Exchange risk premia, expectations formation and "news" in the Mexican peso/U.S. dollar forward exchange rate market

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Abstract

In this article, we investigate expectations concerning the Mexican peso/US dollar exchange rate with the aid of a survey dataset containing market participants' forecasts of the exchange rate and of the interest differential between the peso and the dollar. Our findings indicate that the survey expectations were off by a large and significant constant. At the same time, large average risk premia, as well as time variance in the risk premia, were detected. As to the expectations formation mechanism, market participants tended to react to current (unanticipated) depreciations by expecting future depreciations at the 3-, 6-, and 12-month horizons, as implied by destabilizing expectations models. "News" about the interest differential did not contribute additional predictive power with regard to the peso/dollar exchange rate once a risk premium term is included in a regression equation. Interestingly, a Dornbusch-type overshooting effect is present in the Mexican data. © 2001 Elsevier Science Inc. All rights reserved.

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1. Introduction

The debate regarding the empirical finding that the forward discount is a biased predictor of the future change in the exchange rate continues to be an issue of central concern in the
international financial economics literature (for instance, see the surveys on the efficiency of the forward foreign exchange markets by Engel, 1996; Hodrick, 1987). The early empirical evidence suggests that future exchange rate changes and current interest differentials (i.e., the forward discount) are negatively correlated. That is, relatively high domestic nominal interest rates predict appreciation of the domestic currency. The rejection of forward market efficiency may be attributable to the irrationality of market participants, to the existence of time-varying risk premia, learning about a policy change, or to some combination of these phenomena (see, e.g., Cavaglia, Verschoor, & Wolff, 1993a, 1993b, 1994; Frankel & Froot, 1987a, 1987b; Froot & Frankel, 1989; Lewis, 1995).

Interestingly, this branch of literature concerning the forward discount bias pays little attention to emerging and the lower-income developed economies. In the current article, we aim to provide statistical evidence on the nature of the forward discount bias for the Mexican forward exchange rate market using survey expectations of the Mexican peso/US dollar exchange rate. By gathering independent measures of expectations, it is be possible to decompose the forward discount bias into separate components attributable to risk premia and to expectational errors. The study complements previous work that has largely focused on analyzing survey data for the five most actively traded currencies vis-a-vis the US dollar, and on EMS currencies. In addition, we address the question that was considered earlier by Cavaglia et al. (1993a) and Frankel and Froot (1987a): Which time series process best characterizes investors’ expectations formation?

In the past two decades, the asset market approach has become the principal tool for analyzing movements in exchange rates. From the time when expectations were first introduced into this approach to exchange rate determination, it has been recognized that unexpected events have a qualitatively different effect on the exchange rate from anticipated developments. The exchange rate should change discontinuously in response to new pieces of unanticipated relevant information, whereas anticipated discrete changes are ruled out, since they would represent an unexploited profit opportunity. The asset market approach typically places considerable emphasis on the importance of expectations and changes therein and is generally taken to imply that empirical research on the determinants of exchange rates should relate innovations in exchange rates to innovations in expectations about relevant future fundamentals. In the empirical literature, this approach is often referred to as the “news” approach of exchange rate determination. Since expectations are inherently unobservable, any empirical study on the “news” approach involves choosing a specific model of the process of exchange rate determination and an appropriate method of generating expected values of its driving values. In this article, we follow an alternative route and investigate empirically the relationship among exchange rate returns, “news,” and risk premia using survey data of matched Mexican peso/US dollar exchange and interest rate expectations, thereby at least partially avoiding the problem of artificially generated expectations when using an econometric technique.

1 One interesting study is worth noting. Bansal and Dahlquist (1999) provide empirical evidence regarding the “forward premium puzzle” obtained from emerging and developed economies.

2 See Frenkel and Mussa (1980). Black (1973) was an early introduction of rational expectations and a test of anticipated vs. unanticipated effect of news reports.
This article, presented in six sections, extends the findings of Cavaglia et al. (1993a, 1993b, 1994), Frankel and Froot (1987a, 1987b), and Froot and Frankel (1989) by considering a new survey dataset that covers the Mexican peso relative to the US dollar over the January 1988–May 1992 period. Our results are easily summarized. The survey expectations were off by a large and significant constant. At the same time, a large average risk premium, as well as time variance in the risk premium, was detected. As to the expectations formation mechanism, market participants tended to react to current (unanticipated) depreciations by expecting future depreciations at the 3-, 6-, and 12-month horizons, as implied by destabilizing expectations models. “News” about the interest differential does not contribute additional predictive power with regard to the peso/dollar exchange rate once a risk premium term is included in a regression equation. Interestingly, a Dornbusch-type overshooting effect is present in the Mexican data.

In Section 2, the construction of the exchange rate survey is outlined and summary statistics describing the data are provided. In Sections 3 and 4, we address the principal question of whether rejection of forward market efficiency is attributable to the existence of time-varying risk premia or irrational behavior on the part of economic agents. Alternative models characterizing the formation of exchange rate expectations are considered in Section 5. Finally, in Section 6, we investigate empirically the relationship between interest rate “news” and exchange rate surprises. The empirical results of this investigation on Mexican peso/US dollar exchange rate expectations are summarized in Section 7.

2. The survey data

Since 1985, Business International has conducted a monthly survey of exchange rate expectations covering, among others, the Mexican peso relative to the US dollar, which are published in its Cross Rate Bulletin. For publication purposes, survey participants were asked a few days prior to month’s end to fax 3-, 6-, and 12-month-ahead expectations of the currency, with projections being made from the beginning of the following month. Thus, for instance, the 3-, 6-, and 12-month-ahead expected Mexican peso/US dollar rates recorded on December 27, 1989 reflect a slightly longer forecast horizon as they represent the expected spot rate on April 1, 1990, June 1, 1990, and January 2, 1991, respectively. The dates when the surveys were conducted were recorded as well as the spot rate on that particular day.

Since 1988, survey respondents were also asked to provide their 3-, 6-, and 12-month-ahead expectations regarding domestic interest rates with 1- and 3-month maturities. Since part of our study is concerned with matched interest rate and exchange rate expectations, survey data availability led us to focus our analysis on the 3-, 6-, and 12-month-ahead Mexican peso/US dollar exchange rate and the subsequent interest rate expectations. Actual

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3 For some of the tests conducted in this article, we were able to extend the sample period. Unfortunately, our dataset does not contain information about NAFTA and the tequila crisis imposed on investors’ expectations.

4 Although the notation used in Sections 3–6 will be presented as if the survey were constructed on December 31 (in the example at hand), care has been exercised throughout the empirical analysis to ensure that conditional expectations are computed on the proper information set.
spot and interest rates used in this study were obtained from Datastream. The 30 odd participants of the survey are treasurers of multinationals and private banks residing in four of the world’s continents. Although not all participants will provide their views regarding a particular currency, the response rate is, at worst, 60%. The Cross Rates Bulletin reports the geometric mean forecast of the responses received, thus minimizing the effect of extreme forecasts.

The use of survey data allows the direct measurement of a risk premium: conditional on market efficiency and rational expectations, the forward exchange rate is equal to the expected future spot rate plus a risk premium. Thus, the forward discount can be decomposed into two components — the expected rate of depreciation and the risk premium:

\[ f_{t+k} - S_t = (E_t S_{t+k} - S_t) + P^k_t. \]  

Here, \( S_t \) is defined as the natural logarithm of the spot exchange rate at time \( t \), \( E_t S_{t+k} \) is defined as the expected logarithm of the spot exchange rate at time \( t+k \) formed at time \( t \), \( f_{t+k} \) is defined as the natural logarithm of the forward rate at time \( t \) for delivery at time \( t+k \), and \( P^k_t \) is the associated risk premium. In the remainder of the paper, \( d_t \) is defined as the 3-month domestic interest rate differential, \( i_t - i^*_t \), for deposits starting at time \( t \) and maturing at \( t+3 \), and \( E_t d_{t+k} \) as the expectation of the 3-month interest rate differential at time \( t+k \), conditional on information available at time \( t \). Here, \( i_t \) is the domestic 3-month interest rate and \( i^*_t \) the foreign 3-month interest rate. Because the survey expectations are direct estimates, we do not need to assume any particular model of expected depreciation or of the risk premium. Note that Eq. (1) is essentially the definition of the risk premium. To give Eq. (1) economic content, a model of international asset pricing that describes the determination of \( P^k_t \) is required.\(^5\)

Table 1 provides summary statistics for the actual and expected annualized exchange rate depreciation, the survey forecast error, the forward discount, and foreign exchange risk premium across forecast horizons.\(^6\)

For the period analyzed (January 1, 1988 through May 1, 1992), the survey consistently called for an upward movement in the value of the Mexican peso/US dollar exchange rate, whereas the standard deviation of the mean expected depreciation declines with the forecast horizon. The mean expected rates of depreciation are the same in sign, but smaller in magnitude than the time series average of the forward discount, suggesting the presence of exchange risk premia and thus implying that Mexican and US assets are regarded as imperfect substitutes. It is interesting to note that, in contrast to the absolute values, the standard deviations of the mean survey forecast error decline as the length of the forecast horizon increases from 3 to 12 months. Frankel and Froot (1986) report similar results; however, their empirical observation is reversed in Frankel and Froot (1987a) when four data points are added to the sample.

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\(^{5}\) Equilibrium models of international asset pricing that provide us with such descriptions are presented, for instance, in Adler and Dumas (1983), Hodrick (1981), Hodrick and Srivastava (1984), Roll and Solnik (1977), and Stulz (1981).

\(^{6}\) Denoting the forecast horizon in months as \( k \), annualized returns are obtained by multiplying the log differences by 1200/\( k \).
Table 1
Summary statistics February 1, 1988 through May 1, 1992 (percent per annum)

<table>
<thead>
<tr>
<th>Time to horizon</th>
<th>3 months</th>
<th>6 months</th>
<th>12 months</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Mean</td>
<td>S.D</td>
<td>Mean</td>
</tr>
<tr>
<td>$S_{t+k} - S_t$</td>
<td>20.89</td>
<td>33.77</td>
<td>19.94</td>
</tr>
<tr>
<td>$E_s S_{t+k} - S_t$</td>
<td>36.44</td>
<td>35.15</td>
<td>41.70</td>
</tr>
<tr>
<td>$S_{t+k} - E_s S_{t+k}$</td>
<td>-15.55</td>
<td>40.76</td>
<td>21.75</td>
</tr>
<tr>
<td>$F_t, S_t$ - $S_t$</td>
<td>85.44</td>
<td>74.00</td>
<td>73.12</td>
</tr>
<tr>
<td>$F_{t+k} - E_s S_{t+k}$</td>
<td>49.00</td>
<td>90.16</td>
<td>31.42</td>
</tr>
</tbody>
</table>

Actual depreciation: $S_{t+k} - S_t$, expected depreciation: $E_s S_{t+k} - S_t$; survey forecast error: $S_{t+k} - E_s S_{t+k}$; forward discount: $F_t, S_t - S_t$; and risk premium: $F_{t+k} - E_s S_{t+k}$.

Fama (1984) demonstrates that rejection of the unbiasedness hypothesis in most instances implies that risk premia are more variable than expected rates of depreciation and that the two covary negatively. From Table 1, one notes that the standard deviations of the expected depreciation are smaller than the standard deviation of the risk premium. This contrasts with Froot and Frankel (1989), who find that the variability of the risk premium is generally smaller than that of the expected depreciation. Since the risk premium is just equal to the interest rate differential (or forward discount) less the expected change in exchange rates, a greater variance of the risk premium than the variance of the expected depreciation might be considered plausible: an increase in the interest rate differential is then associated with a decline in expected depreciation and therefore with an even larger rise in the risk premium.  

3. Tests of forward discount bias

Forward market efficiency has generally been tested by regressing the observed change in the spot exchange rate onto the forward discount. Thus, the null hypothesis of unbiasedness implies that it is possible to decompose $S_{t+k} - S_t$ as:

$$S_{t+k} - S_t = \alpha + \beta(F_{t+k} - S_t) + e_{t+k},$$

with $\alpha = 0$, $\beta = 1$, and $e_{t+k}$ has mean zero and is uncorrelated with $F_{t+k} - S_t$. Eq. (2) was estimated by ordinary least squares (OLS) for each forecast horizon ($k = 3, 6,$ and 12 months). Realized spot exchange rates were obtained from Datastream.  

Hansen and Hodrick (1980) demonstrated that, when the forecast horizon is longer than the observational frequency, the forecast error $e_{t+k}$ will be serially correlated. While OLS point estimates of $\beta$ remain consistent in spite of the serially correlated residuals, the OLS standard errors for the

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7. By arbitrage, the forward discount must equal the interest differential. If it did not, then a strategy of borrowing in the foreign currency, changing the proceeds into Mexican pesos, investing those pesos, and then selling them forward would yield a riskless profit.

8. The spot exchange rates at time $t+k$, $S_{t+k}$, used to compute the change in the spot rate are obtained from Datastream on days corresponding to the survey forecast dates. If the forecast date falls on a holiday or weekend, the next business day is chosen.
regression coefficient are biased. This can be corrected via the Newey and West (1987) estimation procedure. Therefore, the $k$-month-ahead forecast equations in this section are estimated with the Newey–West estimator, assuming a moving average process of order $k - 1$ for the monthly $k$-month-ahead forecast errors.\footnote{See Cavaglia et al. (1993a) for a more detailed description. Note that the $k$-month-ahead forecast is, in reality, a $k$-month plus a few days ahead forecast.}

As is well known, the results of many previous studies suggest rejection of the null hypothesis across the spectrum of forward rates. Oftentimes, the estimate of $\beta$ is reliably less than one. In fact, $\beta$ is frequently estimated to be less than zero, evidenced by an average coefficient of $-0.88$ across some 75 published estimates (see Froot & Thaler, 1990).

Table 2A reports the results of fitting Eq. (2) for each forecast horizon via OLS, with Newey–West standard errors. Overall, the results indicate a sound rejection of the null hypothesis that the forward discount of the Mexican peso is an unbiased predictor of the future change in the Mexican peso/US dollar exchange rate. All $\beta$ estimates are statistically significantly negative. Also, the constant term $\alpha$ is significantly positive in all cases, reflecting a large systematic bias. A finding of a negative $\beta$-coefficient implies that, e.g., when Mexican peso interest rates exceed foreign (US dollar) rates, the Mexican peso subsequently tends to appreciate. This is in contrast to the expected depreciation of the Mexican peso dictated by the unbiasedness hypothesis.

Our result is less apparent for some of the cross exchange rates within the European Monetary System (EMS). Cavaglia et al. (1994) find positive $\beta$-coefficients for EMS currencies relative to the Deutschmark. In addition, Bossaerts and Hillion (1991) find positive estimates of $\beta$ for most currencies against the French franc, whereas Flood and Rose (1996) find higher $\beta$-coefficients for economies within the EMS versus the Deutschmark than for economies versus the US dollar.\footnote{Bansal and Dahlquist (1999) present evidence from emerging and lower-income countries that is consistent with economic intuition: a positive domestic interest rate differential predicts a depreciation of the domestic currency.} Within the EMS, changes in the interest differential may only reflect changes in inflationary expectations rather than changes in real interest rates. If this is the case, the bias seems less severe and tests of the bias show $\beta$-coefficients that are positive and nearer one, suggesting that the quality of the forward discount as a predictor of future exchange rates changes is higher in the case of EMS exchange rates.

Two interpretations of these results are common in the literature. Rejection of forward market efficiency has often been attributed to either the failure of rational expectations or the existence of time-varying risk premia. In this context, Frankel and Froot (1987a), Froot and Frankel (1989), and Taylor (1989) demonstrated how survey expectations data can be exploited to ascertain the economic importance of these competing explanations. In Eq. (2), the probability limit of the estimate of the $\beta$-coefficient is [Eq. (3)]:

$$\beta = \frac{\text{cov}(F_{t+k} - S_t, S_{t+k} - S_t)}{\text{var}(F_{t+k} - S_t)}. \quad (3)$$

Defining $u_{t+k}$ to be the $k$-month-ahead expectations forecast error, $S_{t+k} - E_t S_{t+k}$, and using the decomposition of the forward discount in Eq. (1), it follows that [Eq. (4)]:

$$\beta = \beta_1 + 3_2. \quad (4)$$
### Table 2

(A) Tests of forward discount unbiasedness: \( S_{t+k} - S_t = \alpha + \beta_1 (F_{t+k} - S_t) + \epsilon_{t+k} \) from February 1, 1988 through May 1, 1992

<table>
<thead>
<tr>
<th></th>
<th>3 months</th>
<th>6 months</th>
<th>12 months</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \alpha )</td>
<td>0.0462*** (0.0062)</td>
<td>0.0765*** (0.0117)</td>
<td>0.1374*** (0.0234)</td>
</tr>
<tr>
<td>( \beta_1 )</td>
<td>-0.0691*** (0.0181)</td>
<td>-0.0688*** (0.0212)</td>
<td>-0.0720*** (0.0219)</td>
</tr>
<tr>
<td>( \chi^2 )</td>
<td>8670.04*** (0.000)</td>
<td>7326.78*** (0.000)</td>
<td>24370.74*** (0.000)</td>
</tr>
</tbody>
</table>

The standard errors of the coefficients are given in parentheses. The \( \chi^2 \) pertains to the joint hypothesis that \( \alpha = 0 \) and \( \beta_1 = 1 \) (P values are given in parentheses).

*** Denotes rejection at the 1% level for the hypotheses that \( \alpha = 0 \) and \( \beta_1 = 1 \).

(B) Tests of rational expectations: \( E_{t} S_{t+k} - S_{t+k} = \alpha_1 + \beta_1 (F_{t+k} - S_t) + \epsilon_{t+k} \) from February 1, 1988 through May 1, 1992

<table>
<thead>
<tr>
<th></th>
<th>3 months</th>
<th>6 months</th>
<th>12 months</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \alpha_1 )</td>
<td>-0.0656** (0.0292)</td>
<td>-0.1607** (0.0683)</td>
<td>0.3077*** (0.1207)</td>
</tr>
<tr>
<td>( \beta_1 )</td>
<td>0.0772 (0.0611)</td>
<td>0.0996 (0.0950)</td>
<td>0.0544 (0.1400)</td>
</tr>
<tr>
<td>( \chi^2 )</td>
<td>5.76* (0.056)</td>
<td>6.30** (0.043)</td>
<td>7.12** (0.028)</td>
</tr>
</tbody>
</table>

The standard errors of the coefficients are given in parentheses. The \( \chi^2 \) pertains to the joint hypothesis that \( \alpha_1 = 0 \) and \( \beta_1 = 0 \).

* Denotes rejection at the 10% level for the hypotheses that \( \alpha = 0 \) and \( \beta_1 = 0 \).

** Denotes rejection at the 5% level for the hypotheses that \( \alpha = 0 \) and \( \beta_1 = 0 \).

*** Denotes rejection at the 1% level for the hypotheses that \( \alpha = 0 \) and \( \beta_1 = 0 \).

(C) Tests of perfect substitutability: \( E_{t} S_{t+k} - S_t = \alpha_2 + \beta_2 (F_{t+k} - S_t) + \epsilon_{t+k} \) from February 1, 1988 through May 1, 1992

<table>
<thead>
<tr>
<th></th>
<th>3 months</th>
<th>6 months</th>
<th>12 months</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \alpha_2 )</td>
<td>0.1118*** (0.0152)</td>
<td>0.2385*** (0.0300)</td>
<td>0.4692*** (0.0567)</td>
</tr>
<tr>
<td>( \beta_2 )</td>
<td>-0.1464*** (0.2945)</td>
<td>-0.1904*** (0.0450)</td>
<td>-0.2517*** (0.0580)</td>
</tr>
<tr>
<td>( \chi^2 )</td>
<td>1791.91*** (0.000)</td>
<td>1033.23*** (0.000)</td>
<td>663.77** (0.000)</td>
</tr>
</tbody>
</table>

The standard errors of the coefficients are given in parentheses. The \( \chi^2 \) pertains to the joint hypothesis that \( \alpha_2 = 0 \) and \( \beta_2 = 1 \) (P values are given in parentheses).

*** Denotes rejection at the 1% level for the hypotheses that \( \alpha_2 = 0 \) and \( \beta_2 = 1 \).

where [Eq. (5)]

\[
\beta_1 = \frac{\text{cov}(F_{t+k} - S_t, S_{t+k} - E_t S_{t+k})}{\text{var}(F_{t+k} - S_t)}.
\]

and

\[
\beta_1 = \frac{\text{cov}(F_{t+k} - S_t, u_{t+k})}{\text{var}(F_{t+k} - S_t)}.
\]

Under the hypothesis of rational expectations, \( \beta_1 \) will equal zero since the forecast error, \( u_{t+k} \), will be orthogonal to any variable in the set of information known to agents at the time they formed their expectations. Under the hypothesis that the correlation of the risk premium with the forward discount is zero (no time-varying risk premium), \( \beta_2 \) will equal one. In
Section 4, we consider formal tests along these lines. Also, we will consider alternative explanations that were suggested in the literature.

4. Decomposition of the bias: irrationality, exchange risk premia, peso problems, or other factors?

Survey data will be exploited in this section to decompose the forward discount bias into portions attributable to irrational behavior of economic agents or to the existence of time-varying risk premia. Here it is worth mentioning that the irrationality hypothesis is actually not the only explanation for a possible rejection of rationality of expectations. Other prominent explanations involve “peso problems” (see Krasker, 1980; and for learning about government policies, see Lewis, 1995).11 Furthermore, it should be pointed out that over the period analyzed, the Mexican peso/US dollar exchange rate was not freely floating but managed by the authorities and followed two different regimes: (1) from January 1988 to December 1988, the exchange rate was fixed except for a small change in February; and (2) from January 1989 through November 1991, a crawling peg was introduced: the exchange rate was depreciated by an announced 1 peso per US dollar a day, initially. Subsequently, the crawl was slowed to 0.4 pesos, and finally 0.2 pesos in November 1991. From then on, a fluctuation band was introduced, which gradually widened to four percentage points by the end of 1992.

To test for the rationality of the survey exchange rate expectations, we consider a fairly standard test (see Pesaran, 1987) --- the orthogonality test. The orthogonality test aims to assess whether economic agents use information that is available to them efficiently to forecast future exchange rates. The null hypothesis of rational expectations (orthogonality) implies that $\alpha_1 = 0$ and $\beta_1 = 0$ in regressions of the following form:

$$S_{t+k} - E_t S_{t+k} = \alpha_1 + \beta_1 (F_{t+k} - S_t) + v_{t+k}.$$  \hspace{1cm} (8)

where the left-hand-side variable is the survey forecast error. Under the null hypothesis of rational expectations and under the assumption that any measurement error in the survey is orthogonal to the forward discount, the $\beta_1$-coefficient is precisely equal to $\beta_1$ in Eq. (6). Eq. (8) was fitted via OLS for each forecast horizon; standard errors are corrected to allow for a $k - 1$ order moving average as in the estimation of Eq. (2).

Table 2B reports regressions of the forecast error on the 3-, 6-, and 12-month-ahead forward discount. Interestingly, the null hypothesis of rational expectations is rejected, but only in the sense that the survey expectation was off by a large and significant constant in all cases. The slope coefficient $\beta_1$ is never significantly different from zero. Note that the rejection of the rational expectations hypothesis could be due to the “peso problem” factor; i.e., expectations of the future spot exchange rate consistently overestimated the actual future.

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11 In particular, Krasker (1980) demonstrated that in the presence of a small and positive probability of a devaluation, an efficient exchange rate market will imply that the expected value of the future spot rate will reflect the probability of that event. However, as long as the devaluation does not take place within the sample period examined, the expectation of the future spot rate will consistently overestimate the realized future spot rate.
rate (i.e., negative survey forecast errors) in a regime where the exchange rate was not allowed to float freely.

In order to test whether the existence of time-varying risk premia is an important reason for rejection of forward market efficiency, we fitted the following equation \(^1\) 

\[
E_t S_{t+k} - S_t = \alpha_2 + \beta_2 (F_{t+k} - S_t) + \epsilon_t. 
\]  

(9)

The null hypothesis of perfect substitutability implies that \(\alpha_2 = 0\) and \(\beta_2 = 1\). Under the hypothesis that the correlation of the risk premium with the forward discount is zero (no time-varying risk premium), \(\beta_2\) will equal one. By inspection, the \(\beta_2\) coefficient is precisely equal to \(\beta_2\) in Eq. (7), reflecting a deviation from forward discount unbiasedness due to the existence of time-varying risk premia. Similarly, the hypothesis of a zero mean risk premium can be tested by examining whether the \(\alpha_2\) coefficient is significantly different from zero.

The results of fitting Eq. (9) for each forecast horizon are reported in Table 2C. The results indicate quite strongly that there is a large average risk premium, as well as a time-varying risk premium, component for all horizons. Since the risk premium is just equal to the interest rate differential less the expected rate of depreciation, a finding of \(\beta < 0\) implies that the risk premium on Mexican assets must rise with the interest differential and that the covariance of expected depreciation and risk premium is negative. Higher expected inflation in Mexico might sensibly be associated with both greater expected Mexican peso depreciation and increased riskiness of peso-denominated assets. This would be the case if, e.g., higher expected inflation reflects greater uncertainty about the future course of monetary policy.

The existence of time-varying risk premia corroborates some of the results of Cavaglia et al. (1994) for bilateral exchange rates relative to the US dollar and relative to the German mark spanning the same time period. By contrast, Froot and Frankel (1989) found estimates of \(\beta_2\) that were insignificantly different from one for survey-based tests using four of the major currencies relative to the US dollar, suggesting that changes in the forward discount primarily reflect changes in expected depreciation rather than changes in the risk premium. Interestingly, the greater ease of finding evidence of risk premia for smaller currencies is shown to be quite robust across different survey data sets.

5. Models of expectations formation

Empirical evidence demonstrates that the random walk hypothesis is still a relatively accurate characterization of the time series of exchanges rates of major industrialized countries (see the survey on nominal exchange rates by Frankel & Rose, 1995). Indeed, Meese and Rogoff (1983a, 1983b) and Wolff (1987) show that standard models of exchange

\(^1\) When the expected depreciation is on the left-hand side of the regressions, forecast horizons longer than the observational frequency do not themselves imply that the error term is serially correlated, since expectations are formed using only contemporaneous and past information. Therefore, Eqs. (9), (11), and (13) were estimated using the standard OLS procedure.
rate determination fail to outperform the predictive power of the random walk hypothesis even when allowing for time-varying model parameters. It is interesting to investigate how well the random walk performs for the peso/dollar exchange rate. The availability of survey data permits us to test directly how economic agents form their expectations of future appreciation of a currency. For instance, Allen and Taylor (1990) present survey-based evidence that foreign exchange dealers utilize some combination of charts and fundamentals in predicting currency movements, with greater weight being given to fundamentals as the forecast horizon lengthens. In this section, two alternative models of expectations formation are considered — the extrapolative and adaptive models — and a simple test regarding the term structure of expectations is presented. The extrapolative expectations model is considered first. Economic agents extrapolate the most recent trend into the future, formally [Eq. (10)]:

\[ \Delta E_t S_{t+k} = \beta (\Delta S_t). \] (10)

where \( \Delta E_t S_{t+k} \) is the most recent change in the expected exchange rate, \( E_t S_{t-k} - S_t \), and \( \Delta S_t \) is the most recent change in the spot exchange rate. If \( \beta \) is greater than zero, then exchange rate expectations are said to exhibit “bandwagon” effects, and if \( \beta \) equals zero, then expectations are said to be static. In the former case, a current appreciation generates expectations of further appreciation, whereas in the latter, the expected depreciation is equal to zero.\(^{13}\) If we define “speculation” as buying and selling of foreign exchange in response to non-zero expected exchange rate changes, one can interpret a finding of \( \beta > 0 \) (“bandwagon” expectations) as implying that speculation is destabilizing and a finding of \( \beta < 0 \) (distributed lag expectations) as implying that speculation is stabilizing. The following regression equation was fitted for each forecast horizon:

\[ E_t S_{t+k} - S_t = \alpha + \beta (S_t - S_{t-1}) + \epsilon_t. \] (11)

The results of fitting Eq. (11) are reported in Table 3A.

We find that the sign of the significant \( \beta \)-coefficients is positive in the regressions for each forecast horizon. Thus, past exchange rate depreciations are expected to be extrapolated in the future, as implied by destabilizing expectations models. Interpreting the regression coefficient for the Mexican peso/US dollar exchange rate at the 3-month horizon, a current depreciation of 10% in the Mexican peso implies an expected depreciation over the next 3 months of 12.1%. Acting on the basis of such expectations, investors are more likely to sell (buy) when the price of foreign exchange is already low (high), thereby exaggerating (dampening) the original depreciation. This “bandwagon” behavior can create speculative bubbles. This result is largely inconsistent with the empirical findings of Cavaglia et al. (1993a, 1993b), Frankel and Froot (1987a, 1987b), and MacDonald and Torrance (1988). They obtain parameter estimates suggesting stabilizing expectations models for long-term horizons. It should be noted that for the short-term expectations (1 week and 1 month), Frankel and Froot (1987a,

\(^{13}\) The existence of “bandwagon” exchange rate expectations has been a concern of critics of floating exchange rates in that it would render the system unstable (see Nurske, 1944). This view was challenged, however, by Friedman (1953), who argued that speculation would be stabilizing.
1987b), Froot and Ito (1989), and MacDonald and Torrance (1988) corroborate our findings of "bandwagon" effects, the destabilizing case.

Adaptive expectations models were subsequently considered; namely, the expected future spot rate is formed as a weighted average of the current spot rate and the lagged expected rate, or:

\[ E_t S_{t+k} = (1 - b)S_t + bE_{t-k}S_t. \]  

(12)

Alternatively, one can view the expected depreciation as a function of past forecast errors, and then the following equation may be fitted:

\[ E_t S_{t+k} - S_t = \alpha + \beta(E_t S_{t-1} - E_{t-k} S_t) + \epsilon_t. \]  

(13)

Eq. (13) corresponds to Eq. (12) if we set \( \alpha = 0 \) and \( \beta = -b \). The results of fitting the above equation for all forecast horizons are reported in Table 3B. Significantly positive coefficients are obtained for all, but one, forecast horizons. Interpreting the regression coefficient for the Mexican peso/US dollar exchange rate at the 3-month horizon, an unexpected depreciation of 10% in the peso implies an expectation of continued depreciation over the next 3 months of about 2.6%. Thus, long-term Mexican peso/US dollar exchange rate expectations — 3, 6, and 12 months ahead — again appear to be destabilizing significantly. These results are at variance with the results of Cavaglia et al. (1993a, 1993b) for the Japanese yen and EMS currencies relative to the US dollar and those obtained by Frankel and Froot (1987b) for the four most actively traded currencies. From Table 3A and B, it is clear that long-term Mexican peso/US dollar exchange rate expectations display a considerable positive weight on the contemporaneous spot rate, rather than placing all of the weight on the lagged spot rate or on the lagged expected spot rate, and in this sense are destabilizing.

Froot and Ito (1989) examine the consistency of short-run and long-run exchange rate expectations. They find that relative to long-term expectations, shorter-term expectations overreact to an exchange rate shock. We conduct a simple test of consistency of expectations by examining the extent to which shocks or revisions to 12-month-ahead expectations are reflected in short-run (3-month-ahead) expectations. Thus, we fitted the following model for the 3-month-ahead expected exchange rate:

\[ E_t S_{t+3} - S_t = \alpha + \beta(E_t S_{t+12} - E_{t-1} S_{t-11}) + \epsilon_t. \]  

(14)

Interpreting the above, a positive regression coefficient would suggest that a decline in the 12-months-ahead exchange rate forecast of the domestic currency would result in an expected depreciation 3 months ahead. A negative coefficient would suggest that short-run expectations "overreacted" to changes in 12-month expectations.

\[ ^{14} \text{Frankel and Froot (1987b) and MacDonald and Torrance (1988) obtain a regression coefficient for short-term expectations (1 week and 1 month ahead) that is opposite in sign to that of long-term expectations (3, 6, and 12 months ahead).} \]

\[ ^{15} \text{Because OLS estimates would be inconsistent in the context of Eq. (14), we implemented the instrumental variables estimation technique outlined in Hansen (1982). Instruments used were a constant term and lagged exchange rate returns.} \]
### Table 3

(A) Extrapolative expectations: $E_tS_{t+1} - S_t = \alpha - 3(S_t - S_{t-1}) - z_t$, from January 1, 1986 through May 1, 1992

<table>
<thead>
<tr>
<th></th>
<th>3 months</th>
<th>6 months</th>
<th>12 months</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\alpha$</td>
<td>0.0945*** (0.0126)</td>
<td>0.1893*** (0.0228)</td>
<td>0.3562*** (0.0396)</td>
</tr>
<tr>
<td>$\beta$</td>
<td>1.2082*** (0.2259)</td>
<td>1.9738*** (0.3882)</td>
<td>3.4149*** (0.6438)</td>
</tr>
<tr>
<td>$\chi^2$</td>
<td>297.53*** (0.000)</td>
<td>359.92*** (0.000)</td>
<td>386.70*** (0.000)</td>
</tr>
</tbody>
</table>

The standard errors of the coefficients are given in parentheses. The $\chi^2$ pertains to the joint hypothesis that $\alpha = 0$ and $\beta = 0$ ($P$ values are given in parentheses).

*** Denotes rejection at the 1% level for the hypotheses that $\alpha = 0$ and $\beta = 0$.

(B) Adaptive expectations: $E_tS_{t+1} - S_t = \alpha + \beta(S_t - E_t(S_t)) + z_t$, from January 1, 1986 through May 1, 1992

<table>
<thead>
<tr>
<th></th>
<th>3 months</th>
<th>6 months</th>
<th>12 months</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\alpha$</td>
<td>0.1258*** (0.0108)</td>
<td>0.2418*** (0.0213)</td>
<td>0.4720*** (0.0386)</td>
</tr>
<tr>
<td>$\beta$</td>
<td>0.2614*** (0.1125)</td>
<td>0.1877 (0.1246)</td>
<td>0.4055*** (0.0826)</td>
</tr>
<tr>
<td>$\chi^2$</td>
<td>142.71*** (0.000)</td>
<td>132.95*** (0.000)</td>
<td>149.54*** (0.000)</td>
</tr>
</tbody>
</table>

The standard errors of the coefficients are given in parentheses. The $\chi^2$ pertains to the joint hypothesis that $\alpha = 0$ and $\beta = 0$ ($P$ values are given in parentheses).

*** Denotes rejection at the 1% level for the hypotheses that $\alpha = 0$ and $\beta = 0$.

(C) Term structure of expectations. $E_t(S_{t+1} - S_t) = \alpha + \beta(E_t(S_{t+12} - E_t(S_{t+11})) + z_t$, from January 1, 1986 through May 1, 1992

<table>
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<tr>
<th></th>
<th>3 months</th>
<th>6 months</th>
<th>12 months</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\alpha$</td>
<td>0.0928*** (0.0327)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\beta$</td>
<td>2.0988** (0.9420)</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

The standard errors of the coefficients are given in parentheses.

** Denotes rejection at the 5% level for the hypotheses that $\alpha = 0$ and $\beta = 0$.

*** Denotes rejection at the 1% level for the hypotheses that $\alpha = 0$ and $\beta = 0$.

The results of fitting Eq. (14) to our survey data are reported in Table 3C. We find a significantly positive $\beta$-coefficient, suggesting that a positive innovation in the long-run fundamental (proxied by the 12-month-ahead exchange rate forecast) is associated with a positive change in the short-run (3-month) expectation by a larger amount. Our result is at variance with the Cavaglia et al. (1993a, 1993b) and Froot and Ito (1989) finding that exchange rate expectations are not consistent across the maturity spectrum.

To sum up, a number of studies on using survey data find that investors tended to react to current (unanticipated) appreciations by expecting future depreciation at horizons of 1 year, 6 months, or 3 months. At shorter horizons of 1 week to 1 month, however, investors tend to extrapolate recent trends. Expectations at these short horizons appear destabilizing. This lack of unity across different forecast horizons raises the possibility that different types of market participants form their expectations in different ways, with speculators more heavily represented at the short horizon and investors at the long horizons. Our results, however, demonstrate that the pattern across long-term forecast horizons for the Mexican peso/US dollar exchange rate contradicts the long-held concern that speculation based on long-term expectations may be stabilizing. Instead, it makes the
exchange rate too volatile around its long-run equilibrium, and the excess volatility imposes large costs on producers and consumers who, as a consequence, make less efficient allocative decisions.

6. Interest rate “news” and exchange rate surprises

In order to obtain a relationship between interest rate “news” and exchange rate surprises, it is useful to decompose the forecast error resulting from the use of the forward exchange rate as a predictor of the subsequent spot rate (at the maturity date of the forward contract) as follows:

\[ S_{t+k} - F_{t+k} = (S_{t+k} - E_t S_{t+k}) + (E_t S_{t+k} - E_t F_{t+k}). \]  \hspace{1cm} (15)

The forward rate forecast error, \( S_{t+k} - E_t F_{t+k} \), is often referred to as the “return to the forward speculation” or the “excess exchange rate return.” The above identity shows that the forecast error consists of two components: the surprise in the spot rate and the exchange rate risk premium.

In this section, we focus our attention on a number of different regression relationships that are closely linked to Eq. (15) and that can be estimated on the basis of our survey database that contains matched exchange rate and interest rate expectations. Following Frankel (1981), we focus on surprises in interest rate differentials as the most important source of unexpected exchange rate movements.\(^{16}\) We will investigate the sources of relationship (15) by means of the following three regression equations:

\[ S_{t+k} - F_{t+k} = \alpha + \beta_1 (d_{t+k} - E_t d_{t+k}) + \beta_2 (F_{t+k} - E_t S_{t+k}) + \epsilon_{t+k}. \]  \hspace{1cm} (16)

\[ S_{t+k} - F_{t+k} = \alpha + \beta_1 (d_{t+k} - E_t d_{t+k}) + \beta_2 (d_t) + \epsilon_{t+k}. \]  \hspace{1cm} (17)

\[ S_{t+k} - E_t S_{t+k} = \alpha + \beta_1 (d_{t+k} - E_t d_{t+k}) + \beta_2 (d_t) + \epsilon_{t+k}. \]  \hspace{1cm} (18)

Eq. (16) relates the forecast error resulting from the forward rate to “news” about interest differential and the level of the risk premium, which is also directly observable from our survey data. In Eq. (17), we replace the ex-ante measure of the risk premium by the lagged interest differential as a proxy for the risk premium, following Bekaert and Hodrick (1992). In Eq. (18), finally, we relate the innovation in the spot exchange rate to “news” about the interest differential and the level of the lagged interest differential.

In estimating Eqs. (16–18), the difficulty with applying the standard OLS procedure arises because of the contemporaneous variables appearing on the right-hand side of the equations. These will not, in general, be independent of the disturbance term.\(^{17}\) It therefore seems quite unsatisfactory to impose the exogeneity assumption. A general approach to estimation problems of this kind is provided by the method of instrumental variables. In this section,

\(^{16}\) Additionally, to distinguish between the flexible-price and sticky-price versions of monetary models, one can look at possible effects of innovations in interest rates on the price of foreign exchange.

\(^{17}\) This is usually referred to as “simultaneous equation bias.”
Table 4

(A) $S_{t+k} - E_t S_{t+k} = \alpha + \beta_1(d_{t+k} - E_t d_{t+k}) + \beta_2(F_{t+k} - E_t F_{t+k}) + z_{t+k}$ from February 1, 1988 through May 1, 1992

<table>
<thead>
<tr>
<th></th>
<th>3 months</th>
<th>6 months</th>
<th>12 months</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \alpha )</td>
<td>-60.5907 (36.1719)</td>
<td>-11.3658 (26.3266)</td>
<td>-13.3148 (11.6411)</td>
</tr>
<tr>
<td>( \beta_1 )</td>
<td>0.0877 (0.0755)</td>
<td>-0.0400 (0.0816)</td>
<td>-0.0337 (0.0445)</td>
</tr>
<tr>
<td>( \beta_2 )</td>
<td>-0.6371*** (0.2186)</td>
<td>-0.9298*** (0.1965)</td>
<td>-0.8781*** (0.1363)</td>
</tr>
</tbody>
</table>

The standard errors of the coefficients are given in parentheses.

*** Denotes rejection at the 1% level for the hypotheses that \( \alpha = 0 \), \( \beta_1 = 0 \), and \( \beta_2 = 0 \).

(B) $S_{t+k} - E_t S_{t+k} = \alpha - \beta_1(d_{t+k} - E_t d_{t+k}) - \beta_2(d_t) + \gamma_{t+k}$ from February 1, 1988 through May 1, 1992

<table>
<thead>
<tr>
<th></th>
<th>3 months</th>
<th>6 months</th>
<th>12 months</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \alpha )</td>
<td>-5.3372 (8.9474)</td>
<td>8.7621 (9.6477)</td>
<td>5.2884 (3.6856)</td>
</tr>
<tr>
<td>( \beta_1 )</td>
<td>0.0356**</td>
<td>-0.0052 (0.0245)</td>
<td>-0.0118 (0.0124)</td>
</tr>
<tr>
<td>( \beta_2 )</td>
<td>-0.7302*** (0.0371)</td>
<td>-0.6436*** (0.0393)</td>
<td>-0.4635*** (0.0155)</td>
</tr>
</tbody>
</table>

The standard errors of the coefficients are given in parentheses.

** Denotes rejection at the 5% level for the hypotheses that \( \alpha = 0 \), \( \beta_1 = 0 \), and \( \beta_2 = 0 \).

*** Denotes rejection at the 1% level for the hypotheses that \( \alpha = 0 \), \( \beta_1 = 0 \), and \( \beta_2 = 0 \).

(C) $S_{t+k} - E_t S_{t+k} = \alpha + \beta_1(d_{t+k} - E_t d_{t+k}) + \beta_2(d_t) - z_{t+k}$ from February 1, 1988 through May 1, 1992

<table>
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<th></th>
<th>3 months</th>
<th>6 months</th>
<th>12 months</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \alpha )</td>
<td>5.7424 (15.9884)</td>
<td>44.7977**</td>
<td>20.6176 (11.3996)</td>
</tr>
<tr>
<td>( \beta_1 )</td>
<td>-0.0467 (0.0391)</td>
<td>-0.1587*** (0.0569)</td>
<td>-0.1229*** (0.0390)</td>
</tr>
<tr>
<td>( \beta_2 )</td>
<td>-0.0341 (0.0673)</td>
<td>-0.242*** (0.0909)</td>
<td>-0.1248*** (0.0488)</td>
</tr>
</tbody>
</table>

The standard errors of the coefficients are given in parentheses.

* Denotes rejection at the % level for the hypotheses that \( \alpha = 0 \), \( \beta_1 = 0 \), and \( \beta_2 = 0 \).

** Denotes rejection at the 5% level for the hypotheses that \( \alpha = 0 \), \( \beta_1 = 0 \), and \( \beta_2 = 0 \)

*** Denotes rejection at the 1% level for the hypotheses that \( \alpha = 0 \), \( \beta_1 = 0 \), and \( \beta_2 = 0 \)

we implemented the instrumental variables estimation technique outlined in Hansen (1982). Some interesting results emerge from Table 4A, B, and C.

In Table 4A, we find that “news” about the interest differential enters insignificantly in the equation for the difference between the actual exchange rate and the lagged forward exchange rate: information about surprises with regard to the interest differential has no marginal predictive power. Furthermore, we find significant effects of the ex-ante measure of the risk premium for each forecast horizon.

As noted in Section 5, it is widely reported in the literature (see, for instance, Bekaert & Hodrick, 1992) that the lagged interest differential tends to predict movements in the excess return in the foreign exchange market, which are by definition equal to the difference between the spot rate and the lagged forward exchange rate. As a consequence, it is argued that the

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18 Instruments used were a constant term and lagged explanatory variables.
interest rate differential can serve as a proxy for the risk premium in the foreign exchange market. The significant effect of lagged interest differentials in equations for the excess return in the foreign exchange market is confirmed in Table 4B. The insignificant effect of the “news” about the interest differentials is very similar to the results reported in Table 4A; only for the 3-month forecast horizon do we find a significantly positive $\beta_3$-coefficient. As suggested by Frankel (1979), this is consistent with a monetary model of exchange rate determination in which a rise in domestic interest rates may be primarily due to inflationary expectations.

In Table 4C, we find that “news” about interest differentials enters significantly in the equations for the 6- and 12-month forecast horizons. Moreover, the significant coefficients are negative, reflecting that an unexpected rise in the interest differential tends to strengthen the domestic exchange rate, i.e., to reduce $S_{t+1}$. This result indicates that for the 6-month forecast horizon, a 1% unanticipated increase in the interest differential for the Mexican peso will lead to approximately 0.16% (unanticipated) appreciation of the Mexican peso, thereby exhibiting the Dornbusch overshooting effect. In order to sort out whether the significant effects of the interest differential in Table 4B really reflect time-varying risk premia, we estimated the same regression but this time with the difference between the realized spot rate and the expected future spot rate as the dependent variable. If the interest differential truly reflects risk premia, we would expect this variable to enter insignificantly in the equations for the difference between the actual exchange rate and the expected future spot rate in Table 4C. As is apparent from Table 4C, the lagged interest differential is highly significant for the 6- and 12-months horizons, but not for the 3-month horizon.

7. Conclusions

In this article, we have investigated expectations concerning the Mexican peso/US dollar exchange rate with the aid of a survey dataset containing market participants’ forecasts of the exchange rate and of the interest differential between the peso and the dollar. Our study complements previous work that has largely focused on analyzing survey data for the five most actively traded currencies vis-a-vis the US dollar, and on EMS currencies. Overall, the results indicate a sound rejection of the null hypothesis that the forward discount of the Mexican peso is an unbiased predictor of the future change in the Mexican peso/US dollar exchange rate. All estimated $\beta$-coefficients are negative, suggesting that an increase in the interest rate differential is associated with a decline in expected depreciation, since the Mexican peso subsequently appreciates on average.

Furthermore, our findings indicate that the survey expectations were off by a large and significant constant. At the same time, as in most models in which sterilized foreign exchange intervention is effective, variation in the forward discount does reflect a statistically significant degree of variation in the risk premium. Since the risk premium is just equal to the interest differential less the expected change in exchange rates, this implies that the risk premium on Mexican assets must rise with the interest differential and that the covariance of expected depreciation and risk premium is negative. This might be considered plausible: higher expected inflation in Mexico can be associated with both greater expected Mexican peso depreciation and increased riskiness of peso-denominated assets. The existence of time-
varying risk premia corroborates some of the results of Cavaglia et al. (1994) for bilateral exchange rates relative to the US dollar and relative to the German mark spanning the same time period. By contrast, Froot and Frankel (1989) provide evidence for survey-based tests using four of the major currencies relative to the US dollar that changes in the forward discount primarily reflect changes in expected depreciation rather than changes in the risk premium. Interestingly, the greater ease of finding evidence of risk premia for smaller currencies is shown to be quite robust across different survey data sets.

As to the expectations formation mechanism, market participants tended to react to current (unanticipated) depreciations by expecting future depreciations at the 3-, 6-, and 12-month horizons, as implied by destabilizing expectations models. Our results demonstrate that the pattern across long-term forecast horizons for the Mexican peso/US dollar exchange rate contradicts the long-held concern that speculation based on long-term expectations may be stabilizing. Instead, it makes the exchange rate too volatile around its long-run equilibrium, and the excess volatility imposes large costs on producers and consumers who, as a consequence, make less efficient allocative decisions. These results are at variance with the findings of Cavaglia et al. (1993a, 1993b) for the Japanese yen and EMS currencies relative to the US dollar and those obtained by Frankel and Froot (1987a, 1987b) for the four most actively traded currencies. They obtain parameter estimates suggesting stabilizing expectation models for long-term horizons. At shorter horizons of 1 week to 1 month, however, investors tend to extrapolate recent trends.

In addition, we have investigated empirically the relationship among exchange rate returns, “news,” and risk premia using survey data of matched Mexican peso/US dollar exchange and interest rate expectations, thereby at least partially avoiding the problem of artificially generated expectations when using an econometric technique. “News” about the interest differential does not contribute additional predictive power with regard to the peso/dollar exchange rate once a risk premium term is included in a regression equation. Interestingly, a Dornbusch-type overshooting effect is present in the Mexican data.

References


