Elections, Exchange Rates & Central Bank Reform in Latin America

by

Rodolfo Cermeño,* Robin Grier,** & Kevin Grier***

*División de Economía, Centro de Investigación y Docencia Económicas (CIDE), Carret. México-Toluca 3655, 01210 México D.F., Mexico. Email: rodolfo.cermeno@cide.edu.

**Corresponding Author: Department of Economics, University of Oklahoma, Norman OK 73019. Email: rgrier@ou.edu.

***Department of Economics, University of Oklahoma, Norman OK 73019. Email: angus@ou.edu.

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I. Introduction

The interaction of exchange rates and elections is an interesting example of how political economy models are important to understanding developing countries. This interaction is of particular interest in Latin America, where research has shown both that exchange rate pegs are more likely to be abandoned in the aftermath of elections and that the rate of depreciation is higher after elections.

In this paper, we focus on the link between depreciation rates and elections. Our contribution is twofold. First, we employ a statistical model that incorporates a wide range of economic variables as explanatory variables and also takes into account the pervasive conditional heteroskedasticity found in financial data. Second, we test whether the wave of central bank reform that swept the region has had any effect on the existence or strength of the electoral cycle in exchange rates. In a panel of 9 Latin American countries with available macroeconomic data and a history of exogenous election dates, we confirm previous findings that real depreciation intensifies after elections even when modeling the significant conditional heteroskedasticity in these data. We also show for the first time that post-election exchange rate movements are also less predictable.

We go on to test whether central bank reform has influenced the way in which elections affect the RER in Latin America. If reform has been effective at reducing political manipulation of the exchange rate, then any relationship we see between elections and the RER before central bank reform should disappear in the post-reform era. We find that the relationship does disappear after reform and that post-reform real exchange rates are also significantly less volatile.
We believe this work is important for three reasons. First, it shows that political exchange rate cycles in these Latin American countries are, for now at least, a thing of the past. Second, it presents evidence that central bank reform has had real effects, an idea that is controversial in the literature. Third, it demonstrates a useful modeling approach for panels that employ financial data.

Section II below reviews the literature on elections and the exchange rate, Section III sets out the model of the RER that we use in our analysis. Section IV contains our initial results while Section V incorporates and tests for the effects of central bank reform on the electoral cycle in real exchange rates. Section VI concludes.

II. Exchange Rates, and Elections in Latin America

In this section we consider how national elections can influence the evolution of the real exchange rate.¹ There are several theoretical models which imply that governments prefer to delay exchange rate devaluations before elections. Stein & Streb (1998) argue that policymakers postpone exchange rate corrections at least in part because of the increased inflation they often bring. Thus they shift inflation problems to the next period, presumably after the election takes place. Bonobo & Terra (2005) model how interest group competition between the tradable and non-tradable sectors can create uncertainty around elections. Non-tradable interests would prefer an overvalued exchange rate which would effectively lower the costs of imports, while tradable interests would prefer the opposite. A cycle of pre-election appreciation (and subsequent devaluations) may imply that incumbents are courting votes from the non-tradable sectors. Stein & Streb (2004), Stein et al. (2005) and Méon (2004) construct models where there is asymmetric information about the competency of the incumbent executive. Devaluations are a signal of
incompetence, so politicians are keen to postpone any exchange rate depreciations until after an election.

While some of the original evidence for politically influenced exchange rate cycles was anecdotal, recent research has shown a significant empirical relationship between elections and exchange rate fluctuations.\textsuperscript{2} Edwards (1994), in an empirical study of 39 large nominal devaluations in democratic countries, finds that devaluation is significantly more likely in the post-election period. Gavin & Perotti (1997) show that policymakers are more likely to abandon a fixed peg regime early on in their terms and not close to an election period.

In Latin America, Klein & Marion (1997) find in a sample from 1956-1991 that the likelihood of abandoning an exchange rate peg is highest right after a presidential election. Likewise, in a cross country study of 26 Latin American and Caribbean countries, Frieden et al. (2001) show that an upcoming election lowers the probability of a large real exchange rate devaluation (25 percent or greater) by more than 30 percent, while a changes in government raises the probability of a large devaluation by about 10 percent. In a sample of 15 Latin American countries from 1960 to 1994, Stein et al. (2005) present evidence of a significant real exchange rate depreciation in the months after a presidential election. Pascó-Fonte & Ghezzi (2001) and Bonomo & Terra (1999) confirm similar electoral cycles in the real exchange rate for Peru and Brazil, respectively.\textsuperscript{3}

In this paper we are interested in whether the finding that depreciations are delayed until after elections will continue to hold when we (1) embed the test in a generously parameterized statistical model that also controls for conditional heteroskedasticity and cross sectional dependence; and (2) allow a level shift in the conditional variance of the RER during post-
election periods to test for increased electoral uncertainty. While the effect of elections on exchange rate uncertainty is less studied, it is a potentially important phenomenon as exchange rate uncertainty has been linked to reduced exports and even reduced economic growth. An Alesina (1987) style rational partisan model implies that a contested election between political parties with different policy preferences creates uncertainty about the post-election inflation rate (and by extension the post-election real exchange rate). Grier & Grier (2000), using a simple univariate GARCH model, find exactly this result for Mexican inflation.

Once we undertake this initial test we turn to the question of whether central bank reform alters how elections affect real exchange rates and their predictability.

In the next three sections, we set out our statistical model of the real exchange rate and test whether the timing of elections and central bank reform play a significant role in the determination of real exchange rates in our 9 Latin American countries.

III. A Statistical Model of the Real Exchange Rate

A. Variables

Our sample is a monthly panel of 9 Latin American countries from 1980-2000, encompassing a total of forty presidential elections. The specific countries were chosen because they have relatively complete data and their elections are conducted at regular intervals and thus their timing is exogenous to economic variables. In this way we avoid issues of simultaneity between the economy and the election date.

Our real exchange rate data comes both from J.P. Morgan and the IMF’s International Financial Statistics CD-ROM. In these data, higher values of the index imply a higher real
value of the currency under study. Thus, real appreciations are denoted by increases in the index.

In order to test whether elections delay real exchange rate depreciations, we construct an electoral dummy variable called Post with data from Georgetown’s Political Database of the Americas, IFE’s election guide, and Bienen & van de Walle (1991). Post is equal to one for the month of the election and the subsequent five months. Appendix 1 provides a list of election dates for all the countries in the sample. Based on the results discussed above in the literature, we would expect to find a negative and significant coefficient on Post, reflecting the delayed real exchange rate depreciation that occurs after the election date.

We control for economic variables which affect the real exchange rate by adopting as our baseline model the empirical model of Goldfajn & Valdés (1999). They identify terms of trade, trade liberalization, government spending, and the international interest rate as important factors in the determination of the real exchange rate.

Shocks to the terms of trade, either through a fall in the price of exports or an increase in the price of imports, can have a negative income effect on small, open economies (see Diaz-Alejandro (1982)). For example, a rise in the price of imports can cause a reduction in a country’s permanent income, which would reduce people’s demand for non-tradables and cause a depreciation in the real exchange rate. On the other hand, Edwards (1989) identifies a possible substitution effect, in that production may move from the non-tradables to the tradables sector, resulting in an increase in the price of non-tradables and a real exchange rate appreciation.

The effect of a terms of trade shock thus depends on whether the substitution effect is stronger than the income effect. We use a terms of trade index from the World Bank’s World Tables, where 1987 is equal to 100. The data is yearly and is divided by 12 and interpolated.
using July as the base month.

Goldfajn & Valdés use economic openness (measured as the sum of exports and imports as a percentage of GDP) as a proxy for trade liberalization. They argue that a reduction in import tariffs would bring about a fall in the price of non-tradables in order to bring the labor market back into equilibrium. We measure openness as the ratio of exports and imports to GDP. The trade data is available monthly from the IMF IFS CD-ROM. GDP figures are interpolated from yearly data from the same source.

The effect of a permanent increase in the size of government can have two different effects on the price of non-tradables. Increased government spending can cause a real exchange rate appreciation if the government increases overall demand for non-tradable goods. If new government spending instead goes toward imported goods (Goldfajn & Valdés use the example of imported military equipment), then the increased expenditures will be associated with a real exchange rate depreciation. We measure the size of government with monthly data on general government expenditures as a percentage of GDP from the IMF-IFS CD-ROM.

We measure the international interest rate with the 3-month US Treasury bill rate in secondary markets, which is taken from the St. Louis Fed’s FRED database and is reported monthly over the sample. As Goldfajn & Valdés point out, changes in the international interest can have both short run and long-run effects on the real exchange rate. If we assume that savings are inelastic in the short run, then a fall in the international interest rate will translate into increased capital inflows into developing countries. In the long run however, when savings should be more elastic, the result of a lower international rate should mean a drop in net foreign assets.
Finally, we include lagged RER values on the right-hand side of our models to capture any persistence in the series. This is especially important in our case because un-modeled serial correlation can cause spurious findings of conditional heteroskedasticity. We also consider several lags of the economic control variables as nothing in the theory specifies the exact timing of the relationships. We want to let the baseline model fit the data as well as possible in order to make the test for the existence of effects from elections or CBR as stringent and credible as possible.

B. A Model of Conditional Heteroskedasticity in Panels

We seek to estimate how elections and central bank reform affect the RER (both its mean and its predictability) in a panel model that controls for the relevant economic variables and allows for a time-varying covariance matrix. In order to determine an appropriate functional form of the model, we conduct several pre-tests. Specifically, we implement panel unit root tests on the different variables discussed above, finding in each case that the series are not stationary.\textsuperscript{10} We next investigate any possible co-integrating relationships by means of panel unit root tests, of which we found none. These results indicate that the data should be first-differenced and the model estimated in growth rates.

We tested for the existence of individual effects in both the conditional mean and the conditional variance and found pervasive and significant individual effects in the mean and the variance. The differencing described above will remove any individual effect from the mean equation and we will include country specific fixed effects in our conditional variance model.\textsuperscript{11} In addition, we tested for the best fitting lag structure of the independent variables and found that 12 lags of RER growth, and 3-month moving averages of terms of trade, openness, and
government spending are appropriate. The U.S. t-bill rate is contemporaneously related to the RER in these countries and we interpret this correlation as one way causation from the US to Latin America. Our model allows for both conditional and unconditional heteroskedasticity and conditional cross-sectional correlation of the error terms. Equations 1 - 3 below present the specifics.

\[
\Delta \ln(R_{it}) = \alpha_0 + \sum_{j=1}^{12} \alpha_j \Delta \ln(R_{it-j}) + \alpha_5 \Delta \ln(\text{totma}_i) + \alpha_6 \Delta \ln(\text{openma}_i) \\
+ \alpha_7 \Delta \ln(\text{govma}_i) + \alpha_8 \Delta \ln(\text{tbill}_i) + \mu \text{Post}_i + \varepsilon_{it}
\]

\[
h_{iit} = \phi_1 \varepsilon_{iit-1}^2 + \phi_2 h_{iit-1} + \phi_3 B_{r_i} + \phi_4 C_{h_i} + \phi_5 C_{o_i} + \phi_6 C_{r_i} + \phi_7 E_{c_i} + \phi_8 M_{ex_i} \\
+ \phi_9 P_{e_i} + \phi_{10} U_{r_i} + \phi_{11} V_{e_i} + \theta \text{Post}_i \\
\text{For all } i = 1 - N
\]

\[
h_{ikt} = \rho_{ik} h_{iit} h_{kkt} \\
\text{For all } i \neq k
\]

Here i indexes countries and t indexes time. The error terms are assumed to be distributed multivariate normal with mean zero and variance $H_t$. The diagonal elements of $H_t$ are given in equation 2 and the off-diagonal elements in equation 3. The covariance specification follows Bollerslev (1990).

The key coefficients testing for electoral effects on the RER process are $\mu$ and $\theta$. If politicians purposefully delay putting off needed real exchange rate depreciations until after elections, then $\mu$ will be negative and significant. If elections create additional RER uncertainty, then $\theta$ will be positive and significant.
IV. Results

Table 1 presents some preliminary results of estimating equation (1) via Least Squares. The coefficient on Post is negative, sizeable (around -14) and significant at the 0.01 level, supporting the argument that politicians seek to delay potentially politically costly real exchange rate depreciations until after an election.

The terms of trade variable is negative, which indicates that improvements in the terms of trade are associated with decreases in the real exchange rate, but is only weakly statistically significant. The openness variable is negative and significant at the .05 level, meaning that greater levels of openness are correlated with lower real exchange rate values. Changes in the t-bill rate are positively and significantly correlated with real exchange rate appreciations, while government spending does not have a significant effect in this specification.

These initial results confirm and extend what others have argued about how elections can distort the evolution of exchange rates. Even when considering the real exchange rate, exogenous election dates, and a statistical model for the evolution of the RER we find that there is significant post-electoral RER depreciation. We are interested though in whether these OLS residuals show evidence of conditional heteroskedasticity and whether change in the legal status of a nation’s central bank affects the finding of post-electoral RER depreciation.

To formally test for conditional heteroskedasticity we take the squared residuals from Table 1 and regress them on various numbers of lags of the squared residuals. The $R^2$ of these regressions multiplied by the sample size is asymptotically distributed as a $\chi^2$ statistic which we can use to test the null hypothesis of no autocorrelation in the squared residuals. At one lag, the value of the $\chi^2$ statistic is 16.26, which is significant at the 0.001 level. At five (ten) lags, the
calculated $\chi^2$ is 30.37 (41.46), also significant at the 0.001 level. Thus, there is evidence of significant and persistent autocorrelation in these squared residuals.

In Table 2, we control for conditional heteroskedasticity and cross sectional dependence by estimating the system of equations 1, 2 and 3 given above via direct numerical maximization of the log likelihood.$^{15}$ To save space, we do not individually report the estimated coefficients on the 12 lags of the RER, but rather simply report their joint significance level, which is 0.01.$^{16}$ To test the significance of our estimated covariance model, we can compare the maximized value of the likelihood function here to that obtained in Table 1. This likelihood ratio test yields a calculated Chi-square statistic of 1,666, which means that we can reject at the .001 level the null hypothesis that there is no conditional heteroskedasticity or cross-sectional dependence in these data.$^{17}$

Table 2 shows that once we model the conditional heteroskedasticity and cross-sectional dependence in the residuals, the size of the electoral effect in the mean equation drops by around 50% and its significance level falls to 0.05. We also find a positive relationship between the post-election dummy and the conditional variance. $Post$ has a large positive and significant effect on the conditional variance of RER growth (the coefficient is around 1100 which is larger than 8 of the 9 country fixed effects in the variance). Thus by implementing our covariance model we find that the effect of elections on depreciation is still significant and is also accompanied by a large increase in uncertainty associated with the aftermath of elections.$^{18}$

The lower portion of the table contains the estimated cross-country conditional correlation matrix. Twenty-one of the 36 correlations are positive and significant at the 0.05 level. The countries whose exchange rates shocks are most significantly linked to other countries
are Venezuela and Chile, while the countries with the fewest significant links are Uruguay and Ecuador.

As one simple check on the adequacy of our covariance model, we take the normalized residuals from Table 2 and test their squares for autocorrelation. At 1, 5, and 10 lags the calculated $\chi^2$ statistics are 0.23, 1.81 and 2.53 respectively, all of which are statistically insignificant, indicating that the autocorrelation of the squared residuals found in the OLS model has been eliminated.\(^{19}\)

These results confirm the basic finding of significant electorally influenced depreciations in Latin America in a statistical model that controls for a range of economic variables and conditional heteroskedasticity. Further, they show for the first time that there is a second dimension to the phenomenon, namely a significant increase in uncertainty in the aftermath of elections. We now proceed to investigate whether the wave of Central Bank reform that swept the region has had any effect on this electoral exchange rate cycle.

V. The Effects of Central Bank Reform

Between 1988 and 1995, as documented in Jácome (2001), each of the nine countries we study has enacted substantial central bank reform legislation granting increased independence to the central bank. Specifically, Jácome (2001) shows that significant increases in central bank autonomy were accomplished by reforms that (a) changed central bank charters so that the sole (or at least primary) objective would be price stability, (b) reduced the dependence of central banks on the executive branch of government, and (c) increased the economic autonomy of central banks.\(^{20}\) Appendix 1 provides the dates of the reforms in each country.
To test for the effect of these changes on the RER process, we create a variable $Reform_{it}$, which is equal to 0 before reform was undertaken in country $i$ and equal to 1 after that date. In addition, we interact this variable with $Post$ to investigate whether the CBR has significantly changed the relationship between the political business cycle and the mean and conditional variance of the real exchange rate in Latin America. This changes equations (1) and (2) to the following:

\[
\Delta \ln(R_{it}) = \alpha_0 + \sum_{j=1}^{12} \alpha_j \Delta \ln(R_{jt-1}) + \alpha_5 \Delta \ln(totma_{it}) + \alpha_6 \Delta \ln(openma_{it}) + \alpha_7 \Delta \ln(govma_{it}) + \alpha_8 \Delta \ln(tbill_{it}) + \mu Post_{it} + \beta Reform_{it} + \Delta Post_{it} \times Reform_{it} + \epsilon_{it}
\]

\[
h_{it} = \phi_1 \epsilon_{it-1}^2 + \phi_2 h_{it-1} + \phi_3 Br_i + \phi_4 Ch_i + \phi_5 Col_i + \phi_6 Cr_i + \phi_7 Ec_i + \phi_8 Mex_i + \phi_9 Per_i + \phi_{10} Ur_i + \phi_{11} Ven_i + \theta Post_{it} + \chi Reform_{it} + \psi Post_{it} \times Reform_{it} \quad \text{For all } i = 1 - N
\]

With this model we can test whether the timing of CBR is significantly partially correlated with direct changes in the RER process and whether it has any influence on how elections affect the RER process. If central bank reform has been an effective and credible means for Latin American governments to reduce monetary manipulation, then we would expect to see any pre-reform relationship between RER depreciation and elections to be diminished in the conditional mean equation.

Table 3 reports estimates of equations 1', 2' and 3. Again we do not report the 12 individual RER lag coefficients, but simply note that they are significant as a group at the 0.01 level. The four new variables that did not appear in Table 2 ($Reform$ and $Post \times Reform$ in both the conditional mean and conditional variance equations) are jointly significant at the 0.01 level.
and three of them are individually significant. Thus, CBR is an important factor for understanding the evolution of the real exchange rate process in these countries. The Goldfajn-Valdes control variables are jointly significant and generally show greater individual statistical significance than they did in the OLS specification reported in Table 1.

Table 3 shows the following results: (1) ceteris paribus, CBR is associated with a lower average rate of RER depreciation and significantly less RER uncertainty. The reform variable is positive (negative) and significant in the conditional mean (variance) equation, meaning that there is less average depreciation and that the RER process is less unpredictable in the post-reform period; (2) the immediate post-election period is associated with increased rates of RER depreciation in the pre-reform data, but this effect is completely erased in the post-reform observations. Pre-reform, the RER depreciates almost 13% per month in the post-election period, holding other factors constant, while post-reform there is, on net, a small, but not statistically significant, post-election appreciation effect; and (3) once the change in the effect of elections on RER growth is allowed to change in the pre- and post-reform periods, there is no longer any evidence that elections create additional RER uncertainty. Both Post and (Post*Reform) are insignificant in the conditional variance equation.\(^{25}\)

In sum, our results clearly show that CBR is correlated with the real exchange rate process in these nine countries. This is true both directly and in terms of how elections affect the real exchange rate. CBR has on average eliminated electorally motivated RER depreciations, reduced the average rate of RER depreciation and lowered RER uncertainty.
VI. Financial Liberalization as an Alternative Explanation

While the above results are striking, one can never be completely sure that the set of dates associated with CBR are the best set of dates for explaining the demise of the RER election cycle in these countries. Given the amount of time required to estimate a single panel model like the ones in this paper (several hours), it is not feasible to employ a strategy of looking at all possible combinations of break dates. Here though, we consider one alternative explanation, namely that financial openness may have ended the RER election cycle.

To this end we create two new variables: Liberalize and the interaction of Liberalize and Post. Liberalize is a dummy variable equal to 1 after a country has liberalized its capital markets. The data is compiled from work by Campbell et al. (2002, 2005) and summarized in Campbell & Bekaert (2004). Appendix 1 provides the dates of liberalization for each country.

Table 4 reports the results of re-estimating equations 1', 2' and 3' with the inclusion of these new variables. As with previous results, we do not report the separate RER lag coefficients, but they are significant as a group at the 0.01 level. The model has a lot of coefficients but the main result can be summarized succinctly: it is CBR and not financial opening that is significantly correlated with the decline of the RER election cycle in these countries. Specifically, the model shows that in the post-election period in a country with no financial opening and no central bank reform, the RER depreciates around 9 percentage points. After elections where the country has financial opening but no CBR, the RER depreciates around 21 percentage points. After elections where there is both financial opening and CBR the RER does not depreciate at all. If anything, then, we see that controlling for CBR, financial opening
exacerbates the electoral cycle in the RER. While not conclusive, this exercise lends credence to our claim that CBR has been an effective policy tool in these countries.

VII. Discussion

In this paper we investigate the interaction between exchange rate depreciations and elections. Our work extends existing research on the subject in two ways: (1) we use a generously parameterized statistical model which both allows for conditional heteroskedasticity and cross-sectional correlation in the errors and allows us to test for electoral effects on exchange rate uncertainty; and (2) we investigate the role of CBR on the relationship between elections and RER depreciation and the evolution of the RER in general.

We find that real depreciations were significantly delayed until after elections in the pre-CBR data and that the adoption of new CB legislation completely eliminates the effect. We also find that CBR is significantly partially correlated with reduced RER uncertainty and lower average depreciation rates. Further, we show that the onset of CBR is more closely related to the disappearance of the RER electoral cycle than are the date of financial opening in these countries. This is new evidence in favor of the proposition that legal central bank independence can have real economic effects and suggests a policy recommendation that central bank reform accompany democratic reforms in order to mitigate any tendency for elections to create exchange rate instability.
Table 1. A Least Squares Model of RER Growth

\[
\Delta \ln(R_{it}) = 1.63 + .06 \Delta \ln(R_{it-1}) - .12 \Delta \ln(R_{it-2}) - .03 \Delta \ln(R_{it-3}) - .08 \Delta \ln(R_{it-4}) \\
(1.2) \quad (3.1) \quad (5.6) \quad (1.4) \quad (3.8)
\]

+ .01 \Delta \ln(R_{it-5}) - .04 \Delta \ln(R_{it-6}) - .05 \Delta \ln(R_{it-7}) - .02 \Delta \ln(R_{it-8}) \\
(0.6) \quad (2.1) \quad (2.5) \quad (0.8)

+ .01 \Delta \ln(R_{it-9}) + .02 \Delta \ln(R_{it-10}) - .02 \Delta \ln(R_{it-11}) - .004 \Delta \ln(R_{it-12}) \\
(0.5) \quad (1.1) \quad (1.1) \quad (0.2)

- .14 \Delta \ln(tot_{ma}) + .01 \Delta \ln(gov_{ma}) - .04 \Delta \ln(open_{ma}) + 0.06 \Delta \ln(tbill) \\
(1.5) \quad (1.0) \quad (2.0) \quad (3.6)

- 13.9 \text{ Post} \\
(3.3)

\[\text{LLF} = -12396; \text{N=9, T=247}\]
Table 2. A Model of RER Growth, Controlling for Conditional Heteroskedasticity & Cross-Sectional Dependence

\[ \Delta \ln(R_{it}) = 0.67 + \sum_{j=1}^{12} \alpha_j \Delta \ln(R_{it-j}) + 0.09 \Delta \ln(\text{tot}_{ma}) - 0.009 \Delta \ln(\text{open}_{ma}) + 0.011 \Delta \ln(\text{gov}_{ma}) \]
\[ + 0.06 \Delta \ln(\text{tbill}) - 6.91 \text{Post} \]

\[ h_{iit} = 1.6 \varepsilon_{iit-1}^2 + 0.13 h_{iit-1} + 1187.3 \text{Post} + 430.9 \text{Br} + 334.5 \text{Ch} + 188.8 \text{Col} + 74.5 \text{Cr} + 2690.9 \text{Ec} + 220.1 \text{Mex} + 273.2 \text{Per} + 404.3 \text{Ur} + 247.0 \text{Ven} \]

\[ h_{kt} = \rho_{ik} * h_{iit} * h_{kkt} \]

Matrix of the estimated \( \rho_{ik} \)

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Maximized Log-Likelihood = - 10909  N=9  T=247

Numbers in parentheses are asymptotic t-statistics, while the number in brackets is the marginal significance level of the 12 lagged RER growth coefficients (which sum to 0.10).

Correlation coefficients in bold are significant at the .05 level or better.
Table 3: The Effect of Central Bank Reform on the Election-RER relationship

\[
\Delta \ln(R_{it}) = -5.8 + \sum_{j=1}^{12} \alpha_j \Delta \ln(R_{it-j}) + 0.11 \Delta \ln(tot_{ma}) - 0.017 \Delta \ln(open_{ma}) + 0.010 \Delta \ln(gov_{ma}) \\
+ 0.06 \Delta \ln(tbill) - 9.82 Post + 15.4 Post*Reform + 4.22 Reform \\
\]

\[
h_{it} = 1.6 e_{it-1}^2 + 0.13 h_{it-1} + 446.1 Post + 702.5 Post*Reform - 88.5 Reform + 421.7 Br \\
+ 386.2 Ch + 213.8 Col + 121.1 Cr + 2828.7 Ec + 279.6 Mex + 285.9 Pe \\
+ 477.3 Ur + 248.7 Ven \\
\]

\[
h_{ikt} = \rho_{ik} * h_{it} * h_{kt} \\
\]

Matrix of the estimated \( \rho_{ik} \)

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Maximized Log-Likelihood = -10889  N=9  T=247

Numbers in parentheses are asymptotic t-statistics while, the number in brackets is the marginal significance level of the 12 lagged RER growth coefficients (which sum to 0.06).

Correlation coefficients in bold are significant at the .05 level or better.
Table 4: The Effect of Central Bank Reform & Financial Liberalization on the Election-RER relationship

\[
\Delta \ln(R_{it}) = 0.16 + \sum_{r=1}^{12} \alpha_j \Delta \ln(R_{it-j}) + 0.13 \Delta \ln(\text{tot}_{ma}) - 0.027 \Delta \ln(\text{open}_{ma}) + 0.010 \Delta \ln(\text{gov}_{ma}) \\
+ 0.06 \Delta \ln(\text{tbill}) - 9.35 \text{Post} + 20.4 \text{Post*Reform} + 1.56 \text{Reform} \\
+ 2.38 \text{Liberalize} - 12.5 \text{Post*Liberalize}
\]

\[
h_{it} = 1.6 e_{it-1}^2 + 0.12 h_{it-1} + 53.7 \text{Post} - 97.4 \text{Post*Reform} - 216.1 \text{Reform} - 68.9 \text{Liberalize} \\
+ 1790.27 \text{Post*Liberalize} + 537.0 \text{Br} + 509.3 \text{Ch} + 416.2 \text{Col} + 326.9 \text{Cr} \\
+ 3002.6 \text{Ec} + 442.6 \text{Mex} + 426.7 \text{Pe} + 783.2 \text{Ur} + 436.7 \text{Ven}
\]

\[
h_{ikt} = \rho_{ik} * h_{it} * h_{kt}
\]

Matrix of the estimated \( \rho_{ik} \)

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Maximized Log-Likelihood = -10865  \( N=9 \)  \( T=247 \)

Numbers in parentheses are asymptotic t-statistics while, the number in brackets is the marginal significance level of the 12 lagged RER growth coefficients (which sum to 0.06).

Correlation coefficients in bold are significant at the .05 level or better.
### Appendix 1: Presidential Elections, Central Bank Reform & Financial Liberalization

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<th>Liberalization</th>
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References


Edwards, Sebastian, 1989, Temporary terms of trade disturbances, the real exchange rate, and the current account, *Economica* 56 (223), 343-357.


Georgetown Election Data (http://www.georgetown.edu/pdba/Elecdata/Calendar)


Harvey, Campbell and Geert Bekaert, 2004, A detailed chronology of important financial, economic, and political events in emerging markets. (http://www.duke.edu/~charvey/Country_risk/couindex.htm).

IFE’s election guide (http://www.ifes.org/eguide/elecguide.htm).


Given that there is not a one-to-one, nor even necessarily a monotonic, relationship between nominal and real exchange rate fluctuations, studying only the nominal exchange rate can give misleading inferences about whether politics affects the economically more important real exchange rate. A nominal depreciation is not the only way a government can lower its real exchange rate. It could also keep the domestic inflation rate consistently lower than the inflation rate of its trading partners by restricting the money supply. Similar to a nominal devaluation though, deflation before an election would also be politically painful.

Cooper (1971) argues that nominal devaluations are costly to developing country governments, especially finance ministers, who often are fired right after a devaluation takes place. Ben-Porath (1975) shows that the Israeli government has never devalued its currency less than eighteen months before an election.

On the other hand, Leblang (2003) finds no evidence for increased post-election devaluations in a large panel study. In his data, the number of attacks is higher post-election than pre-election, but governments seem to consistently defend against these attacks in the post-election period.

For example, Grier & Hernández (2004) show that exchange rate uncertainty has a negative and significant effect on the growth rate of Mexican industrial production, while Grier & Smallwood (forthcoming 2007) show that real exchange rate uncertainty negatively effects developing country exports.

Further, given the surprising frequency with which candidates run on one platform only to deliver another in Latin America, it is intuitively plausible that elections might generate increased uncertainty about the time path of macro variables. Carlos Menem in Argentina, Alberto Fujimori in Peru, and Fernando Henrique Cardoso in Brazil are recent examples of this phenomenon.

Several of the countries in the sample have experienced a transition from military to civilian rule during the 1980-2000 time period, which could potentially create a situation of reverse causation. That is, a military government may be handing back the reins of government to civilians because the economy is in a bad state. Haggard & Kaufman (1995) point to the initial democratic elections in Brazil and Uruguay as two cases where economic problems may have sped up the process of democratization. In results not reported in this paper, we investigated this possibility by creating a separate Post dummy for these two cases and found that it (the supplemental dummy) was insignificant in all specifications.

For 7 of the 9 countries (all but Costa Rica and Uruguay) in the sample, we use monthly trade-weighted real exchange rate data from J.P. Morgan www.jpmorgan.com/MarketDataInd/Forex/currIndex.html (Data retrieved 10/01/02). We were
able to add Costa Rica and Uruguay using the IMF-IFS CD-ROM. The correlation coefficient between the two RER measures for countries which appear in both data sets ranges from .94 to .99.

8. In the case of run-off elections, Post is equal to one for the month of the run-off and the subsequent five months.

9. They also point out that “this result depends on the assumption that the cross price elasticities of excess demand of non-tradables with respect to both exportables and importables are positive.”

10. Details of these tests are available on request. When we consider the data country by country using traditional unit root tests, the results are overwhelmingly similar to the panel test results.

11. To pre-test for individual effects in the conditional variance, we take the squared OLS residuals, construct estimated individual error variances for each country and then test the null hypothesis that all the error variances are equal.

12. The 3 month moving average includes the second through fourth lags of the variables.

13. That is to say, we assume that the U.S. t-bill rate is uncorrelated with the error term in the exchange rate equation.

14. This type of panel-garch model was first proposed by Cermeño & Grier (2005).

15. Engle (1982) demonstrates in the univariate case, that while least squares is still the best linear estimator in this setting, it can be extremely inefficient compared to the non-linear estimator accounting for the conditional heteroskedasticity.

16. In these data, RER growth is not very persistent, as indicated by the small sum of the lag coefficients. The initial values tend to be negative and significant while the later values are generally positive and significant.

17. The critical value at the .005 level with 9 degrees of freedom is 21.96.

18. Grier & Hernandez (2004) show for the case of Mexico that real exchange rate uncertainty has a significant negative partial correlation with the growth of industrial production.

19. Recall that the OLS specification showed that 12 lags were sufficient to deal with correlation in the level of the residuals but that there was very strong correlation in the squared OLS residuals.

20. Central banks in Brazil, Chile, Colombia, Costa Rica, Mexico, and Peru now have price stability as their main objective. Executive appointments to the central banks of Brazil, Chile, Colombia, Costa Rica, Mexico, and Venezuela must now be confirmed by the legislature. In addition, reforms in Chile, Mexico, Peru, and Venezuela have all outlawed central bank credit to the government except in cases of emergency.
21. The important date is when the reforms started to have effects, which could potentially be either before (through expectations) or after (through inertia) the enactment date. In the absence of any outside information, we take the enactment date as the relevant date. If we are systematically wrong, our reform variable will be biased against finding any effects of reform.

22. The effect of CBR on RER uncertainty is less clear cut as it revolves around the relative predictability of governments vs. central bankers as well as the credibility of the undertaken reform. If monetary policy is more predictable under an independent central bank than it is under direct government control, then CBR should lower uncertainty. If, however, the reform is not fully credible, agents’ expectations will be a weighted average of the outcomes predicted for each policymaker (that is, the independent central bank and the government), and the variance of the forecast errors may well increase. A finding that CBR reduces uncertainty would be evidence in favor of the idea that, on average, Latin American CBR was credible and that central bank policies are generally more predictable than the government’s exchange rate policies.

23. Franzese (1999, 2003) uses a model where all the independent variables would be interacted with the reform variable. In our case, this is not computationally feasible given the number of extra coefficients involved and the difficulty of estimating large numbers of coefficients using nonlinear methods, especially given that our panel approach greatly exacerbates these problems. Our advantage is that we can investigate the effect of reform on uncertainty.

24. Swinburne & Castello-Branco (1991) argue that laws which purport to safeguard CB independence may be useless if there are other ways (that is, non-statutory ways) in which politicians can influence the central bank. If this is the case in our sample, then we should see no difference between the pre- and post-reform period.

25. We do not emphasize the exact quantitative effect of CBR because we are dealing with intercept shifts and it is unwise to interpret the intercept literally in a linear regression. However, it is the case that the intercept of the mean equation shifts from negative and significant to positive and significant as does the post-election coefficient. The intercept of the conditional variance equation declines by a bit more than 40%.

26. It is worth noting that if we leave out CBR, we find that financial opening reduces the RER election cycle, though the model with CBR only fits the data marginally better than the model with financial opening only.