the theoretical framework is laid out, and the empirical literature on UIP is briefly surveyed. Section 3 concisely describes the recent exchange rate regimes that have been in place in Colombia, and in section 4 the data is analysed and the results presented. The paper ends with a brief summary in section 5.

2 Uncovered Interest Parity: Theory and Empirical Findings

UIP is a fundamental theory in exchange rate economics. However, the literature has developed a strong consensus, that UIP performs poorly in empirical studies. In the first section of this chapter the UIP hypothesis is derived. Thereafter the empirical literature on UIP is briefly surveyed. The chapter ends with a section discussing why exchange rates behave differently in the short and long term, and why UIP might perform better in the longer term.

2.1 The Uncovered Interest Parity Hypothesis

The hypothesis of UIP states that market forces equilibrate the expected return on an uncovered investment in a foreign currency to the return on the riskless option of investing in local currency.[1] If E_t is the expectations operator and E_tS_{t+T} denotes the expected value at time t of the spot exchange rate S at time $t + T$, the UIP hypothesis can be expressed as

$$E_tS_{t+T}\,(1 + r_t^*) = S_t\,(1 + r_t) \qquad (2.1)$$

where S_t is the exchange rate at time t, and r_t and r_t^* are domestic and foreign interest rates at time t. The exchange rate S is measured in units of domestic currency per unit of foreign currency.

Rewriting equation (2.1) yields

$$(E_tS_{t+T} - S_t)\,/\,S_t = (r_t - r_t^*)\,/\,(1 + r_t^*) \qquad (2.2)$$

[1] This hypothesis emerges as a special case from the conceptual framework generally referred to as the international asset pricing model. See Hodrick (1987) or Meese (1989) for a general discussion as well as for references to the literature.

If s is the logarithmic value of S, this would with approximations imply that[2]

$$E_t s_{t+T} - s_t \approx r_t - r_t^* \tag{2.3}$$

Note that UIP can be written for any duration of time T. Consequently, if UIP is valid, the whole expected future time path of the spot exchange rate can be derived from the yield curves of the domestic and foreign interest rates.

The UIP hypothesis in itself as expressed by equation (2.3) is, however, not very interesting. The absence of reliable data for the expected future exchange rate makes it difficult to reach definitive conclusions about its validity.

Testing UIP normally involves combining it with the assumption that investors have rational expectations and that the *expected* future spot rate, even if not observable in itself, can be regarded as an unbiased predictor of the *actual* future spot rate. It is, therefore, assumed that on average

$$E_t s_{t+T} = s_{t+T} \tag{2.4}$$

and equation (2.3) can be rewritten as

$$s_{t+T} - s_t \approx r_t - r_t^* \tag{2.5}$$

[2] Note that for $1 + r_t^*$ and $1 + r_t$ close to 1,

$$E_t S_{t+T} / S_t = 1 + (r_t - r_t^*) / (1 + r_t^*) \approx 1 + r_t - r_t^*$$

and therefore

$$E_t s_{t+T} - s_t \approx r_t - r_t^*$$

We have so far assumed that the foreign and domestic investments are of equal risk, or, alternatively, that investors are risk neutral. However, if investors are risk averse and if the domestic investment carries a larger default risk, equation (2.5) can be rewritten as

$$s_{t+T} - s_t \approx r_t - r_t^* - \varphi_t \tag{2.6}$$

where φ_t represents the risk premium of the domestic interest rate associated with the difference in default risk between the domestic and foreign investments. In practice this should equal the interest rate spread between domestic and foreign investments when both are denominated in the same currency.

The relationships stated by equation (2.5) and (2.6) are tested by estimating[3]

$$(s_{t+T} - s_t) = a + b\,(r_t - r_t^*) + \varepsilon_t \tag{2.7}$$

or

$$(s_{t+T} - s_t) = a + b\,(r_t - r_t^*) - c\,\varphi_t + \varepsilon_t \tag{2.8}$$

where a, b and c are regression coefficients and ε_t is the forecasting error realised at time $t + T$ from a forecast of the exchange rate made at time t. For UIP to be valid, a should approximately equal zero, b should approximately equal one, and c should approximately equal minus one. The condition that the parameter estimate of b should be approximately equal to unity, is generally referred to as the *unbiasedness hypothesis*.[4] If equation (2.7) is estimated for investments of different risk (under the assumption of risk averse investors), the parameter estimate of a should, furthermore,

[3] See Meese (1989) for a discussion on these specification forms.
[4] Meredith and Chinn (1998), p. 4.

reflect the difference in risk, and should, therefore, not equal zero. In the case where the domestic investments are more risky than the foreign investments, the former should yield an interest premium over the latter, even if the exchange rate is expected to remain constant. The parameter estimate of a should, thus, in such a case be negative.

2.2 Review of Empirical Studies

Even if UIP in its simplicity is a very attractive theoretical hypothesis, empirical tests have almost universally failed it. The standard practice is to estimate equation (2.7) using ordinary least squares or generalised least squares, and then to investigate how close the parameter estimates are to their expected values.[5] As discussed in the previous section, the slope parameter b should approximately be equal to unity. If interest rate data are coming from investments of similar risk, which have been the case in many studies, the constant a should, furthermore, not be significantly different from zero.

An extensive literature has by now rejected the UIP hypothesis empirically.[6] In Froot (1990), 75 published studies are surveyed. The large majority of these reject the unbiasedness hypothesis that the slope parameter b in equation (2.7) should equal unity. Not only are the parameter estimates of b in these studies significantly less than one, but in most of the studies they are negative. None of the studies yields a parameter estimate exceeding unity. The average estimate of b in fact equals minus 0.88. Similar results have been reported by other surveys, including MacDonald and Taylor (1992), Isard (1995), and Lewis (1995).

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[5] Some later papers have used other techniques, such as, for example cointegration frameworks, but the results are similar. See, for example, Wu (1999).

[6] Initial studies include Bilson (1981), Longworth (1981), as well as the seminal paper by Meese and Rogoff (1983).

However, the common perception that the failure of UIP indicates that short-term exchange rate movements are best characterised as a random walk might not be true. As discussed, most studies have found exchange rates to move *inversely* with interest rate differentials.[7]

A number of studies have, nevertheless, concluded that even if UIP fails in the short term, its validity increases with the term of the investment.[8] Meredith and Chinn (1998) tests the UIP hypothesis using 10-year government bond yields of the G-7 countries. Their results differs starkly from the short-term studies. For all the currencies, the estimated slope coefficients are positive, with four of the six values lying closer to unity than to zero, and in all of the cases except for one, the hypothesis that *b* equals zero can be rejected.

2.3 Exchange Rate Determination in the Short and Long Term

We concluded in the earlier section that even if there is little empirical support for UIP in the short term, long-term studies have yielded much more favourable results. Two possible explanations are here suggested.

First, in the short term there might exist a second relationship between the interest rates and exchange rate, in addition to that stipulated by UIP, as emphasised by McCallum (1994). Monetary authorities in many countries use short-term interest rates as a monetary policy instrument. These are adjusted in response to undesired exchange rate movements. A negative exogenous shock to the economy leads to a depreciation of the exchange rate. To counter this, short-term interest rates are raised. When the shock fades away, the exchange rate appreciates, and short-term interest

[7] Meredith and Chinn (1998).
[8] See, for example, Berk and Knot (2001), McCallum (1994), and Meredith and Chinn (1998).

rates are lowered.[9] In this case, high interest rates are, in fact, followed by an exchange rate appreciation, which is in line with empirical findings, but which is in direct conflict with what is stated by the UIP hypothesis.

Second, a number of studies have shown that, due to incomplete information in the short term, the behaviour of market participants is to a large extent based on *technical analysis* of short-term trends or other patterns in the observed behaviour of the exchange rate.[10] In support of such behaviour, simulations have shown short-term trading strategies based on technical analysis to generate significant profits.[11] The long-term behaviour of exchange rates are, on the other hand, much more governed by fundamentals.

Models have been developed where *feedback traders* coexist with *fundamentalists* as market participants.[12] The former, which are also sometimes referred to as *chartists*, base their trading strategies on the recent history of exchange rates, while the latter base their strategies on analysis of economic fundamentals. In these types of models, the fundamentalists have the predominant influence of exchange rates in the long term. However, risk aversion together with substantial uncertainties regarding news and new information, leads to feedback traders dominating the market in the short term. This implies that short-term exchange rates will vary much more widely than is justified by changes in fundamentals. Feedback trading can be regarded as rational in the context of incomplete information and a continuous learning process. Information generated by the trading process itself gives feedback on the assessment of other market participants.[13]

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[9] Even if central banks are not intervening to defend the exchange rate, they often tend to hold short-term interest rates relatively constant. This implies that short-term interest differentials will vary much less than other variables influencing the risk of the country, which might be a potential explanation of the negative parameter estimates of b. See Isard (1988) as well as Boyer and Adams (1988).
[10] See, for example, Taylor and Allen (1992).
[11] See Cumby and Modest (1987), Dooley and Shafer (1983), and Sweeny (1986).
[12] See, for example, Kyle (1985), Frankel and Froot (1990), and Cutler, Poterba and Summers (1990).

According to the theories presented here, short-term exchange rate behaviour will, consequently, be dominated by the behaviour of feedback traders, by short-term economic shocks, and by the response of monetary policy to such shocks, while long-term exchange rate to a large extent are governed by economic fundamentals. Exchange rates will, therefore, be much more volatile in the short term than in the long term.

[13] See, for example, Lyons (1993).

3 The USD/COP Exchange Rate

In this study we are using data from January 1994 until present. The USD/COP exchange rate was characterised by a crawling band regime from January 1994 until September 1999, after which it has been floating freely. The first section of this chapter discusses the different exchange rate regimes and what impact they have had on the exchange rate. The second section looks at the change in the exchange rate together with the interest rate differentials between the United States and Colombia.

3.1 The Different Exchange Rate Regimes in Colombia

From 1967 and up until 1991, the exchange rate regime in Colombia was defined by a crawling peg. The Colombian peso was pegged to the US dollar at a pre-specified exchange rate and was not allowed to depart significantly from this rate. This exchange rate was, furthermore, devalued daily at a pre-determined and continuous devaluation rate. The exchange rate regime was combined with a system of thorough capital controls, where all foreign exchange transactions had to be made through the Banco de la República. This period had earned Colombia a reputation for outstanding macroeconomic stability. The rate of inflation was high but stable at between 20 and 30 per cent per year, and a recession had not been experienced since 1931. However, things were to change dramatically during the second part of the 1990s.[14]

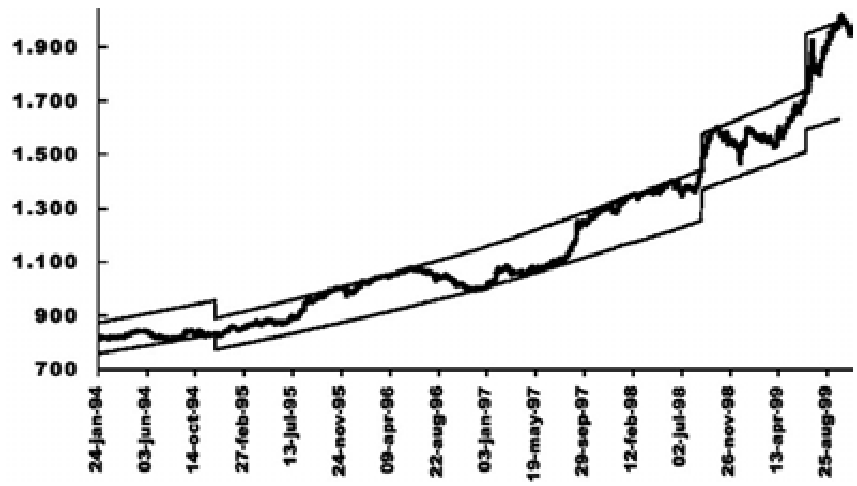
The crawling peg regime was abolished in June 1991. In 1989 the rate of daily devaluations had been increased to counter a sharp fall in international coffee prices and a deterioration in the trade balance. The latter was due to the opening up of the economy. However, inflation was running at a high rate, which required more restrictive monetary policies, and in June 1991, the decision was taken to introduce a

[14] For a thorough discussion on the Colombian exchange rate regimes, see Villar and Rincón (2000), as well as Cárdenas (1997). The discussion here draws heavily from Villar and Rincón (2000).

more flexible exchange rate regime. A market for foreign exchange was created, where the exchange rate was freely determined.[15] However, the Banco de la República continued to maintain an official exchange rate set as a crawling peg, and this came to act as a ceiling of the market exchange rate. In practice the new exchange rate regime was a managed floating regime with many similarities to a crawling exchange rate band.

In January 1994, the central bank introduced an official crawling band regime. This was to regain control over monetary variables, after a period of very low real interest rates in combination with very large capital inflows. The exchange rate was allowed to fluctuate around a pre-determined central rate, which initially was to be continuously devalued at an annual rate of 11 per cent. The actual exchange rate could depart with as much as 7 per cent from the central rate. As shown by figure 3.1, the regime, in fact, very much resembled a managed float, since the limits of the band were shifted several times, and since the band was relatively wide. "In this sense, the currency band was not supposed to create obstacles in the process of adjustment of the exchange rate but to guarantee a more orderly and gradual adjustment when such a process was grounded in fundamental macroeconomic changes".[16]

[15] The market traded Exchange Rate Certificates (Certificados de Cambio) which were US dollar denominated interest bearing papers issued by the Banco de la República. See Villar and Rincón (2000), pp 27ff.
[16] Villar and Rincón (2000), p. 30.



The USD/COP exchange rate

Figure 3.1. The Colombian exchange rate band 1994 – 1999

Source: Villar and Rincón (2000), p. 31.

In September 1998 the exchange rate band was shifted, following speculative pressure for a devaluation. The economy experienced difficulties, and the GDP growth rate in 1998 was a mere 1 per cent. In 1999 the economic crisis worsened significantly. In June the band was shifted for the second time in less then a year and a new macroeconomic programme was announced. Colombia was, however, running a large fiscal deficit, and the credibility of the currency band system was rapidly deteriorating. In September 1999 the Government reached an agreement with the International Monetary Fund on a 3-year macroeconomic adjustment programme. As part of this agreement, the exchange rate band was dismantled, and the exchange rate was allowed to float freely.

The floating regime is close to a free float. The Banco de la República has two mechanisms with which it can intervene in the markets.[17] First, the central bank auctions a limited amount of foreign exchange put options every month. This is used for accumulation of international reserves. Second, the bank can intervene in the market to reduce extreme short-term exchange rate volatility. This is done through the auctioning of put or call options on foreign exchange, but is only used if the average exchange rate of a given day deviates more than 4 per cent from its 20-day moving average. The second type of intervention has, in fact, not been used until earlier this year.

Figure 3.2 shows the exchange rate development since 1994, and Figure 3.3 shows the exchange rate variability. It is apparent from Figure 3.2 that the exchange rate has left its path of a long-term stable depreciation rate. Nevertheless, the rate of domestic consumer price inflation has fallen significantly from an average of 16.6 per cent from 1994 to 1998 to a rate of 9.4 per cent in 2001. The short-term variability of the exchange rate, as shown by Figure 3.3, has, on the other hand, not changed significantly. If we calculate the average absolute weekly change for the periods January 1994 to September 1999 and October 1999 to August 2002 we receive values of 0.72 per cent and 0.68 per cent respectively. Short-term variability has, thus, not been influenced by the change in exchange rate regime.

[17] Villar and Rincón (2000), p. 34.

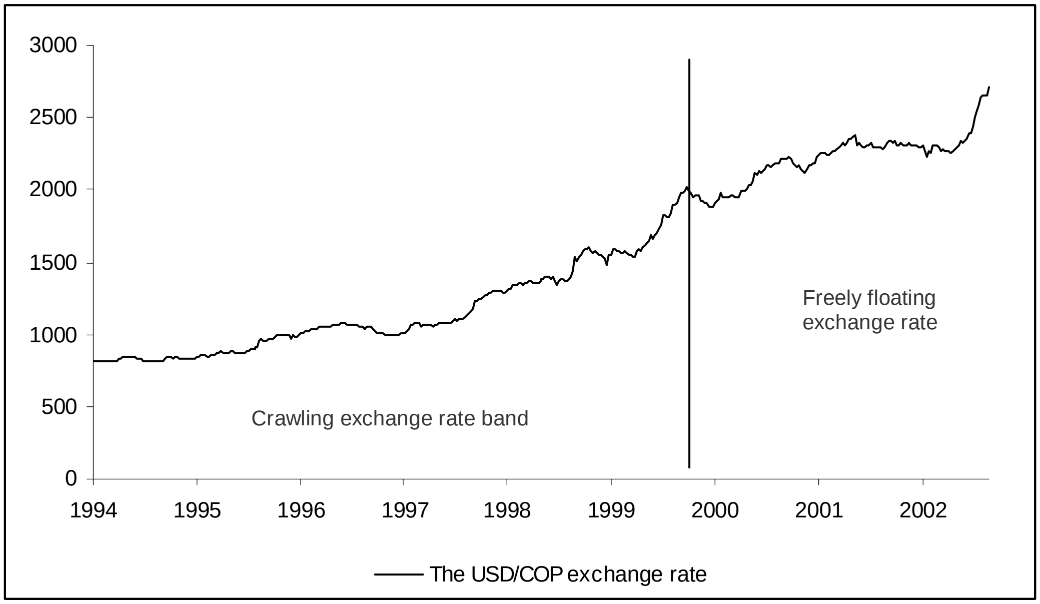


Figure 3.2. The USD/COP exchange rate under the different regimes

Source: Banco de la República.

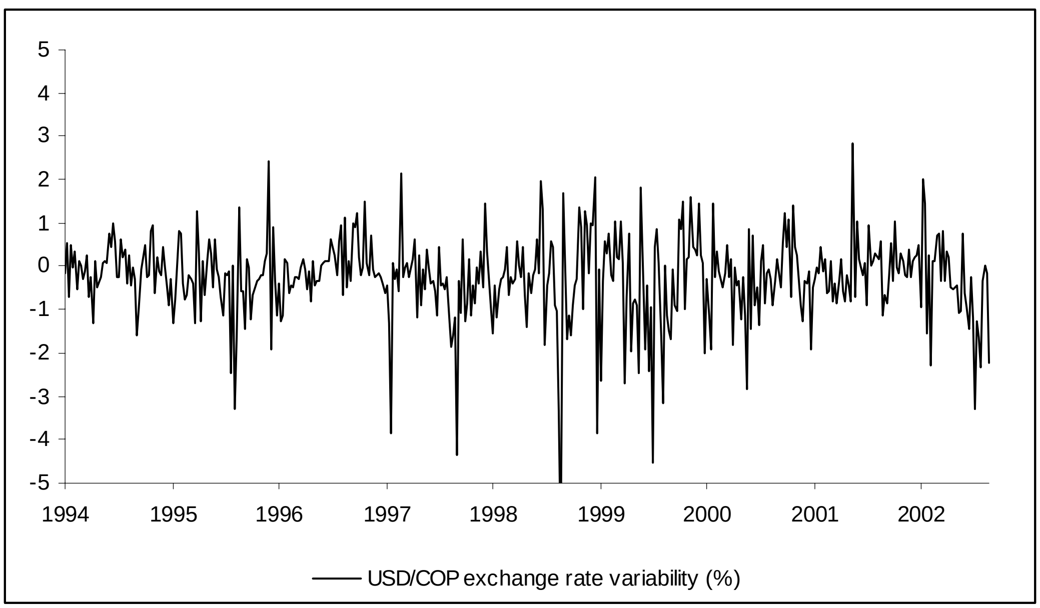


Figure 3.3. Short-term variability of the USD/COP exchange rate, expressed as percentage change from previous week

Source: Banco de la República.

3.2 The Exchange Rate and the Interest Rate Differentials

The UIP hypothesis is tested through estimation of equation (2.7), as discussed in the previous chapter. This states that the change in the exchange rate, $s_{t+T} - s_t$, should equal the interest rate differential, $r_t - r_t^*$. Figure 3.4, 3.5 and 3.6 graph the changes in the exchange rate against the interest rate differentials for 3, 6 and 12 months duration respectively.[18]

From these graphs we make the following observations: First, the 12-month time series seem to follow each other closely after October 1996, as shown by Figure 3.6. However, for the 3-month duration, the time series seem relatively independent of each other, as indicated by Figure 3.4. Second, all the time series look non-stationary, apart from the 3-month change in the exchange rate. Non-stationarity is a condition for cointegration, and we might, therefore, suspect that a valid cointegrating relationship does not exist in the 3-month case. This will be tested in the next chapter.

The pattern in Figure 3.6 is very interesting. It seems that the UIP relationship might hold from October 1996 onwards. However, during this time Colombia went through a considerable macroeconomic transition. Inflation was brought down from double-digit rates to single-digit rates, and the growth rate of the money supply fell significantly. The GDP growth rate slowed, and in 1999 the country experienced its first recession since the 1930s.[19] These variables all have an influence upon the interest rate and the exchange rate.[20] It is quite probable that the macroeconomic transition caused the clear shift in the rate of exchange rate depreciation as well as in the interest rate differentials, illustrated in Figure 3.6.

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[18] The data used is discussed in the following chapter.
[19] A recession is defined as two consecutive quarters of negative growth.
[20] See, for example, seminal papers by Dornbusch (1976), and Hooper and Morton (1982), or the overview in Isard (1995).

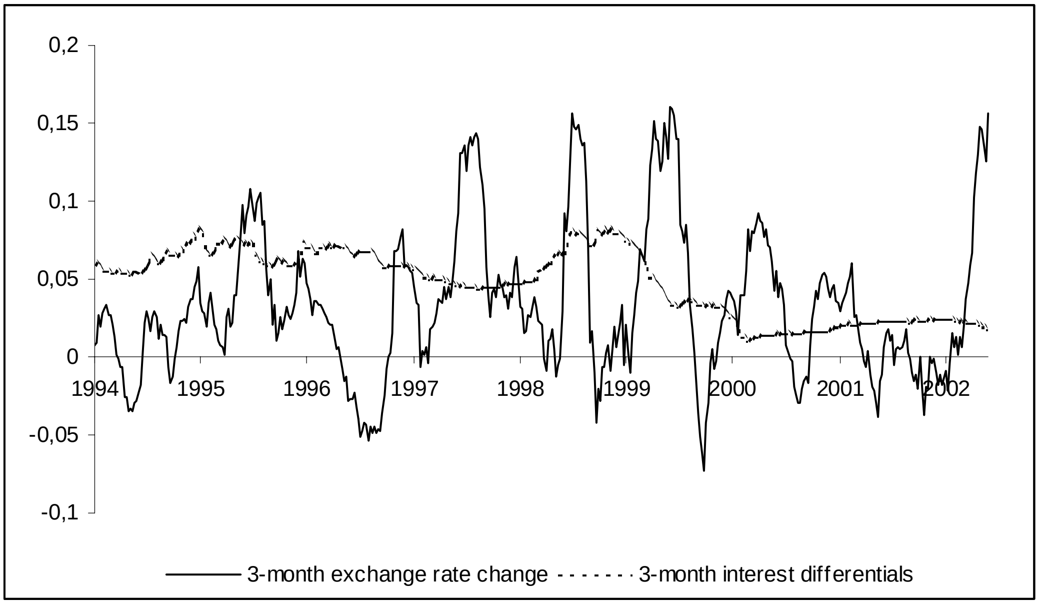

Figure 3.4. The 3-month change in the USD/COP exchange rate together with the differential between US and Colombian 3-month interest rates

Note: Interest differentials are expressed as the difference between the 3-month return on the investments in absolute value, and not as an annualised return.

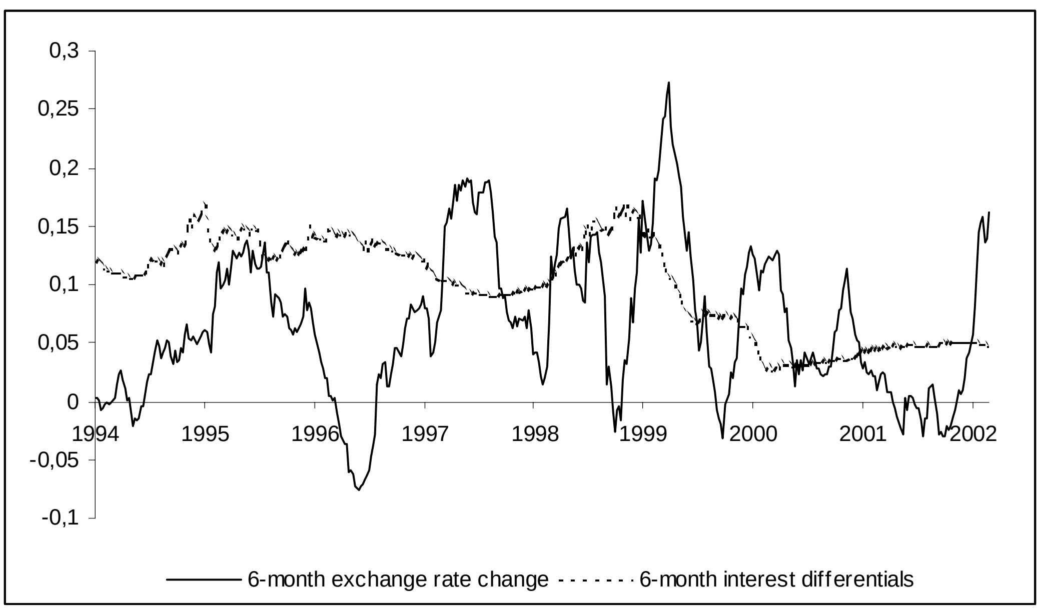

Figure 3.5. The 6-month change in the USD/COP exchange rate together with the 6-month differential between US and Colombian 6-month interest rates

Note: Interest differentials are expressed as the difference between the 6-month return on the investments in absolute value, and not as an annualised return.

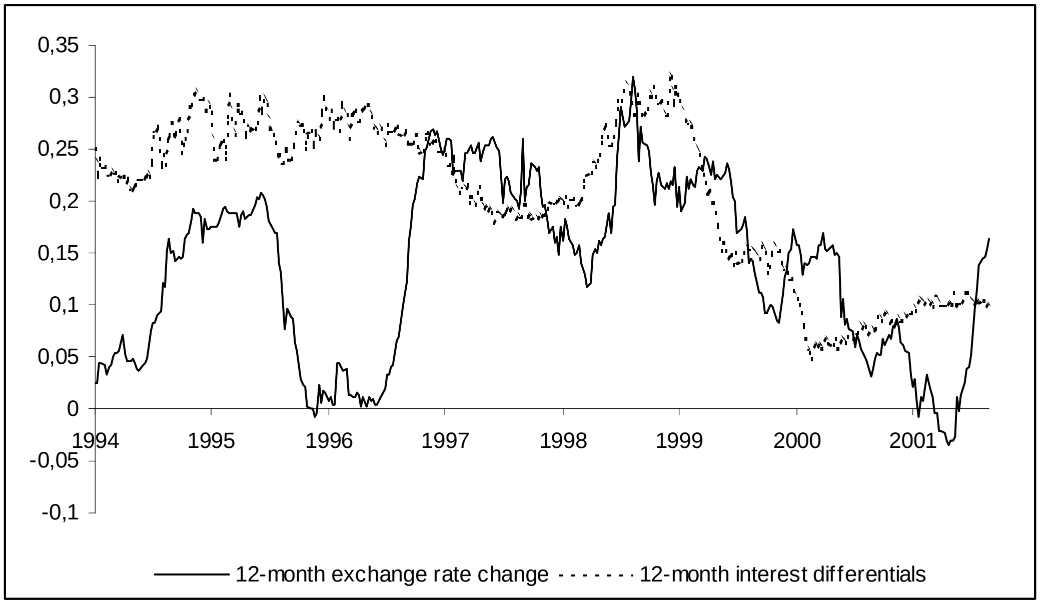

Figure 3.6. The 12-month change in the USD/COP exchange rate together with the 12-month differential between US and Colombian 12-month interest rates

4 Empirical Analysis

This chapter tests the UIP hypothesis for the USD/COP exchange rate. In the first section of the chapter the time-series data used is defined and discussed. In the second section the regression analysis is performed and the results are analysed. Some conclusions of the research are drawn in the next chapter.

4.1 The Data Set

To test the UIP hypothesis for the USD/COP exchange rate we use weekly data from January 1994, when the crawling-band regime was introduced, up until August 2002.[21] As discussed in the previous chapter, Colombia had two different exchange rate regimes during this period. Between January 1994 and September 1999, exchange rate movements were restricted by a crawling band. In September 1999, the crawling band was dismantled and the exchange rate was allowed to float freely, which has been the prevalent exchange rate regime since then.

The UIP hypothesis is tested using 3, 6 and 12-month interest rates. For the United States we are using interest rates of US Treasuries with constant maturities. For Colombia we are using certificates of deposits of the banking system to define interest rates. Interest rates used are those for the DTF, CDT 180 and CDT 360, which are 90-day, 180-day and 360-day certificates of deposits respectively. The source of the US data is the Federal Reserve, and that of the Colombian data is Banco de la República.

[21] The source used is Banco de la República.

As Figure 4.1 shows, the difference between the 3, 6 and 12-month Colombian interest rates are relatively small, while the difference between Colombian and US interest rates are significant.

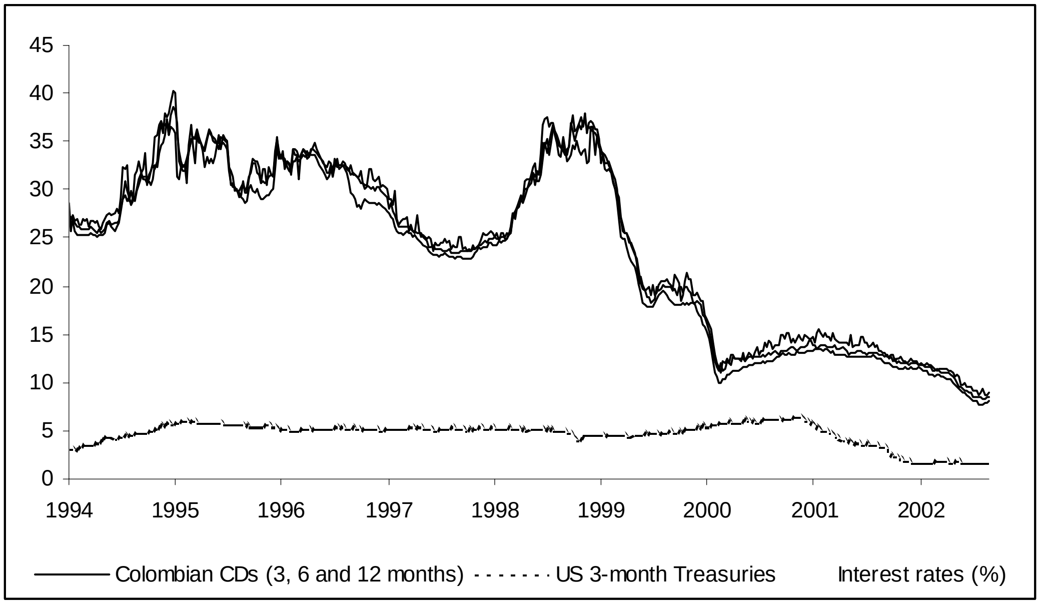

Figure 4.1. Interest rates in Colombia and in the United States

Note: The Colombian interest rates have followed each other closely. The yield curve has generally been normal, implying that the 12-month rate in the graph is the highest and the 3-month rate the lowest. During the interest rate peaks, the yield curve was, however, inverted.

Source: Banco de la República, and the Federal Reserve.

Table 4.1. Ratings of Colombian banks as well as of Colombian and US sovereigns

Entity Rated	Moody's Rating (long-term deposits or government bonds)
Banks (Colombia)	
BBV Banco Ganadero S.A.	Ba3
Bancafe S.A.	Ba3
Banco Popular S.a.	Ba3
Banco de Bogota	Ba3
Bancolombia S.A.	Ba3
Sovereigns	
Colombia (domestic/foreign currency)	Baa2/Ba2
United States	Aaa

Note: The terminology of Moody's differs from that of the other main rating agencies. The Ba3 rating corresponds to the BB-, Baa2 to BBB, Ba2 to BB and Aaa to AAA.
Source: Moody's Investors Service.

The UIP hypothesis normally assumes investments to be riskless. However, in Colombia, investments carry significant risk, and the choice of interest rates, therefore, needs some attention. The main Colombian banks are rated BB- (or Ba3 using the Moody's terminology), as shown in Table 4.1, and we will therefore assume that certificates of deposits of the Colombian banking system will carry a similar risk. This is far below the AAA rating of US sovereign debt, and also far below the AA rating of US bank debt.[22] However, the ratings of Colombian banks are not too far away from the BBB rating of Colombian sovereign debt denominated in the domestic currency.[23] We will, therefore, accept it as a good proxy for Colombian interest rates.

The interest differential between Colombian and US interest rates, will partly be related to the expected change in the exchange rate, according to the UIP hypothesis, and partly to the different risks of default of the different investments. To capture this risk, we will use data on the spread between the 10-year US dollar denominated Colombian global bonds and the 10-year US Treasuries.[24] As shown in Table 4.1, the former are BB rated, which is a rating similar to that of the Colombian banking system, while the latter are AAA rated.

The sovereign spread is illustrated in Figure 4.2. In 1997 the spread was below 200 basis points, and it rose only marginally during the Asian crisis in the second half of 1997. However Colombia was, together with most Latin American countries, severely affected by the Russian default of August 1998, when the spread rose to well over 1000 basis points. Since then the spread has remained volatile and at much higher

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[22] The main US banks are rated AA or AA+ according to Moody's and Standard & Poor's. It would, thus, make little difference if certificates of deposits of the US banking system were used instead of US Treasuries as a base for US interest rates.
[23] Even if the difference is four rating notches, it should not significantly influence the results of the study undertaken here.
[24] Ideally we should have used the spread of investments of 3, 6 and 12-month maturity, but data on such spreads do not exist. The spread data used is, furthermore, a measure of the default risk of Colombian sovereign US dollar denominated securities, while we are using peso denominated certificates of deposits of the banking system as the source of domestic interest rates. We can, however, use the spread of the former as an indicator of the default risk of the latter, if we assume that

levels than experienced in 1997. Recent fiscal problems have induced another surge in the spread, to levels similar to that during the Russian crisis.

the default risks of these two groups of securities are highly correlated, which is not an unreasonable assumption. The source for the data is Banco de la República.

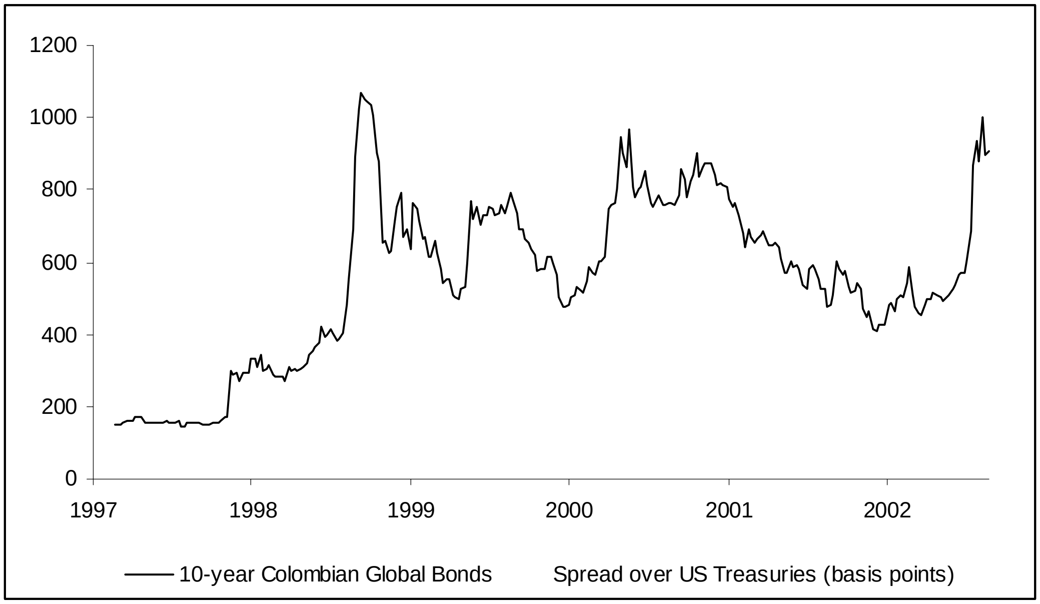

Figure 4.2. The sovereign spread between Colombian and US government securities

Source: Banco de la República.

To estimate equation (2.7) we need to construct time series for the change in the exchange rate, $s_{t+T} - s_t$,[25] as well as for the interest differential, $r_t - r_t^*$. The change in the exchange rate is computed for the duration of time T of 3, 6 and 12 months, which with weekly data correspond to 13, 26 and 52 weeks respectively. The interest rate differentials are computed as the differences between the interest rates on Colombian certificates of deposits and on US Treasuries for the maturities of 3, 6 and 12 months. Note that the interest rate differentials should be expressed as quarterly, semi-annual and annual returns respectively for the 3, 6 and 12-month investments, to be comparable to the percentage change in the exchange rate during these periods. Consequently, the interest differential for the 3-month rate should be divided by four to be comparable with the 3-month change in the exchange rate. The interest differential for the 6-month rate should be divided by two.

[25] Note that exchange rate data is logarithmic.

4.2 Regression Analysis and Results

To test the UIP hypothesis we will investigate the existence of a valid cointegrating vector between the changes in the exchange rate and the interest rate differentials. We are doing this using the Engle-Granger methodology,[26] testing for long-term equilibrium relationships.[27]

We start by testing the time series for unit roots. These tests are summarised in Table 4.1. All the time series are integrated of order one, I(1), apart from the 3-month change in the exchange rate, which is stationary. This is in line with our earlier conclusion from studying the graphs of the time series.

Table 4.1. Unit root tests for the time series

Variable	Dickey-Fuller Statistic	Differentiated Variable	Dickey-Fuller Statistic
3-month time series			
$(s_{t+3} - s_t)$	ADF(32) = -3.31	$\Delta(s_{t+3} - s_t)$	ADF(32) = -4.42
$(r_t - r_t^*)$	ADF(1) = -0.75	$\Delta(r_t - r_t^*)$	ADF(1) = -10.95
6-month time series			
$(s_{t+6} - s_t)$	ADF(30) = -2.24	$\Delta(s_{t+6} - s_t)$	ADF(34) = -3.35
$(r_t - r_t^*)$	ADF(8) = -0.80	$\Delta(r_t - r_t^*)$	ADF(8) = -5.47
12-month time series			
$(s_{t+12} - s_t)$	ADF(24) = -2.37	$\Delta(s_{t+12} - s_t)$	ADF(32) = -3.20
$(r_t - r_t^*)$	ADF(2) = -0.54	$\Delta(r_t - r_t^*)$	ADF(2) = -13.48
Spread			
φ_t	ADF(1) = -1.97	$\Delta\varphi_t$	ADF(1) = -9.62

Note: The Augmented Dickey-Fuller test is used to test for unit roots. The value in parentheses is the order of the lag used. The null hypothesis in each case is that the variable is integrated of order one and, thereby, non-stationary. The 5 per cent rejection region for non-stationarity for the Augmented Dickey-Fuller statistic is ADF < -2.89, according to Fuller (1976).

[26] See Engle and Granger (1987), or for an overview, Enders (1995).

[27] We are here not interested in the short-term dynamics, and, therefore, we do not need to estimate the error-correction model.

To investigate the existence of a cointegrating vector, we estimate equation (2.7) as well as equation (2.8) using OLS methodology. These equations are restated here, for simplicity.

$$(s_{t+T} - s_t) = a + b\,(r_t - r_t^*) + \varepsilon_t \tag{4.1}$$

$$(s_{t+T} - s_t) = a + b\,(r_t - r_t^*) - c\,\varphi_t + \varepsilon_t \tag{4.2}$$

For UIP to be valid, a should approximately equal zero, b should approximately equal one, and c should approximately equal minus one. Nevertheless, equation (4.1) normally assumes that the domestic and foreign investments are of similar risk. This is, however, not the case for Colombian and US investments, where the former carries a significantly higher default risk than the latter. In this case, the parameter estimate of a should be negative when equation (4.1) is estimated, as discussed in section 2.1.

If we estimate equation (4.1) for the 3, 6 and 12-month time series, we receive the results presented in Table 4.2.

For the 3-month case, the parameter estimate of a is positive and significantly different from zero, while the parameter estimate of b is of the wrong sign and not significantly different from zero. However, even if this contradicts the UIP hypothesis, it is in line with other empirical tests, that usually get a negative parameter estimate for b for shorter time durations. More serious is that the Augmented Dickey-Fuller statistic indicates that the error terms are non-stationary,[28] and we can, therefore, accept the null-hypothesis that a valid cointegrating vector does not exist. This result is hardly surprising, since the changes in the exchange rate are stationary, while the interest differentials are I(1), as shown earlier in Table 4.1.

[28] This also implies that the parameter estimates of the OLS regression are not valid.

For the 6 and 12-month case, the parameter estimate of *a* is again positive and significantly different from zero. However, now the parameter estimate of *b* is of the right sign, even if it is not significantly different from zero in the 6-month case. In both cases the existence of a valid cointegrating vector can be rejected, since the error terms are non-stationary, as shown by the augmented Dickey-Fuller statistics.

Table 4.2. Regressions for the time period January 1994 to August 2002

3-month time series

$$(s_{t+3} - s_t) = \underset{(6.38)}{0.036} - \underset{(-0.45)}{0.049}\,(r_t - r_t^*)$$

$R^2 = 0.00$; Adjusted $R^2 = 0.00$; ADF(26) = -2.66

6-month time series

$$(s_{t+6} - s_t) = \underset{(6.24)}{0.051} + \underset{(1.84)}{0.141}\,(r_t - r_t^*)$$

$R^2 = 0.01$; Adjusted $R^2 = 0.01$; ADF(36) = -3.05

12-month time series

$$(s_{t+12} - s_t) = \underset{(6.29)}{0.074} + \underset{(5.53)}{0.300}\,(r_t - r_t^*)$$

$R^2 = 0.01$; Adjusted $R^2 = 0.01$; ADF(18) = -2.86

Note 1: Interest differentials are expressed as the difference between the return on the domestic and the foreign investments in absolute value, and not as an annualised return (apart from the 12-month case, when these are the same).

Note 2: The t-ratios are given in parentheses below the parameter estimates. The Augmented Dickey-Fuller statistic tests the null hypothesis that the residuals are integrated of order one and, thereby, non-stationary. The 5 per cent rejection region for non-stationarity for the Augmented Dickey-Fuller statistic within the Engle-Granger framework is ADF < -3.17. See Enders (1995), p. 383. The value in parentheses shows the order of the lag used for the augmented Dickey-Fuller test.

In the previous chapter, we concluded that even if a cointegrating relationship does not exist for the whole period, it seems to exist after October 1996, as indicated by Figure 3.5. We, therefore, rerun the regressions using data for the period October 1996 to August 2002. The results are presented in Table 4.3.

Table 4.3. Regressions for the time period October 1996 to August 2002

3-month time series

$$(s_{t+3} - s_t) = \underset{(4.52)}{0.029} + \underset{(2.21)}{0.319}\,(r_t - r_t^*)$$

$R^2 = 0.02$; Adjusted $R^2 = 0.02$; ADF(17) = -4.47

6-month time series

$$(s_{t+6} - s_t) = \underset{(3.62)}{0.031} + \underset{(6.17)}{0.583}\,(r_t - r_t^*)$$

$R^2 = 0.12$; Adjusted $R^2 = 0.12$; ADF(18) = -5.34

12-month time series

$$(s_{t+12} - s_t) = \underset{(2.45)}{0.021} + \underset{(16.96)}{0.801}\,(r_t - r_t^*)$$

$R^2 = 0.53$; Adjusted $R^2 = 0.53$; ADF(12) = -3.27

These results look very interesting. First, the error terms in all three cases are stationary, indicating the existence of a valid cointegrating relationship. The parameter estimates of b are of the right sign and significantly different from zero in all the cases, and in the 6 and 12-month cases they are, furthermore, closer to unity than to zero.[29]

In all three cases the parameter estimate of *a* is positive and significantly different from zero. This might at first be rather puzzling, since *a* according to theory should be negative. Colombian investments carry a significant default risk relative to US investments and should, therefore, yield an interest rate premium even if the currency is not expected to depreciate, and, hence, *a* should be negative. However, the parameter estimate of *b* is *not* equal to one, and the depreciation of the exchange rate has, therefore, according to these estimations, not been fully explained by the interest differential. If *b* in the 12-month case is restricted to one, the parameter estimate of *a* will be minus 0.012, which indeed is negative. It is also significantly different from zero, with a t-statistic of minus 3.41. In the 6-month and 3-month case the parameter estimate of *a* would be minus 0.002 and 0.002 respectively. In none of these cases the parameter estimate is significantly different from zero. T-statistics are minus 0.56 and 0.55 respectively.

We can, therefore, conclude that the results, particularly for the 12-month case, are supportive of the existence of a UIP relationship. For the 12-month case, the explanatory value (the adjusted R squared) is 0.53, which, furthermore, is very good.

However, we should at this point emphasise that, even if the frequency of the data used is relatively high (weekly data), the time period from October 1996 to August 2002 is actually too short to draw any definite conclusion of a stable long-run relationship.[30] The shortness of the time series used is, consequently, a weakness in our results.

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[29] It is, however, not possible to determine whether the parameter estimates are significantly different from one in the OLS framework used. According to the t-tests the parameter estimates are significantly different from one. But since the variables used are non-stationary and the error terms, therefore, autocorrelated, the confidence intervals yielded by the t-test are normally underestimated. See Johnston (1991) pp. 310ff. Therefore we cannot draw any definite conclusion.

[30] See Juselius (1994) for a discussion on this subject.

We continue to test the UIP hypothesis, now including the spread, and using equation (4.2). We only have data for the spread from February 1997 and onwards and consequently estimate the equation for the period February 1997 to August 2002. The results are presented in Table 4.4.

Table 4.4. Regressions for the time period February 1997 to August 2002 including sovereign spreads

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3-month time series

$$(s_{t+3} - s_t) = \underset{(5.73)}{0.062} + \underset{(1.16)}{0.177}\,(r_t - r_t^*) - \underset{(-3.79)}{2.092}\,\varphi_t$$

$R^2 = 0.02$; Adjusted $R^2 = 0.01$; ADF(17) = -4.77

6-month time series

$$(s_{t+6} - s_t) = \underset{(7.45)}{0.101} + \underset{(4.98)}{0.477}\,(r_t - r_t^*) - \underset{(-6.64)}{2.176}\,\varphi_t$$

$R^2 = 0.26$; Adjusted $R^2 = 0.25$; ADF(18) = -4.41

12-month time series

$$(s_{t+12} - s_t) = \underset{(5.65)}{0.079} + \underset{(13.95)}{0.696}\,(r_t - r_t^*) - \underset{(-5.00)}{0.772}\,\varphi_t$$

$R^2 = 0.55$; Adjusted $R^2 = 0.55$; ADF(8) = -3.81

Note: Both interest differentials and spreads are expressed as the difference between the return on the domestic and foreign investments in absolute value, and not as an annualised return (apart from the 12-month case, when these are the same).

Also these results look very good. In all the regressions the parameter estimate of c is of the right sign and significantly different from zero, and in the 12-month case it is not significantly different from minus one. The error terms are all stationary, indicating the existence of a cointegrating relationship. The parameter estimates of b are all significant and of the right sign. They are however further away from unity then in the previous regression. The parameter estimate of a are also in all the cases higher than in the previous regression.

Colombia abandoned the crawling band in September 1999 in favour of a freely floating exchange rate. Studying Figure 4.5 leads us to conclude that this might have changed the relationship between the changes in the exchange rate and the interest rate differentials. If we rerun the regressions with and without the sovereign spread for the period after September 1999, this yields the results presented in Table 4.5 and 4.6.

Table 4.5. Regressions for the time period October 1999 to August 2002

3-month time series

$$(s_{t+3} - s_t) = \underset{(7.49)}{0.099} - \underset{(-5.81)}{3.605} (r_t - r_t^*)$$

$R^2 = 0.20$; Adjusted $R^2 = 0.19$; ADF(13) = -0.90

6-month time series

$$(s_{t+6} - s_t) = \underset{(5.06)}{0.092} - \underset{(-2.75)}{1.104} (r_t - r_t^*)$$

$R^2 = 0.06$; Adjusted $R^2 = 0.05$; ADF(4) = -1.28

12-month time series

$$(s_{t+12} - s_t) = \underset{(5.45)}{0.124} - \underset{(-2.14)}{0.505} (r_t - r_t^*)$$

$R^2 = 0.04$; Adjusted $R^2 = 0.03$; ADF(2) = -1.26

Table 4.6. Regressions for the time period October 1999 to August 2002 including sovereign spreads

3-month time series

$$(s_{t+3} - s_t) = \underset{(5.77)}{0.156} - \underset{(-6.32)}{4.446} (r_t - r_t^*) - \underset{(-2.40)}{2.504} \varphi_t$$

$R^2 = 0.23$; Adjusted $R^2 = 0.22$; ADF(5) = -2.05

6-month time series

$$(s_{t+6} - s_t) = \underset{(4.24)}{0.166} - \underset{(-3.48)}{1.612} (r_t - r_t^*) - \underset{(-2.11)}{1.576} \varphi_t$$

$R^2 = 0.09$; Adjusted $R^2 = 0.08$; ADF(1) = -0.81

12-month time series

$$(s_{t+12} - s_t) = \underset{(6.11)}{0.283} - \underset{(-3.66)}{0.890} (r_t - r_t^*) - \underset{(-3.86)}{1.791} \varphi_t$$

$R^2 = 0.17$; Adjusted $R^2 = 0.15$; ADF(1) = -1.36

The error terms in all these regressions are non-stationary, which indicates that the cointegrating relationships that existed earlier might have been broken. However, the regression period is too short for this latest test to provide for any definite conclusion. It is also apparent that the parameter estimates of b are significantly different from zero, but of the wrong sign. We, consequently, conclude that the UIP relationship seems to have been significantly weakened, if not broken totally, after the introduction of the floating exchange rate. Due to the shortness of the time period, these results might, however, not be stable, and will need to be confirmed by a future study when longer time series exist.

5 Conclusion

This paper has tested the UIP hypothesis with the USD/COP exchange rate, using weekly data for the period from January 1994, when the crawling-band regime was introduced, up until August 2002. We have used ordinary least squares to test equation (2.7) and (2.8) for Engle-Granger cointegration, which is in line with many earlier studies.

The study has yielded several important results. The UIP hypothesis seem to hold for the period October 1996 to August 2002, even if the shortness of the time series used might put the stability of the results into question. The UIP relationship, furthermore, seem to have been significantly weakened towards the end of the period, if not broken completely. The results are summarised in Table 5.1. We have shown earlier, that for all the regressions presented in Table 5.1, the error terms are stationary, and a valid cointegration relationship, consequently, exist. Estimates of the slope parameter b are all of the right sign, even if not significantly different from zero in the 3-month case with spread (the 3-month case without spread is, furthermore, a border case). For the 12-month cases, as well as for the 6-month case without spread, the slope parameter is closer to unity than to zero. Estimates of the spread parameter c are all significant and of the right sign.

These test results are among the most favourable of any study of the UIP hypothesis, which might be somewhat surprising. They differ starkly from the almost unanimous rejection of the hypothesis shown by earlier papers.

It is, furthermore, apparent from the results in Table 5.1, that the validity of the UIP relationship increases with the term of the investment, which is in line with earlier studies. While the results for the 3-month duration are relatively weak, the UIP hypothesis receive significant support from the 12-month results.

Table 5.1. Summary of regression results for October 1999 to August 2002

Time to maturity	Estimate of *a* value	t-stat	Estimate of *b* value	t-stat	Estimate of *c* value	t-stat
Without spread – equation (2.7)						
3 months	0.029	(4.52)	0.319	(2.21)	N/A	
6 months	0.031	(6.32)	0.583	(6.17)	N/A	
12 months	0.021	(2.45)	0.801	(16.96)	N/A	
With spread – equation (2.8)						
3 months	0.062	(5.73)	0.117	(1.16)	-2.092	(-3.79)
6 months	0.101	(7.45)	0.477	(4.98)	-2.176	(-6.64)
12 months	0.079	(5.65)	0.696	(13.95)	-0.772	(-5.00)

Note 1: The lower half of the table (estimation of equation (2.8)) uses data from the period February 1997 to August 2002, since spread data was not available for earlier dates.

Note 2: In all the cases the error terms have been shown to be stationary, and a valid cointegrating vector, therefore, exists.

Colombia introduced its crawling-band exchange-rate regime in January 1994. However, this study indicates that a valid UIP relationship did not exist until mid-1996. This relationship seem to have been considerably weakened towards the end of the period, and we might suspect that this had to do with the change of exchange rate regime in 1999, when the exchange rate band was abolished in favour of a free float. This would be in line with research presented by Lewis (1988, 1989), who argue that market participants do not have complete information and are engaging in a process of rational learning. When an exchange rate regime has been changed, it will take some time for market participants to learn to correctly form their expectations of future exchange rates under the new regime. If this is the case, the validity of the UIP relationship should again increase when the floating rate regime has been in place for some time.

There is, however, also a possibility that the significant UIP relationship found from October 1996 onwards could turn out to be a temporary one-off occurrence in

Colombian exchange rate history. As discussed earlier, Colombia underwent a considerable macroeconomic transition at this time, which should have had a significant impact on both the rate of depreciation of the exchange rate and on the interest rate differentials. This transition might, indeed, have been the main reason why the UIP relationship holds during this period.

Some recent studies have, nevertheless, indicated that UIP might actually work better in emerging economies than in developed economies, on which it has normally been tested. Bansal and Dahlquist (2000) concludes that the negative correlation between the expected currency depreciation and interest rate differentials is mainly confined to developed economies. Flood and Rose (2001) finds that UIP works better for countries in crisis, but does, however, not find any difference between rich and poor countries in general.

Whether the strong support for UIP in Colombia since October 1996 is a temporary occurrence or whether the UIP relationship will actually hold also in the future can only be determined by future studies.

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