

The Exchange Rate and The Term Structure of Interest Rates in Mexico

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Abstract

Mexico adopted a floating exchange rate regime in December 1994. The Bank of Mexico's monetary policy gives attention to maintain "orderly conditions in foreign exchange markets." The Bank of Mexico relies primarily on the control of the overnight interest rate in conducting its monetary policy. The question arises whether this extremely short-term interest rate is the relevant instrument to achieve exchange rate objectives. Theory suggests that the relationship between the exchange rate and the term structure of interest rates can be complicated and counterintuitive when investors are risk averse. In this paper, we pursue an empirical investigation on the effect of the term structure of interest rates on the exchange rate for Mexico. This information could be useful to understand and manage the operation of the Mexican floating exchange rate regime.

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I. INTRODUCTION

The summer of 1997 saw a flurry of developing countries moving toward floating exchange rate regimes. The most notable episodes were the currency collapses in Southeast Asia, which forced monetary authorities in these countries to abandon long-standing currency pegs. More recently, several South American countries joined the fray, allowing their currencies to float in 1999.

Floating currencies in the context of capital mobility restore, of course, the possibility of having a true monetary policy. Perhaps one of the most widespread consensus across academic and policy-making circles that is likely to guide monetary policy of the recent floaters, is on the solid relationship between interest rates and the exchange rate. Conventional wisdom, based on the Mundell-Fleming tradition and Dornbusch's (1976) contribution, but also in more recent papers with risk neutral investors, is that increases in any interest rate would appreciate the domestic currency. A corollary of this proposition is that the term structure of interest rates is irrelevant for the determination of the exchange rate. As this paper will argue, reliance on this conventional wisdom is unwarranted.

For a start, the uncovered interest parity implied by risk neutrality is typically rejected by the short-horizon data (see, e.g., Isard (1995) and Engel (1997) for recent surveys). As Engel (1997) emphasizes, one form of the rejection found in many recent papers is that regressions of the future depreciation on the current forward premium (which is equal to the short-term interest differential under the covered interest parity) yield negative estimates of the slope coefficient. This is called the forward premium anomaly (see Backus, Foresi, and Telmer (1998) for a recent discussion). It is difficult to find an economic explanation of the forward premium anomaly. For example, Mark and Wu (1998) shows that the standard consumption-based capital asset pricing model is unable to explain the forward premium anomaly.

Meredith and Chinn (1998) recently found another stylized fact: the forward premium anomaly does not exist in long-horizon data. When they regress long-horizon exchange rates onto long-term interest rate differentials for the G-7 countries, all the coefficients on interest rate differentials are of the correct sign. There is much indirect empirical evidence that the uncovered interest parity holds better in the long-run than in the short-run. Cointegration results in Meese and Rogoff (1988) and Edison and Pauls (1993) suggest that long-term interest rate differentials are more important than short-term interest rate differentials in the long-run determination of the real exchange rates. Baxter (1995) finds that the correlation between real exchange rates and real interest differentials is generally positive at trend and business-cycle frequencies, and is stronger for long-term interest differentials. Boughton (1988) finds for the United States, Germany and Japan that the empirical performance of asset-market models of exchange rates can be improved by including information about the term structure of interest rate differentials. Eichenbaum and Evans (1995) and Mark (1995) have found that some implications of uncovered interest parity do not hold in the short-run,

but do hold in the long-run. Byeon and Ogaki (1999) use cointegration techniques and find that the short-term interest rate differential and the long-term interest rate differential have the opposite effects on the exchange rate.

An economic model can be consistent with these short-run and long-run stylized facts about the forward premium anomaly if short-term domestic and foreign bonds are complements, but long-term domestic and foreign bonds are substitutes in the model. Ogaki (1999) constructs such a model by assuming that the investors are risk averse and have a short investment horizon.¹ In this model with domestic one-period and two-period bonds, as well as foreign bonds, domestic one-period bonds and foreign bonds are complements even when investors are close to risk neutral. The model predicts a complicated relationship between the exchange rate and the term structure of interest rates.

The intuition behind this result arises from Ogaki's (1990) concept of indirect complementarity, which relates to the risk structure of interest rates under the assumption of risk aversion. If domestic one-period interest rates unexpectedly rise, investors with a one-period investment horizon holding domestic two-period bonds suffer a capital loss. If the domestic currency appreciates as a result of this increase, then holders of foreign bonds also suffer a capital loss. Therefore, as long as an increase in one-period interest rates is associated with an appreciation of the domestic currency, risk averse agents have an incentive to avoid holding both domestic two-period bonds and foreign bonds. Hence these two assets are likely to be strong substitutes. Since domestic one-period and two-period bonds are also likely to be strong substitutes, and since a substitute of a substitute is an indirect complement, domestic one-period bonds and foreign bonds must be strong indirect complements.

When the indirect complementarity dominates the direct substitutability between domestic one-period bonds and foreign bonds, the relationship between exchange rates and interest rate differentials can be counterintuitive and at odds with the conventional wisdom. Suppose that the one-period real interest rate in the domestic country rises relative to foreign real one-period interest rates. If the domestic two-period real interest rate also rises relative to their foreign counterparts, then the domestic currency appreciates as predicted by standard models. However, if the domestic two-period real interest rate does not rise, then the domestic currency depreciates. This is because economic agents substitute away from domestic two-period bonds, and this increases the demand for foreign bonds. In order to achieve equilibrium for foreign bonds, the domestic currency depreciates now, which decreases the demand for foreign bonds because the depreciation gives rise to an expectation that the domestic currency will appreciate in the future. The intuition for the one-period and

¹ McCallum (1994) provides an explanation for evidence against uncovered interest parity based on policy reactions. Our explanation is complementary to his because McCallum's model requires an error term which allows the exchange rate to deviate from the level implied by the exact uncovered interest rate parity relationship. Alvarez, Atkeson, and Kehoe (1999) constructs a model of segmented asset markets which can be consistent with the forward premium anomaly.

two-period interest rates in this model can be applied to any short-term and longer-term interest rates as long as the investment horizon of the investors is consistent with the term to maturity of the short-term bond.

In this paper, we pursue an empirical investigation on the effect of the term structure of interest rates on the exchange rate for Mexico. A floating exchange rate regime was adopted in Mexico following the collapse of the exchange rate band in December 1994. However, as in many other developing countries, Mexico's historical experience with the operation of floating regimes has been brief. The Bank of Mexico has adopted a monetary policy framework that consists of targeting base money. The ultimate objective is to control inflation, but attention is also given to maintain "orderly conditions in foreign exchange markets." In achieving its objectives, the Bank of Mexico relies primarily on the control of day to day liquidity, particularly through the use of targets for the overdraft balances that commercial banks can have with the Bank of Mexico. When monetary authorities attempt to restore order to the foreign exchange market, it is usually done through the management of the banking system's liquidity, with a view to affect overnight or "funding" interest rates. The question arises whether this extremely short-term interest rate is the relevant instrument to achieve foreign exchange market objectives. It is possible, as suggested above, that in this short-term interest rate may have counterintuitive effects on the exchange rate. If this were the case, the Bank of Mexico could, in principle, use traditional open market operations to affect longer-term interest rates for this purpose.

An illustrative example of the importance of the relation between the exchange rate and the term structure of interest rates comes from events in Mexico in the aftermath of the Russian crisis of 1998. What used to be a normal-looking yield curve before the Russian crisis, became highly inverted at the onset of the crisis just as the Mexican peso became under strong selling pressure and was depreciating quite rapidly. Thus, at least during this episode, the behavior of the term structure of interest rates did not conform to the conventional wisdom. Moreover, the consequence of these big shifts in the yield curve was a large capital loss for holders of short-term instruments, especially Mexican banks.

In spite of the importance of the subject, there has been virtually no research on the relationship between the term structure of interest rates and the exchange rate in the case of Mexico. Regarding other related studies, Tapia (1990) finds a positive relationship between the nominal exchange rate and the three-month interest rate differential for the period 1978-87. Werner (1996) finds evidence in favor of the portfolio balance model with his test for the relevance of relative asset supplies in explaining the currency risk premium for Mexico.

Unfortunately long-term bond markets are not well developed in Mexico. That forced us to focus on the term structure of money market interest rates in this paper. We investigate how one- and three-month Cetes interest rates affect the exchange rate in Mexico. Assuming that the investment horizon is one-month, we expect that the three-month interest differential to have the standard effect on the exchange rate: a rise in the three-month Cetes rate tends to

appreciate the Mexican peso. If the indirect complementarity dominates the direct complementarity, then the one-month interest rate differential will have the opposite effect on the exchange rate after controlling for the three-month interest rate differential. Our empirical results are consistent with this view, notwithstanding the focus on such a short horizon. This information could be useful to understand and manage the operation of the Mexican floating exchange rate regime, as well as provide some guidance to other developing countries.

II. A MODEL OF EXCHANGE RATE AND THE TERM STRUCTURE OF INTEREST RATES

In this section, we present a partial equilibrium model of exchange rate determination based on the work of Ogaki (1999), which motivates our empirical investigation. The model is a three-asset extension of the two-asset models of Driskill and McCafferty (1980) and Fukao (1983). The demand of foreign bonds is derived endogenously in a rational expectations equilibrium for the covariance of the (one-period) interest rate and the exchange rate. For simplicity, the price level is assumed to be constant. Alternatively, all variables can be considered to be measured in real terms. Investors are assumed to be identical and live for two periods, and the same number of investors is assumed to be born every period. Let $B_{1,t}$, $B_{2,t}$, and $B_{F,t}$ denote domestic one-period bonds, domestic two-period bonds, and foreign bonds, respectively. The one-period and two-period bonds are discount bonds which will pay one unit of the domestic currency after two periods and one period, respectively. Let q_t be the price of two-period bonds during period t , r_t be the domestic one-period bonds interest rate and R_t be the domestic interest rate on two-period bonds. Then the price on two-period bonds is:

$$q_t = \frac{1}{(1 + R_t)^2}. \quad (1)$$

and the rate of return of holding two-period bonds for one period is:

$$r_{2,t} = \frac{1}{q_t} \left(\frac{1}{1 + r_{t+1}} - q_t \right) \quad (2)$$

From (1) and (2), we obtain

$$r_{2,t} = \frac{(1 + R_t)^2}{(1 + r_{t+1})} - 1 \cong 2 R_t - r_{t+1}. \quad (3)$$

Define the risk premium for two-period bonds, $\theta_{2,t}$, to be the difference between the expected rate of return for two-period bonds and that for one-period bonds:

$$\mathbf{q}_{2,t} = E_t(r_{2,t}) - r_t = 2 \left[R_t - \frac{1}{2} \{r_t + E_t(r_{t+1})\} \right], \quad (4)$$

where E_t is the expectation operator conditional on the information set in period t , Ω_t . We assume that Ω_t includes the current and past values of r_t , R_t , r_t^* , and s_t , where r_t^* is the foreign one-period interest rate and s_t is the natural log of the exchange rate expressed in terms of the domestic currency.

Then the one-period holding rate of return on foreign bonds in terms of the domestic currency is

$$r_{F,t} = r_t^* + s_{t+1} - s_t. \quad (5)$$

Let the risk premium for the foreign bonds, $\theta_{F,t}$, be defined as the difference between the expected return on foreign bonds and the domestic one-period interest rate:

$$\mathbf{q}_{F,t} = r_t^* + E_t(s_{t+1}) - s_t - r_t. \quad (6)$$

The representative investor in period t maximizes his expected utility defined over next period's wealth

$$E_t(u_t(w_{t+1})) \quad (7)$$

subject to the budget constraint,

$$B_{1,t}^d + B_{2,t}^d + B_{F,t}^d = w_t \quad (8)$$

where $B_{i,t}^d$ represent the investor's holding of asset i , the superscript d denotes demand, w_t is the domestic currency value of his assets in the beginning of the period t , and w_{t+1} satisfies

$$w_{t+1} = B_{1,t}^d(1 + r_t) + B_{2,t}^d(1 + r_{2,t}) + B_{F,t}^d(1 + r_{F,t}) \quad (9)$$

In this partial equilibrium model, the stochastic processes for the interest rates are exogenously given, and the utility function is parameterized. The equilibrium exchange rate satisfies the condition that $B_{F,t}^d$ is equal to $B_{F,t}^s$ where $B_{F,t}^s$ is the supply of the foreign bonds to the domestic residents. As is customary elsewhere, the supply of foreign bonds is equal to the cumulative current account balance, and follows the dynamic equation:

$$B_{F,t}^s = B_{F,t-1}^s + C_t \quad (10)$$

where C_t is the current account balance in the period t . Neglecting interest received by holders of foreign bonds, C_t is assumed to satisfy

$$C_t = -a + b s_t + u_t, \quad (11)$$

where a and b are positive numbers, and u_t is the trade shock which is assumed to be white noise with variance σ_u^2 .

Suppose that w_{t+1} is normally distributed conditional on Ω_t , and that the utility function is of the CARA class, so that the absolute risk aversion $-u''/u'$ is a positive constant ζ . Then the demand function for foreign bonds and the demand function for two-period bonds are

$$B_{F,t}^d(E_t(r_{F,t}), r_t, E_t(r_{2,t})) = \mathbf{y} \mathbf{q}_{F,t} - \mathbf{y} \mathbf{q}_{2,t} \quad (12)$$

$$B_{2,t}^d(E_t(r_{F,t}), r_t, E_t(r_{2,t})) = \mathbf{y} \frac{\mathbf{s}_s^2}{\mathbf{s}_r} \mathbf{q}_{2,t} - \mathbf{y} \mathbf{f} \mathbf{q}_{F,t}, \quad (13)$$

where

$$\mathbf{y} = \frac{1}{\mathbf{z} \mathbf{s}_s^2 (1 - \mathbf{r}_{sr}^2)} \quad (14)$$

$$\mathbf{f} = -\frac{\mathbf{s}_{sr}}{\mathbf{s}_r^2} \quad (15)$$

$$\mathbf{s}_s^2 = E_t \left\{ \left[s_{t+1} - E_t(s_{t+1}) \right]^2 \right\} \quad (16)$$

$$\mathbf{s}_r^2 = E_t \left\{ \left[r_{t+1} - E_t(r_{t+1}) \right]^2 \right\} \quad (17)$$

$$\mathbf{s}_{sr} = E_t \left\{ \left[s_{t+1} - E_t(s_{t+1}) \right] \left[r_{t+1} - E_t(r_{t+1}) \right] \right\} \quad (18)$$

$$\mathbf{r}_{sr} = \frac{\mathbf{s}_{sr}}{\sqrt{\mathbf{s}_s^2} \sqrt{\mathbf{s}_r^2}}. \quad (19)$$

As in (12), the demand function for foreign bonds depends on σ_{sr} , the covariance conditional on Ω_t between the exchange rate and the one-period interest rate, and σ_s^2 , the conditional variance of the exchange rate. At the same time, the stochastic process of the exchange rate and σ_{sr} depend on the demand function for foreign bonds. Therefore, it is necessary to solve for a rational expectations equilibrium in which the values of σ_{sr} and σ_s^2 that the investors expect are consistent with the stochastic process of the exchange rate implied by the demand function consistent with these values of σ_{sr} and σ_s^2 . Given these second moments, (12) and (13) give the demand functions for foreign bonds and two-period bonds as functions of expected returns. Hence it is possible to define substitution and income effects for changes in expected returns as in Blanchard and Plantés (1977) and Royama and Hamada (1967).

According to (13), the sign of ϕ determines whether foreign bonds and two-period bonds are substitutes or complements. From (15), the sign of ϕ is determined by the sign of the conditional covariance between the exchange rate and the one-period interest rate, σ_{sr} . In the rational expectations equilibrium discussed below σ_{sr} is always negative, so ϕ is positive. Therefore, foreign bonds and two-period bonds are substitutes. An intuitive explanation of this result is as follows. If the one-period interest rate rises unexpectedly in the next period, the price of two-period bonds falls and two-period bond holders suffer an unexpected capital loss. Because the negative σ_{sr} implies that the exchange rate appreciates when the one-period interest rate rises on the average, bond holders will suffer an additional unexpected loss if they hold foreign bonds. Because of this, foreign bonds and two-period bonds are substitutes if σ_{sr} is negative: the demand for foreign bonds declines when the expected return for two-period bonds rises.

The stochastic processes of one-period interest rates are assumed to be as follows:

$$r_t = \mathbf{m} + e_t + e_{t-1} + \mathbf{e}_t \quad (20)$$

$$R_t = \mathbf{m} + e_t + \frac{1}{2} e_{t-1} \quad (21)$$

$$r_t^* = \mathbf{m} \quad (22)$$

where e_t and \mathbf{e}_t are, respectively, an interest rate shock with a duration of two periods and a temporary interest rate shock with a duration of one period, and μ is constant. It is assumed

that e_t and \mathbf{e}_t are white noise with variance σ_e^2 and $\sigma_{\mathbf{e}}^2$, respectively, and that they are independent of each other and of u_t .

The conditional expectation is assumed to coincide with the best linear prediction. It follows that

$$E_t(r_{t+1}) = \mathbf{m} + e_t, \quad (23)$$

and from (4) and (21),

$$\mathbf{q}_{2,t} = -\mathbf{e}_t. \quad (24)$$

As shown in (24), the assumption employed here is that only the interest rate shock with the two period duration is transmitted to the two-period interest rate, so that the risk premium is equal to the temporary interest rate shock.

Define $\eta = \sigma_e^2 / \sigma_{\mathbf{e}}^2$, which may be called the measure of substitution between one-period and two-period bonds. If $\eta = 0$, then the risk premium for two-period bonds will always be zero, implying that one-period and two-period bonds are perfect substitutes. The greater is η , the smaller the degree of the substitution.

Solving the equilibrium conditions for the foreign bond market as a difference equation system of $E_t(s_{t+\tau})$, with respect to τ , provides the unique saddle point solution,²

$$s_t = \bar{s} - \left[\frac{I-I}{b} u_t \right] - [I(1+I)e_t + I e_{t-1}] + [I(\mathbf{f}-I)\mathbf{e}_t] - \left[\frac{I-I}{b} B_{F,t-1} \right], \quad (25)$$

where $\bar{s} = a/b$ is the long-run equilibrium exchange rate which clears the current account, and

$$I = 1 + \frac{b}{2\mathbf{y}} - \frac{b}{2\mathbf{y}} \sqrt{1 + \frac{4\mathbf{y}}{b}}. \quad (26)$$

It is easy to check $0 < \lambda < 1$, $\partial\lambda/\partial\psi > 0$, $\lim_{\psi \rightarrow 0} \lambda = 0$, and $\lim_{\psi \rightarrow \infty} \lambda = 1$.

Equation (25) shows that the investors' expected values of σ_{sr} and σ_s^2 affect the exchange rate dynamics through λ and ϕ . On the other hand, the exchange rate dynamics in (25) imply certain values of σ_{sr} and σ_s^2 , which need to be consistent with the investors' expected values in the rational expectations equilibrium.

Before discussing the solution for the equilibrium, it is worth noting the nature of (25), which will be the basis of our empirical work below. The discrepancy between the actual and the long-run equilibrium exchange rate is explained by the trade shock (in the first term in brackets), the interest rate shock that is transmitted to the two-period interest rate (the second term in brackets), the interest rate shock that is not transmitted to the two-period interest rate (the third term in brackets), and the cumulative current account balance (in the fourth term in

² For derivation details, refer to Ogaki (1999).

brackets). The trade shock, which tends to give rise to current account surplus, makes the domestic currency appreciate. Prolonged increases in the one-period interest rate make the domestic currency appreciate, though the contemporaneous shock e_t has a greater effect than the previous period's shock e_{t-1} . As the cumulative current account balance becomes greater, the domestic currency's appreciation increases: for investors to have incentives to hold more foreign bonds, the domestic currency must appreciate at present, so that investors can anticipate that it will depreciate in the future.

All of these effects are consistent with the conventional wisdom and go in the standard directions. The shock that is not transmitted to the two-period interest rate, e_t , however, has a perverse effect if the relative magnitude of the indirect complements relationship, ϕ , is greater than 1. This is because the indirect complementarity of one-period and foreign bonds will exceed the direct substitutability if $\phi > 1$.

The term ϕ may be obtained by solving for the rational expectation of covariance. By equating the conditional covariance of the one-period interest rate and the exchange rate implied by (25) and σ_{sr} used to define ϕ , we obtain σ_{sr} in the rational expectations equilibrium:

$$\mathbf{s}_{sr} = -\frac{\mathbf{l}(\mathbf{l} + \mathbf{h})(\mathbf{l} + \mathbf{h})\mathbf{s}_e^2}{\mathbf{l} + \mathbf{h} + \mathbf{hl}} < 0. \quad (27)$$

Therefore, by (15),

$$\mathbf{f} = \frac{\mathbf{l}(\mathbf{l} + \mathbf{h})}{\mathbf{l} + \mathbf{h} + \mathbf{hl}} > 0. \quad (28)$$

In the rational expectations equilibrium, the conditional covariance between the exchange rate and the one-period interest rate, σ_{sr} , is negative. Hence, foreign bonds and two-period bonds are substitutes as explained in the previous section. The measure of the relative magnitude of the indirect complements relationship, ϕ , is positive. The main issue for the purpose of this paper is whether ϕ is greater or less than one.

In order to examine the magnitude of ϕ in (27), we need to know how λ depends on the underlying parameters of the model. Ogaki (1996) analyzes this dependence by solving for the rational expectation of the conditional variance of the exchange rate. He shows that there exists a unique rational expectations equilibrium for the model, and that ϕ is greater than one when the investor's risk aversion parameter, α , and the measure of the degree of substitution between one-period and two-period bond, η , are close to zero. Because ϕ can be smaller than one for other parameter configurations, whether or not ϕ is greater than one is an empirical question.

III. DATA AND PRELIMINARY CONSIDERATIONS

We study the effect of the term structure of interest rates on the exchange rate for Mexico, using the United States as the foreign country. The data series for Mexico are monthly, and cover the period from January 1983 to June 1997. The nominal one-month and three-month interest rates are monthly averages of the weekly auctions of the 28 day and 91 day Cetes (T-Bills), respectively, and were taken from *Indicadores Económicos* (Banco de México). The consumer price index is used as the price level. The data for United States cover the same period. The one-month and three-month Euro Market rates are used for the nominal interest rates. The weekly data obtained from the INTLINE data base are converted into monthly data within AREMOS II. The consumer price index is also used as the price level. The nominal exchange rate data, expressed as the price of the U.S. dollar in terms of the Mexican peso, were also from *Indicadores Económicos*. The real exchange rate is calculated as the nominal exchange rate multiplied by the U.S. price level divided by the Mexican price level. The data are not seasonally adjusted. The inflation rate is the first difference of the log of the price level. The real interest rate is the log of one plus the nominal interest rate minus the expected inflation rate, which is estimated as explained below.

For Mexico, the three-month nominal interest data are missing for August 1986, September 1986, and November 1988 (no Cetes were sold in the weekly auctions for these months). A vector autoregression (VAR) of order five for the inflation rate, the one-month nominal interest rate, and the three-month nominal interest rate was used to assign values to these periods. The VAR was estimated using the data up to July 1986, and the projected values for August 1986 and September 1986 were assigned to these periods. Then the VAR was re-estimated using the data up to October 1988, and the projected value for November 1988 was assigned to this period.

The model is specified in real interest rates, so an estimate of expected inflation was constructed. The three-month real interest rate differential, $(r_{3,t} - r_{3,t}^*)$, is used to capture the effect of the interest rate shock with the longer duration in the model, e_t . In order to estimate the real interest rates, a VAR of order five for the inflation rate, the one-month nominal interest rate, and the three-month nominal interest rate was estimated for each country using the whole sample period. Then the three-period ahead forecast of the inflation rate was calculated from the estimated VAR in each period. The forecast was subtracted from the three-month nominal interest to calculate the real interest rate.

We use the normalized one-period interest rate, $r_{N,t}$ to capture the effect of the temporary interest shock. The normalized one-period interest rate is defined as follows:

$$r_{N,t} = \frac{1}{3} \sum_{k=0}^2 E_t(r_{1,t+k}) - r_{3,t}, \quad (29)$$

where $r_{1,t}$ and $r_{3,t}$ are the one-month, and three-month nominal interest rates, respectively. The normalized one-month interest rate is the negative risk premium for the three-month

bonds, just as $_t$ is the negative risk premium for the two-period bonds in the model. It captures the movements in one-month interest rates as a deviation from three-month interest rates. The normalized interest rate is estimated using the one-period and two-period ahead forecasts from the estimated VAR for each country.

Table 1 reports results for Said and Dickey's (1984) test for a unit root. Let $SD(p)$ be Said and Dickey's test with the order of autoregression equal to p is denoted by $SD(p)$. The table reports $SD(1)$, $SD(5)$, and $SD(10)$ for each series. A constant is included for the regression. The first row reports results for the log real exchange rate, s . None of the test statistics is significant at the 5 percent level. Hence the null hypothesis that the log real exchange rate is unit root nonstationary cannot be rejected. For the three-month real interest rate differential, $(r_{3,t} - r_{3,t}^*)$, $SD(1)$ is significant, but $SD(5)$ and $SD(10)$ are not significant at the 5 percent level. For the normalized one-month interest rate differential, $(r_{N,t} - r_{N,t}^*)$, $SD(1)$ and $SD(5)$ are significant, but $SD(10)$ is not significant at the 5 percent level. Hence the results for the interest rate differentials depend on the order of the autoregression. Standard economic theory suggests that these interest rate differentials can hardly be unit root nonstationary. However, these ambiguous test results suggest that there exists a persistent component in each interest rate differential, which might be locally approximated by unit root nonstationarity.

IV. THE ECONOMETRIC METHOD

The model in Section II abstracts from some features of data, such as nonstationarity of the real exchange rate, so that it can be solved for the rational expectations equilibrium. In our empirical analysis, we follow the spirit of the reduced form of the model, Equation (25), while incorporating important features of data. The cointegrating regression we use cannot be viewed as a structural regression, but should be viewed as a method to find a stylized fact about the relationship between the exchange rate and the term structure of interest rates. While interesting, it is beyond the scope of this paper to derive structural econometric models for the exact model in Section II in the presence of cointegrating relationships.

The regression we employ is of the form

$$s_t = \mathbf{m} + \mathbf{a}(r_{3,t} - r_{3,t}^*) + \mathbf{b}(r_{N,t} - r_{N,t}^*) + \mathbf{h}_t, \quad (30)$$

where the three-month interest rate differential, $(r_{3,t} - r_{3,t}^*)$, is used to capture the effect of the interest rate shock with the longer duration in the model, e_t .³ We use the normalized one-period interest rate, $r_{N,t}$ to capture the effect of the temporary interest shock.

³ Note that the model in Section 2 suggests an regression of a real exchange rate onto short-term and long-term real interest rate differentials. Frankel (1979) and Boughton (1988) regress exchange rates onto both short-term and long-term interest differentials. Apart from the fact that they do not interpret their regressions in terms of cointegration, our regressions

Because the null hypothesis that the real exchange rate is difference stationary cannot be rejected, we view (30) as a cointegrating regression. We apply Park's (1992) Canonical Cointegrating Regression (CCR) to (30).⁴ This procedure allows us to test the null hypothesis of cointegration and the deterministic cointegration restriction. According to our model solution α is negative as in the standard exchange rate models. However, the sign of β can be positive if the indirect complementarity between the one-month domestic and foreign bonds dominates their direct substitutability.

To understand CCR, consider a cointegrated system

$$y = X_t' \mathbf{g} + \mathbf{h}_t \quad (31)$$

$$\Delta X_{2t} = \mathbf{n}_t, \quad (32)$$

where $X_t = (1, X_{2t})$, y_t and X_{2t} are difference stationary, and η_t and \mathbf{v}_t are stationary with zero mean. Here y_t is a scalar and X_t is a $(n-1) \times 1$ random vector. Let

$$w_t = (\mathbf{h}_{1t}, \mathbf{n}_{1t}') \quad (33)$$

Define $\Phi(i) = E(w_t w_{t-i}')$, $\Sigma = \Phi(0)$, $\Gamma = \sum_{i=0}^{\infty} \Phi(i)$, and $\Omega = \sum_{i=-\infty}^{\infty} \Phi(i)$. Here Ω is the long run covariance matrix of w_t . Partition Ω as

$$\Omega = \begin{pmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{21} & \Omega_{22} \end{pmatrix} \begin{matrix} I \\ n-1 \end{matrix} \quad (34)$$

and partition Γ conformably. Define

$$\Omega_{11.2} = \Omega_{11} - \Omega_{12} \Omega_{22}^{-1} \Omega_{21} \quad (35)$$

and $\Gamma_2 = (\Gamma'_{12}, \Gamma'_{22})'$. The CCR procedure assumes that Ω_{22} is positive definite, implying that X_t is not itself cointegrated (see, e.g., Phillips (1986) and Engle and Granger (1987)). This assumption assures that $(1, -\gamma)$ is the unique cointegrating vector (up to a scale factor).

The OLS estimator in (31) is super-consistent in that the estimator converges to γ at the rate of T (sample size) even when ΔX_{2t} and \mathbf{v}_t are correlated. The OLS estimator, however, is not asymptotically efficient.

Consider transformations

are different from theirs because they use real and nominal interest rate differentials. Frankel uses the real short-term interest rate differential and the nominal long-term interest rate differential. Boughton uses the nominal short-term interest rate differential and the real long-term interest rate differential.

⁴ Byeon (1996) applies different asymptotically efficient estimation methods for similar cointegrating regressions with G-7 country data, and finds that his results are similar.

$$y_t^* = y_t + \Pi_y' w_t \quad (36)$$

$$X_{2t}^* = X_{2t} + \Pi_x' w_t,$$

and $X_{2t}^* = (1, X_{2t}^*)'$. Because w_t is stationary, y_t^* and X_{2t}^* are cointegrated with the same cointegrating vector $(1, -\gamma)$ as y_t and X_t for any Π_y and Π_x .

The idea of the CCR is to choose Π_y and Π_x , so that the OLS estimator is asymptotically efficient when y_t^* is regressed on X_{2t}^* . This requires

$$\Pi_y = \Sigma^{-1} \Gamma_2 \mathbf{g} + (0, \Omega_{12} \Omega_{22}^{-1})' \quad (37)$$

$$\Pi_x = \Sigma^{-1} \Gamma_2.$$

In practice, long-run covariance parameters in these formulas are estimated, and estimated Π_y and Π_x are used to transform y_t and X_{2t}^* . As long as these parameters are estimated consistently, the resultant CCR estimator is asymptotically efficient.

The CCR estimators have asymptotic distributions that can be essentially considered as normal distributions, so that their standard errors can be interpreted in the usual way.⁵ An important property of the CCR procedure is that linear restrictions can be tested by χ^2 tests that are free from nuisance parameters. We use χ^2 tests in a regression with spurious deterministic trends added to (31) to test for cointegration. For this purpose, the CCR procedure is applied to a regression

$$y_t^* = \mathbf{m} + \sum_{i=1}^q \mathbf{y}_i t^i + \mathbf{g} X_{2t}^* + \mathbf{h}_t^*.$$

Let $H(0, q)$ denote the standard Wald statistic to test the hypothesis $\psi_1 = \psi_2 = \dots = \psi_q = 0$ with the estimate of the variance of η^* replaced by $\Omega_{11.2}$ (see Park (1990) for details). Then $H(0, q)$ converges in distribution to a χ^2_q random variable under the null hypothesis of cointegration, while it diverges under the alternative hypothesis of no cointegration because the trend terms become spuriously significant.

During the 1983:1-1997:6 period used in our estimations, there have been exchange regime changes in Mexico (see Santaella and Vela 1996). These may have caused shifts in the

⁵ The CCR estimators are asymptotically efficient, but there are other asymptotic efficient estimators by Phillips and Hansen (1990). Monte Carlo experiments in Park and Ogaki (1991) show that the CCR estimators have better small sample properties in terms of the mean square error than Johansen's ML estimators. Following Monte Carlo based recommendations by Park and Ogaki (1991) and Han and Ogaki (1991), we used the prewhitening method and report third stage CCR estimates and fourth stage CCR $H(p, q)$ test statistics.

long-run relationship between the real exchange rate and the real interest rate differentials, even though it is theoretically possible for the relationship to be stable over different regimes. For this reason, it is necessary to test whether or not there have been structural breaks in the regression (30). We use Hansen's (1992) tests for parameter instability for this purpose.

Hansen's tests are developed for Phillips and Hansen's (1990) fully modified estimation method, but can be adapted to the CCR method as follows because these two methods are very similar.⁶ In the regression model (31), we now allow g to be different over two subsamples:

$$\mathbf{g} = \begin{cases} \mathbf{g}_a & \text{if } t \leq \tau \\ \mathbf{g}_b & \text{if } t > \tau \end{cases}$$

where τ is the timing of the structural break. We consider two tests developed by Hansen (1992) for the null hypothesis $H_0: \gamma_a = \gamma_b$. The null hypothesis is the same for both tests, but they differ in the treatment of alternative hypotheses: for the first test, the timing of the structural break is known, while for the second test, it is unknown.

The first test, denoted by F_τ , is essentially the classical Chow test for a known structural break date of τ . Under the null hypothesis of no break, we transform the data as in (35). Define

$$d_t = \begin{cases} 1 & \text{if } t \leq \tau \\ 0 & \text{if } t > \tau \end{cases}$$

Then apply the OLS to

$$y_t^* = d_t X_t^{*'} \mathbf{g}_a + (1 - d_t) X_t^{*'} \mathbf{g}_b + \mathbf{h}_t^* \quad (38)$$

Then F_τ is the standard Wald statistic to test the hypothesis $\gamma_a = \gamma_b$ with the estimate of the variance of η^* replaced by $\Omega_{11.2}$. Then F_τ has an asymptotic chi-square distribution with degrees of freedom equal to the number of parameters in γ .

For the second test, the timing of the structural break is unknown. The test statistic, denoted by SupF , is defined by $\text{SupF} = \sup F$ where supremum is taken over the

⁶ We thank Joon Park for his helpful advice on adapting the tests.

range of $0.15T \leq \tau \leq 0.85T$, where T is the sample size. Asymptotic critical values of this test statistic are tabulated in Hansen (1992).

V. EMPIRICAL RESULTS

Our empirical results are consistent with the proposition that the term structure of interest rates matters for the exchange rate in the case of Mexico. The first row of Table 2 presents cointegrating regression results when the whole sample period is used. The sign of the point estimate of α is negative and statistically significantly different from zero. A negative α is in the standard direction of increases in interest rates leading to exchange rate appreciation as expected from the conventional wisdom, and is also consistent with our model in Section II. However, contrary to the traditional intuition, the point estimate of β is positive and significantly different from zero. A natural explanation arising from our theoretical analysis is that the indirect complementarity dominates the direct substitutability. The implication of this result is that increases in the three-month interest rate differential will, *ceteris paribus*, be associated with a depreciating Mexican peso in the long-run.

As suggested above, the possibility of structural breaks is an issue in the case of Mexico. Table 2 shows that the SupF test is significant at the 5 percent level, indicating that a structural break is present. The F_τ test statistics were computed for possible structural break dates from November 1985 to December 1995. The F_τ statistics are below their 5 percent critical value from November 1985 to July 1986, above their 5 percent critical value from December 1985 to March 1988, and below their 5 percent critical value from April 1988 to December 1995. The F_τ attains the maximum value of 15.881 in November 1987, which is above its 1 percent critical value.

The timing of the structure break has a natural explanation. A comprehensive stabilization plan was implemented in December 1987 in Mexico. Hence these parameter instability test results indicate that the stabilization plan changed the cointegrating relationship between the real exchange rate and the term structure of the real interest rate differentials.

Because the tests for parameter instability used in this paper assume only one structural break, cointegrating regressions were applied to the sample period from January 1988 to May 1997 in order to confirm that there is no additional structural break. The second row of Table 1 presents these cointegrating regression results.⁷ It is noteworthy that the estimate of α continues to be significantly negative, while the estimated β remains significantly positive, thus maintaining the same qualitative relation of the term structure of interest rates to the exchange rate despite the structural break.

⁷ The long-run covariance parameters for these results used for the third stage and fourth stage CCR are taken from the third stage and fourth stage CCR reported in the first row of Table 1.

After the stabilization plan was implemented, several exchange regimes have been adopted in Mexico ranging from a fixed exchange rate regime to a floating regime. The test results for parameter instability, however, indicate that the cointegrating relationship between the real exchange rate and the term structure of the real interest rate differentials may have been stable over this period. The SupF test statistic is not significant at the conventional significance levels. No individual F statistic computed from December 1988 to June 1996 is above its 38 percent critical value.

With an additional assumption that the normalized real one-month interest rate differential is predetermined with respect to both the real three-month interest rate differential and the real exchange rate, and that the three-month interest rate differential is predetermined with respect to the real exchange rate, one can compute the orthogonalized impulse response functions. These assumptions are stringent, but it is often of interest to examine an example of dynamic responses. Under these assumptions, the orthogonalized impulse response functions are defined with respect to the disturbances in the structural error correction model (see, e.g., Kim, Ogaki and Yang 1999).

Imposing the estimates of the cointegrating vector for the sample period of 1988:1-1997:5, we estimate the error correction model with OLS for each equation. This method to estimate the error correction model is the same as Engle and Granger (1987) except that the first step estimator for the cointegrating vector here is from the asymptotically efficient CCR estimator rather than the inefficient OLS estimator. From the estimated error correction model, the VAR representations for the levels of the three variables in the system are computed. The orthogonalized impulse responses are calculated with the standard method from this VAR representation (see, e.g., Hamilton 1994 for the method).

The orthogonalized impulse response functions of the real exchange rate to the real interest rate differential shocks are plotted in Figure 1. Because the variables can be locally approximated by unit root nonstationary and are cointegrated, interest rate shocks have permanent effects on the exchange rate, as implied by the cointegrating vector. The initial impact of a one-percent positive shock to the one-month interest rate is a 0.102 increase in the exchange rate, while a one-percent positive shock to the three-month interest rate is a 0.075 percent fall in the exchange rate. Thus, the estimated initial impact works in the same direction as implied by the cointegrating vector for the long-run for each real interest rate differential.

After the impact effects, the dynamics become very complicated. One month after the interest rate shock, the response is in the opposite direction from the initial impact for each interest rate differential. From the second month on, the response of the exchange rate to the three-month interest rate shock reverts to being negative, and remains negative as implied by the cointegrating vector. However, the response of the exchange rate to orthogonalized shocks to the one-month interest rate differential does not revert immediately to the long-run cointegrating relationship. From the first month on, the response to the one-month interest

rate shock is negative for a long time, and it takes 152 months for the response to become positive. Thus a counterintuitive result from these impulse responses, which is nonetheless possible in our model, is that the Mexican peso depreciates in the medium-run when the one-month interest rate rises in Mexico.

VI. CONCLUSIONS

Conventional wisdom is that increases in any interest rate would appreciate the domestic currency, and thus that the term structure of interest rates is irrelevant for the determination of the exchange rate. This paper has argued that this conventional wisdom can be challenged both on theoretical and empirical grounds.

We study the relationship between the term structure of interest rates and the exchange rate for Mexico. Our empirical results suggest that interest rates may have counterintuitive effects on the exchange rate, which may arise from the complementarity between one-month domestic bonds and foreign bonds. The cointegrating regression estimated in this paper indicates that one-month and three-month interest rate differentials have opposing effects on the exchange rate for Mexico. While increases in the one-month interest rate differential will, other things equal, affect the exchange rate in the standard appreciation direction, increases in the three-month rate differential tend to depreciate the exchange rate. Therefore, the effect of the one-month interest rate differential on the exchange rate depends on how the three-month interest rate moves in response to changes in the one-month interest rate. Imagine for the moment that Bank of Mexico affects liquidity conditions in order to raise the one-month interest rate, intending to cause an appreciation of the peso. If the three-month interest rate rises in response to the rise in the one-month rate, then the peso will tend to appreciate as in the standard exchange rate model that assumes risk neutral investors. However, if the three-month interest rate does not rise, then the peso will tend to depreciate.

Because the exchange regime of Mexico has been changing, we performed structural break tests for parameter instability for the cointegrating regression of the real exchange rate onto the one-month and three-month real interest rate differentials. Even though the test results indicate breaks in the cointegrating relationship around December 1987 when a comprehensive stabilization plan was implemented in Mexico, they indicate that the cointegrating relationship has been stable since then despite several exchange regime changes.

The empirical results obtained in this paper could be useful to understand and manage the operation of the Mexican floating exchange rate regime. They show that attention must be given to the term structure of interest rates when considering exchange rate objectives. The fact that the traditional intuition may not work, should also provide a warning to other developing countries have now adopted floating exchange rate regimes.

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Table 1. Unit Root Test Results

	SD(1)	SD(5)	SD(10)
s_t	-1.877	-1.742	-2.425
$\Gamma_{3,t} - \Gamma_{3,t}^*$	-6.435	-2.786	-2.559
$\Gamma_{N,t} - \Gamma_{N,t}^*$	-6.165	-4.286	-2.530

Note: SD(p) is the Said and Dickey's (1994) test with order of autoregression equal to p.

Table 2. Canonical Cointegrating Regression Results

Sample Period (1)	$\hat{\mathbf{a}}$ (2)	$\hat{\mathbf{b}}$ (3)	H(0,1) (4)	H(0,2) (5)	SupF (6)
1993:7-97:5	-2.922 (0.278)	5.427 (1.107)	0.134 (0.714)	0.768 (0.681)	15.881 (0.049)
1988:1-97:5	-3.061 (0.4420)	6.976 (1.2860)	1.225 (0.2680)	1.248 (0.536)	2.112 (>0.200)

Note.-In cols. 2 and 3, standard errors are in parentheses. Cols. 4 and 5 report χ^2 test statistics for stochastic cointegration with one and two degrees of freedom, respectively. P-values are in parentheses. Col. 6 reports SupF test statistic for a structural break with an unknown timing.

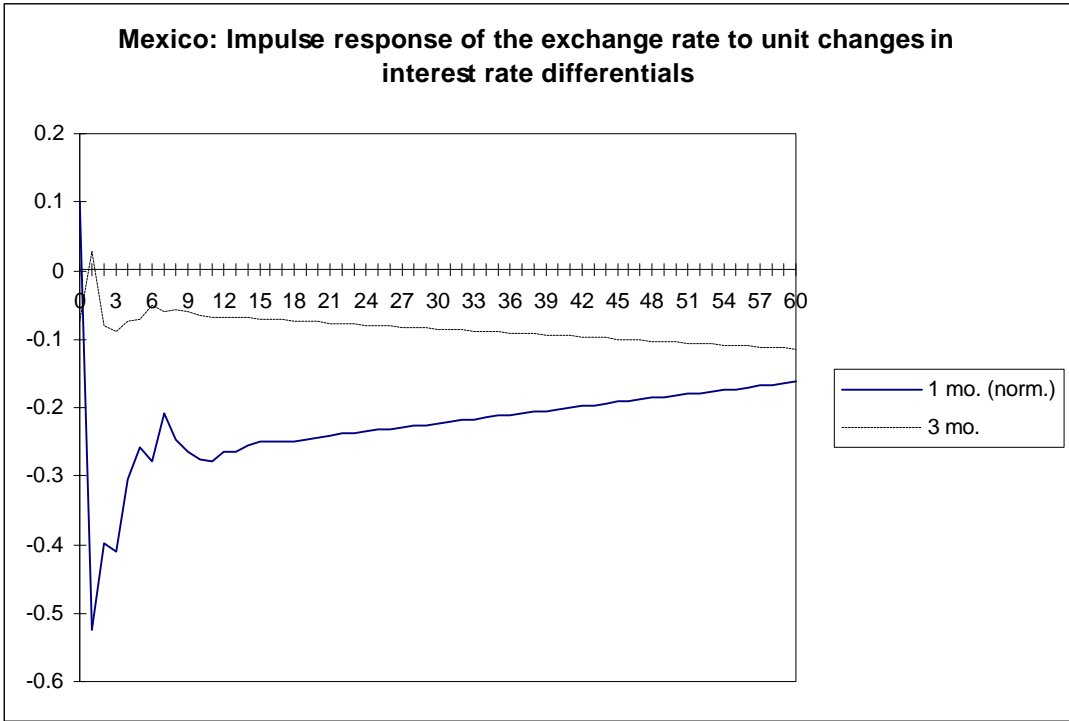


Figure 1.