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Elections, exchange rates and reform in Latin America☆

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ABSTRACT

In this paper, we study the link between real exchange rate (RER) depreciation and elections in Latin America. Our contribution is threefold. First, we employ a statistical model that takes into account the pervasive conditional heteroskedasticity found in financial data and includes a wide range of macroeconomic variables as regressors. 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. Third, we test an additional hypothesis, namely, that financial liberalization may also be an important variable explaining changes in electoral effects on the real exchange rate. In a panel of 9 Latin American countries with available macroeconomic data and a history of exogenous election dates, we confirm the 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 in the literature, that post-election exchange rates are significantly 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 be mitigated in the post-reform era. We find that the relationship disappears after reform and that post-reform real exchange rates are also significantly less volatile. Finally, we show that financial liberalization seems to have a stronger effect on the conditional variance of the RER than does central bank reform, but reform has a stronger impact on the conditional mean.

1. Introduction

Political economy models are important to understanding the behavior of real world economic policymakers, especially in developing countries. In one interesting example, research has shown that, in Latin America, exchange rate pegs are more likely to be abandoned in the aftermath of elections and that the rate of both nominal and real depreciation is higher after elections (e.g., Stein et al., 2005).

In this paper, we study the link between depreciation rates and elections in Latin America in several novel ways. First, we employ a statistical model that takes into account the pervasive conditional heteroskedasticity often found in financial data. This innovation is technically important because un-modeled conditional heteroskedasticity can cause estimated regression coefficients to be quite inefficient and hence produce misleading inferences. It is also substantively important because, for the first time in the literature, it allows us to test whether elections affect the predictability of the real exchange rate (hereafter, RER).[1]

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 effect on exchange rates. Since these reforms (which we discuss in detail in Section 5) are designed to insulate the Central Bank from politics, we expect that they might well reduce any electorally induced movements in the real exchange rate. However, if they are

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[1] We use the words unpredictability and uncertainty rather than volatility in the paper because we are trying to distinguish between possibly predictable movements (volatility) and unpredictable movements (at least unpredictable given the model) which capture uncertainty. We are estimating the conditional variance of the error terms of the model so in that sense we are measuring uncertainty about any forecast using the model. This variance does depend on observed outcomes, and that is why it is called the conditional error variance, but it is a good measure of the unpredictability of the conditional mean in the next period. In practice, the term volatility has come to mean unpredictable variation in financial econometrics, but the distinction is still worth making. A moving standard deviation around the mean measures volatility but not necessarily uncertainty if the movements are predictable. The volatility of temperatures around an annual average is much greater than the actual unpredictability of temperatures. The variance of the error term in regression has always been a measure of the unpredictability of a variable, the only thing different here is that we are modeling systematic movements in the error variance.

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nothing more than "window dressing", they may have no effect on the process.

Third, we test an additional hypothesis, namely, that financial liberalization may also be an important variable explaining changes in electoral effects on the real exchange rate. As we argue in Section 6, to the extent that financial liberalization increases the economic costs of political manipulations (by allowing capital to flow out of the country in response), it may be a dampening factor on electorally induced exchange rate movements.

In a panel of nine Latin American countries with available macroeconomic data and a history of exogenous election dates, we show that real depreciation intensifies after elections even after controlling for relevant economic variables and modeling the significant conditional heteroskedasticity in these data. We also show for the first time in the literature that post-election exchange rate movements are also significantly less predictable.

We further find that the relationship between elections and real exchange rate depreciation disappears after central bank reform and that post-reform real exchange rates are also significantly less volatile. Finally, we show that when we add financial liberalization to the model, it seems to have a stronger effect on the conditional variance of the RER than does central bank reform, while reform has a stronger impact on the conditional mean.[2]

In what follows below, Section 2 reviews the literature on elections and the exchange rate, Section 3 sets out the model of the RER that we use in our analysis. Section 4 contains our initial results while Section 5 incorporates and tests for the effects of central bank reform on the electoral cycle in real exchange rates. In Section 6, we test if financial liberalization can explain the end of the RER election cycle in Latin America. Section 7 concludes.

2. Exchange rates and elections

There are two main approaches in the literature on exchange rate and elections.[3] The first one uses political budget cycle models where the incumbent government uses economic policy to enhance its reelection chances. The second approach, which is more empirically oriented, studies whether politically costly depreciations are delayed until after an election. While early papers focused on the nominal exchange rate, the literature has increasingly emphasized movements in the real exchange rate.

2.1. Political budget cycle approach

In Stein and Streb (1998), the public dislikes inflation because it erodes the real value of their savings. Politicians want to signal their competence to the public by avoiding inflation in the months preceding an election, which leads to necessary devaluations being delayed until after elections.[4] Stein and Streb (2004) test for the existence of a political budget cycle in nominal exchange rates for a sample of 26 Latin American and Caribbean countries from 1960 to 1994. They find no evidence of pre-electoral appreciation, but do show that the average rate of nominal depreciation is two percentage points greater in the months following an election. Stein et al. (2005) show a similar political budget cycle in a sample of real, rather than nominal, exchange rates in 15 Latin American and Caribbean countries from 1960 to 1994. Bonomo and Terra (2005) emphasize the real exchange rate as well and demonstrate how a cycle of government spending on non-tradables could lead to real appreciation in the months before an election.[5]

2.2. Delayed depreciation approach

Other papers make an argument about political incentives for delayed devaluations, but do not do so specifically in the context of the political business cycle. For instance, Edwards (1994) studies 39 large devaluations (defined as a depreciation of at least 15%) in a sample of developing countries. He argues that democratic regimes are strongest when they are first elected and would naturally choose to devalue the exchange rate in this period. He shows that devaluations tend to take place right after an election in democratic regimes, but not so in dictatorial ones.

Gavin and Perotti (1997) find in a sample of 13 Latin American countries from 1968 to 1995 that the probability of a government abandoning an exchange rate peg (and moving to a flexible regime) is highest right after a national election. Klein and Marion (1997) test a similar hypothesis in a sample of 17 Latin American countries from 1956 to 1991. They argue that politicians weigh the economic costs of having a misaligned exchange rate with the political cost of devaluing right before an election. Like Gavin and Perotti (1997), they find that governments are most likely to abandon an exchange rate peg right after an election.

More recent evidence has bolstered these results. Frieden et al. (2001) study the pattern of real exchange rate movements in a sample of 26 Latin American and Caribbean countries from 1960 to 1994. They look at sample averages and find that the exchange rate is 3.7 times more likely to suffer a large depreciation (of at least 25%) in the post-election period as it is in the pre-election one.

2.3. Our approach

Our paper is novel in several ways. First, we test for post-electoral depreciations in the real exchange rate while controlling for other potentially relevant economic variables, conditional heteroskedasticity, and cross-sectional dependence. Much of the empirical literature on elections and exchange rates in Latin America has simply regressed the exchange rate on a set of electoral dummy variables. We control for other factors which might influence the real exchange rate, including the variables that Goldfajn and Valdés (1999) show are important determinants of the real exchange rate: the terms of trade, government expenditures, trade openness, and the international interest rate.[6]

Second, we test for electoral effects in the conditional variance. Much less work has been done on the question of whether elections generate increased exchange rate uncertainty, which is a potentially important phenomenon as it has been linked to reduced exports and even reduced economic growth.[7] There are several

[2] However, as we discuss further in Section 6, the interaction variables for central bank reform and financial liberalization are highly colinear, making our conclusions here only tentative. Our results generally support the idea advanced by Shi and Svensson (2006), that the existence of political cycles is conditional on the specific institutions in place within countries.

[3] For our purposes, we concentrate on papers that deal explicitly with Latin America.

[4] Note that while the authors do not explicitly recognize the link to the political budget cycle literature, the key element in their model, whereby debt today may lower inflation or devaluation today at a cost of more inflation tomorrow, is the same. Similar to this paper, Méon (2004) argues that voters can only indirectly determine the competency of the incumbent by his or her management of the exchange rate and for this reason, incumbents try to avoid a devaluation right before an election.

[5] Here the asymmetric information is not about how competent the incumbent government is but rather the uncertainty of whether the government will favor sectors lobbying for a depreciated real exchange rate, even when the population prefers a more appreciated rate. There are also several case studies of the real exchange rate and electoral cycles. Bonomo and Terra (1999) study the Brazilian case and find evidence of delayed devaluations (but not of a significantly appreciated exchange rate in the pre-election period). Pascó-Font and Ghezzi (2001) show pre-election appreciation and post-election depreciation of the Peruvian real exchange rate, as do Grier and Hernández-Trillo (2004) for the Mexican case.

[6] Bonomo and Terra (1999), however, do use Golfajn and Valdés equilibrium real exchange rate, which is determined by regressing the RER on the set of control variables that we discuss in the text.

[7] McKenzie (1999) surveys 32 empirical papers. For a recent contribution see Grier and Smallwood (2007) or Baum et al. (2004).

reasons to believe that elections may create exchange rate uncertainty. First, although elections resolve uncertainty in rational partisan models à la Alesina (1987), as policy is completely determined by the party that wins the election, the ex-ante forecast for the post-election period is a weighted average of the optimal policy of the two competing parties and the forecast error is thus quite likely to be larger right after the election than before (when the forecast is just based on the optimal policy of the party in power).

A second reason that uncertainty may be higher in the post-election period is that, at least in the case of exchange rate policy, knowing the party that won the election may not be sufficient to predict what policy they will undertake. Milesi-Ferretti (1995) and Méon (2001) both build theoretical models where governments can have incentives to act against their ideologies in the short run. Méon (2001) points out some real world anomalies, such as the fact that a Socialist government took France into the European Monetary System (EMS) while a Conservative government took England out. In addition, the statistical work of Eichengreen et al. (1995) shows no correlation between the ideology of the incumbent government and the occurrence of devaluations, currency flotations, or speculative attacks. Garfinkel et al. (1999) test in a sample of six G-7 countries whether surprise elections create uncertainty about the future exchange rate. Specifically, they calculate how surprising election outcomes were in each country and then measure the exchange rate forecast error (as implied by future contracts) before and after elections. They find that the elections that are characterized as the most surprising are associated with significantly larger forecast errors on average.

In the specific case of Latin America, candidates run on one platform only to deliver another with surprising frequency. Carlos Menem in Argentina, Alberto Fujimori in Peru, and Fernando Henrique Cardoso in Brazil are recent examples of this phenomenon. We thus posit that forecast errors for the real exchange rate will be greater immediately following elections both for Alesina-style reasons and Milesi-Ferretti and Méon-style reasons. We will incorporate a level shift in the conditional variance of the RER during post-election periods into our model to test for increased post-electoral uncertainty.

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

3. A statistical model of the real exchange rate

Our sample is a monthly panel of 9 Latin American countries (Brazil, Chile, Colombia, Costa Rica, Ecuador, Mexico, Peru, Uruguay, and Venezuela) from 1980–2000, encompassing a total of forty presidential elections. These particular countries were chosen because they have relatively complete data and their elections are at fixed dates and thus the timing is exogenous to economic variables. In this way we avoid issues of simultaneity between the economy and the election date.[8]

3.1. Variables

Our real exchange rate data comes both from J.P. Morgan and the IMF's International Financial Statistics CD-ROM.[9] 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.

We focus on the real exchange rate instead of the nominal one for four reasons. First, the real exchange rate is more directly linked to economic outcomes. There is not a one-to-one, nor even necessarily a monotonic, relationship between nominal and real exchange rate fluctuations and studying only the nominal exchange rate can give misleading inferences about whether politics affects the economically more important real exchange rate.[10] Second, while some of the countries in our sample have had periods of fixed nominal exchange rates, the real exchange rate in these countries could still fluctuate over the political cycle. Third, the hyperinflation that some of the countries experienced in the 1980s makes a time series analysis of the nominal exchange rate fairly problematic. Lastly, focusing on the real exchange rate conforms to the trend in the literature on electoral cycles and exchange rates, which has been moving away from studying nominal rates towards real rates.[11]

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, the International Foundation for Electoral Systems' Election Guide, and Bienen and van de Walle (1991). *Post* is equal to one for the month of the election (or in the case of a run-off election, the month of the run-off) and the subsequent five months. Appendix A 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 and 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

[8] Several 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 and 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.

[9] 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.

[10] 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.

[11] Bonomo and Terra (2005, p. 151) write that "Recent empirical studies on Latin American countries' exchange rate policy have identified a new type of electoral cycle: the real exchange rate (RER) tends to be more appreciated than average in the months preceding elections and more depreciated than average in the months following elections."

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 and 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.

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 and 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 and 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.

3.2. Statistical model

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. When we use tests that assume a common root, like the Levin et al. (2002) test, we generally fail to reject the null of a unit root. However, when we use tests that allow a separate unit root for each country, like the Im et al. (2003) test, we generally reject the null of a unit root.

We choose to use first differenced data in this paper for several reasons. First, most of the existing literature studies depreciation rates and not exchange rate levels. Second, first differencing will sweep out individual effects in the mean equation. Third, the statistical complications that arise from incorrectly differencing stationary data are less severe than the complications that arise from failing to difference non-stationary data (in the absence of co-integration). However, we have estimated our models in levels and the results are fully consistent with the results reported here using differences.[12]

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 (*totma*), openness (*openma*), and government spending (*govma*) are appropriate, where the moving average includes the second through fourth lags of the variables. The U.S. t-bill rate (*tbill*) is contemporaneously related to the RER in these countries and we interpret this correlation as one way causation from the US to Latin America. That is, we assume that the U.S. t-bill rate is uncorrelated with the error term in the exchange rate equation.

Our model, which is based on Cermeño and Grier (2006), allows for both conditional and unconditional heteroskedasticity and conditional cross-sectional correlation of the error terms.

The equations below present the specifics.

$$\Delta \ln(R_{it}) = \alpha_0 + \sum_{j=1}^{12}\left(\alpha_j \Delta \ln\left(R_{it-j}\right)\right) + \alpha_{13}\Delta \ln(totma_{it}) + \alpha_{14}\Delta \ln(openma_{it}) + \alpha_{15}\Delta \ln(govma_{it}) + \alpha_{16}\Delta \ln(tbill_t) + \mu Post_{it} + \varepsilon_{it} \tag{1}$$

$$h_{iit} = \Phi_1 \varepsilon^2_{iit-1} + \Phi_2 h_{iit-1} + \Phi_3 \mathrm{Br}_t + \Phi_4 \mathrm{Ch}_t + \Phi_5 \mathrm{Col}_t + \Phi_6 \mathrm{Cr}_t + \Phi_7 \mathrm{Ec}_t + \Phi_8 \mathrm{Mex}_t + \Phi_9 \mathrm{Per}_t + \Phi_{10} \mathrm{Ur}_t + \Phi_{11} \mathrm{Ve}_t + \theta Post_{it} \tag{2}$$

$$h_{ikt} = \rho_{ik} {}^* h_{iit} {}^* h_{kkt} \quad \text{For each country pair.} \tag{3 – 38}$$

Here i indexes countries and t indexes time. The error terms are assumed to be distributed multivariate normal with mean zero and time-varying variance–covariance matrix H_t, which is a 9×9 symmetric matrix that is positive definite for all periods t. The 9 diagonal elements of H_t (i.e. the h_{iit}) are given in Eq. (2) and the 36 unique off-diagonal elements (the h_{ikt}) are estimated in Eq. (3-38). The covariance specification follows Bollerslev (1990). Br, Ch, Col, Cr, Ec, Mex, Per, Ur, and Ve are country dummies representing Brazil, Chile, Colombia, Costa Rica, Ecuador, Mexico, Peru, Uruguay and Venezuela, respectively. In this way we allow for country specific unconditional error variances and time-varying conditional cross-sectional dependence.

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

4. Results

Table 1 presents some preliminary results of estimating Eq. (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 real exchange rate depreciation, 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 a more depreciated real exchange rate. 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.

While our paper is focused on the delayed depreciation/devaluation phenomenon, several papers in the literature present a full political business cycle model of the exchange rate that also allow for pre-election appreciation. We investigate the empirical relevance of

[12] These results are available from the authors upon request.

Table 1
An OLS model of the RER and elections.

$$\Delta\ln(R_{it}) = \underset{(0.9)}{1.35} + \underset{(2.8)}{.06}\,\Delta\ln(R_{it-1}) - \underset{(5.7)}{.12}\,\Delta\ln(R_{it-2}) - \underset{(1.5)}{.03}\,\Delta\ln(R_{it-3}) - \underset{(3.8)}{.08}\,\Delta\ln(R_{it-4}) + \underset{(0.6)}{.01}\,\Delta\ln(R_{it-5}) - \underset{(1.98)}{.04}\,\Delta\ln(R_{it-6}) - \underset{(2.4)}{.05}\,\Delta\ln(R_{it-7}) - \underset{(0.8)}{.02}\,\Delta\ln(R_{it-8}) + \underset{(0.5)}{.01}\,\Delta\ln(R_{it-9})$$
$$+ \underset{(1.0)}{.02}\,\Delta\ln(R_{it-10}) - \underset{(0.9)}{.02}\,\Delta\ln(R_{it-11}) - \underset{(0.2)}{.005}\,\Delta\ln(R_{it-12}) - \underset{(1.5)}{.15}\,\Delta\ln(tot_{ma}) - \underset{(2.1)}{.05}\,\Delta\ln(open_{ma}) + \underset{(.99)}{.01}\,\Delta\ln(gov_{ma}) + \underset{(3.4)}{0.07}\,\Delta\ln(tbill) - \underset{(3.4)}{14.7}\,Post$$

Numbers in parentheses are t-statistics.
Log-likelihood function = −11,925.
$N = 9$; $T = 239$.

Table 2
The RER and elections with a time-varying conditional covariance.

A. Mean equation
$$\Delta\ln(R_{it}) = \underset{(1.1)}{0.67} + \sum_{j=1}^{12} \underset{[.01]}{\alpha_j}\,\Delta\ln(R_{it-j}) + \underset{(1.9)}{.09}\,\Delta\ln(tot_{ma}) - \underset{(1.0)}{.009}\,\Delta\ln(open_{ma}) + \underset{(2.9)}{.011}\,\Delta\ln(gov_{ma}) + \underset{(6.9)}{.06}\,\Delta\ln(tbill) - \underset{(2.5)}{7.13}\,Post$$

B. Conditional variance equation
$$h_{iit} = \underset{(15.3)}{1.6}\,\varepsilon^2_{iit-1} + \underset{(5.2)}{.13}\,h_{iit-1} + \underset{(4.8)}{1177.5}\,Post + \text{country dummy variables}$$

C. Conditional correlations

	Estimated ρ_{ik}							
	Ch	Col	Cr	Ec	Mex	Per	Ur	Ven
Br	−.01	**.18**	.14	−.07	.06	.07	.12	.03
Ch		**.44**	**.33**	**.21**	**.33**	**.23**	**.26**	.05
Col			**.32**	.10	**.25**	**.21**	**.29**	.01
Cr				.05	**.14**	**.32**	**.20**	**.12**
Ec					.07	.06	**.13**	−.004
Mex						**.19**	**.22**	.09
Per							**.20**	.06
Ur								**.18**

Log-likelihood function = −10,909.
N=9; T=239.
Part A gives the panel equation for the conditional mean of RER Growth. 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).
Part B gives the conditional variance panel equation. Nine country dummies were estimated here but are not reported for reasons of space.
Part C presents the 36 estimated conditional error correlations for each country pair. Correlation coefficients in bold are significant at the .05 level or better.

this possibility in our sample by including a pre-election variable *Pre* that equals 1 for the 6 months before the election and zero otherwise. We re-estimate the model presented in Table 1 and find that the *Pre* variable is positive (a coefficient of 0.95) but completely insignificant (t-statistic of 0.22), while all the other coefficients are essentially unchanged from the values reported in Table 1.[13]

These initial results confirm and extend what others have argued about how elections can distort the evolution of exchange rates. Even when considering exogenous election dates and controlling for relevant economic variables, 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 follow the standard approach of testing for ARCH effects in the time series literature and adapt it to panel data. Specifically, 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 χ^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 χ^2 statistic is 16.26, which is significant at the 0.001 level. At five (ten) lags, the calculated χ^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 given above via direct numerical maximization of the log likelihood.[14] The model (and subsequent ones) consists of 38 equations, a panel equation for the conditional mean, a panel equation for the conditional variance which allows for country specific fixed effects, and 36 covariance equations (one for each unique country pair).

In Part A of the table, we report the panel equation for the conditional mean of RER growth. 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. The small sum of the lag coefficients indicates that RER growth is not very persistent in these data. The initial values tend to be negative and significant while the later values are generally positive and significant.

[13] These results are not reported directly but are available from the authors upon request. The lack of significant pre-electoral appreciation is consistent with the empirical findings of Frieden et al. (2001), Stein and Streb (2004), and Stein et al. (2005). We go on to test whether there is any significant relationship between the pre-election period and conditional variance. In a PBC model where all incumbents manipulate the money supply before an election (Lohmann, 1998), there would be no increased uncertainty associated with this period. In a model where there is a separating equilibrium, however, where politicians try to signal their competence with money supply increases, we would expect to see an increase in pre-electoral uncertainty. We test for this possibility by including *Pre* in both the conditional mean and conditional variance equations and find that the coefficients on *Pre* in the mean and conditional variance equations are never positive and significant.

[14] 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.

The terms of trade variable is now positive and significant at the .05 level, indicating that improvements in the terms of trade are associated with real exchange rate appreciation. The openness variable is negative, but is no longer significantly related to the real exchange rate. Government expenditures and changes in the t-bill rate are positively and significantly correlated with real exchange rate appreciations at the .01 level. The coefficient on *Post* is still negative and significant at the .01 level, but the size of the electoral effect drops by around 50% from Table 1.

In Part B we report the panel equation for the conditional variance. The coefficients on the nine country dummy variables are individually significant and significantly different from one another but are not reported to save space. The ARCH and GARCH coefficients are positive and significant at the .01 level. We 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 1,100, 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.[15]

Part C of the table contains the estimated cross-country conditional correlations which come directly from the 36 estimated conditional covariance equations. Twenty 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 Uruguay and Chile, while the countries with the fewest significant links are Venezuela and Ecuador.

As a 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 χ^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.

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 central bank reform in the region has had any effect on this electoral exchange rate cycle.

5. 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 (1) changed central bank charters so that the sole (or at least primary) objective would be price stability, (2) reduced the dependence of central banks on the executive branch of government, and (3) increased the economic autonomy of central banks. For example, 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. Appendix A provides the dates of the reforms in each country.

In theory, these reforms work to insulate central banks from political pressure and as such, if they are effective, should reduce the effect of elections on the path of the real exchange rate. 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. Note that 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. In addition, we interact this variable with *Post* to investigate whether 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.

The effect of CBR on RER uncertainty is less clear cut as it revolves around the relative predictability of governments versus 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.

Controlling for central bank reform changes Eqs. (1) and (2) to the following:

$$\Delta \ln(R_{it}) = \alpha_0 + \sum_{j=1}^{12}\Big(\alpha_j \Delta \ln\Big(R_{it-j}\Big)\Big) + \alpha_{13}\Delta \ln(totma_{it}) + \alpha_{14}\Delta \ln(openma_{it}) + \alpha_{15}\Delta \ln(govma_{it}) + \alpha_{16}\Delta \ln(tbill_t) + \mu Post_{it} + \beta Reform_{it} + \Lambda Post_{it}{*}Reform_{it} + \varepsilon_{it} \qquad (1')$$

$$h_{iit} = \Phi_1 \varepsilon^2_{iit-1} + \Phi_2 h_{iit-1} + \Phi_3 \mathrm{Br}_t + \Phi_4 \mathrm{Ch}_t + \Phi_5 \mathrm{Col}_t + \Phi_6 \mathrm{Cr}_t + \Phi_7 \mathrm{Ec}_t + \Phi_8 \mathrm{Mex}_t + \Phi_9 \mathrm{Per}_t + \Phi_{10} \mathrm{Ur}_t + \Phi_{11} \mathrm{Ven}_t + \theta Post_{it} + \chi Reform_{it} + \Psi Post_{it}{*}Reform_{it} \quad \text{For all } I. \qquad (2')$$

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. Swinburne and 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.

Table 3 reports estimates of Eqs. (1′), (2′) and (3-38). To easily compare the results of Tables 2 and 3, we provide an additional

[15] Note that the coefficient on government expenditures is now positive and significant. Ames (1987) shows evidence of fiscal cycles in Latin America by finding that government expenditures rise before elections and fall afterwards. The fact that *Post* is positive and significant while controlling for government spending means that the post-electoral effect on exchange rate depreciation goes beyond any effect that may be due to the fiscal cycle.

Table 3
The effect of central bank reform on the election-RER relationship.

A. Mean equation

$$\Delta\ln(R_{it}) = \underset{(0.8)}{-.59} + \underset{[.01]}{\sum_{j=1}^{12}} \alpha_j\,\Delta\ln(R_{it-j}) + \underset{(2.3)}{.11}\,\Delta\ln(tot_{ma}) - \underset{(1.7)}{.017}\,\Delta\ln(open_{ma}) + \underset{(2.6)}{.01}\,\Delta\ln(gov_{ma}) + \underset{(6.9)}{.06}\,\Delta\ln(tbill) - \underset{(2.9)}{9.68}\,Post + \underset{(2.2)}{13.6}\,Post{*}Reform + \underset{(3.9)}{4.19}\,Reform$$

B. Conditional variance equation

$$h_{iit} = \underset{(15.3)}{1.6}\,\varepsilon^2_{iit-1} + \underset{(5.3)}{.13}\,h_{iit-1} + \underset{(.97)}{464.8}\,Post + \underset{(1.26)}{730.9}\,Post{*}Reform - \underset{(3.0)}{88.7}\,Reform + country\ dummy\ variables$$

C. Conditional correlations

	Estimated ρ_{ik}							
	Ch	Col	Cr	Ec	Mex	Per	Ur	Ven
Br	.004	**.18**	**.15**	−.07	.07	.07	**.12**	.04
Ch		**.43**	**.34**	**.21**	**.34**	**.24**	**.25**	.05
Col			**.30**	.09	**.25**	**.20**	**.28**	.01
Cr				.06	.13	**.32**	**.20**	.11
Ec					.07	.07	**.13**	−.008
Mex						**.20**	**.22**	.08
Per							**.20**	.06
Ur								**.17**

Log-likelihood function = −10,909.
N=9; T=239.
Part A gives the panel equation for the conditional mean of RER Growth. 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).
Part B gives the conditional variance panel equation. Nine country dummies were estimated here but are not reported for reasons of space.
Part C presents the estimated conditional error correlations for each country pair. Correlation coefficients in bold are significant at the .05 level or better.

table (Table 5), which compares the political and reform coefficients from each of the different models. In Part A of the table, we again 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 Goldfajn–Valdés control variables are of the same sign and significance levels as those in Table 2, with the only exception that openness is now significant at the .05 level. *Post* is negative and significant at the .01 level, while the two new variables that did not appear in Table 2 (*Reform* and *Post*Reform*) are each positive and significant at the .01 level.

These results indicate that CBR is an important factor for understanding the evolution of the real exchange rate process in these countries. The reform variable is positive and significant in the conditional mean equation, meaning that there is less average depreciation after central bank reform. In addition, 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 10% per month in the post-election period, holding other factors constant, while post-reform there is, on net, a small, but not statistically significant from zero, post-election appreciation effect.

Panel B presents the panel equation for the conditional variance with the inclusion of the two new dummy variables. Like the previous table, we find that the country dummy variables are individually significant and also significantly different from one another. The ARCH and GARCH coefficients remain positive and significant. The coefficient on *Post*, however, is now insignificantly different from zero, while the two new variables, *Reform* and *Post* Reform*, are jointly significant at the .01 level. The effect of CBR on post-electoral uncertainty is not clear cut. In Table 2, *Post* had a large positive coefficient in the conditional variance equation. Here in Table 3, *Post* and (*Post*Reform*) are each insignificant in the conditional variance equation. However, the magnitude of the effect of elections on uncertainty after the enactment of CBR is quite large (it is given by the sum of the two coefficients and is thus equal to almost 1200). While CBR has eliminated electoral effects in the conditional mean, it may have made the RER less predictable right after elections as well, indicating that perhaps the reforms were not fully credible.

In sum, our results clearly show that CBR significantly changed 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.

6. Financial liberalization and the RER

While the above results are striking, one can never be completely sure that CBR is the only relevant political reform that could explain the demise of the RER election cycle in the region. In this section, we consider one plausible additional explanation, namely that financial liberalization has helped to end the RER election cycle.

Financial liberalization makes any political manipulation of the exchange rate more costly because it, at heart, removes controls on the flow of financial capital into and out of a country. When a country opens up to, and receives, foreign capital flows, it generally wants to keep those flows. Electorally inspired exchange rate movements could cause damaging capital outflows. In sum, financial liberalization increases the cost of politically-based exchange rate manipulation and thus might well mitigate it.

To explore this possibility, 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 Harvey et al. (2002, 2005) and summarized in Harvey and Bekaert (2004). Appendix A provides the dates of liberalization for each country.

Table 4 reports the results of re-estimating our model with the inclusion of these new variables. Again, for ease of comparison, Table 5 presents all of the political and reform coefficients from the different models. In Panel A, we present the panel equation for the conditional mean of the RER. 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 economic control variables are virtually unchanged from Table 3. The two new variables, *Liberalize* and *Post*Liberalize*, are individually and jointly insignificant. The coefficients on *Post* and *Post*Reform* continue to show significant post-election depreciation before the enactment of CBR and none afterward.

Table 4
The effect of central bank reform and financial liberalization on the election-RER relationship.

A. Mean equation

$$\Delta \ln(R_{it}) = -\underset{(0.7)}{.58} + \sum_{j=1}^{12} \underset{[.01]}{\alpha_j} \Delta \ln(R_{it-j}) + \underset{(2.3)}{.11} \Delta \ln(tot_{ma}) - \underset{(1.8)}{.018} \Delta \ln(open_{ma}) + \underset{(2.7)}{.011} \Delta \ln(gov_{ma}) + \underset{(7.2)}{.07} \Delta \ln(tbill) - \underset{(2.5)}{9.08}\, Post + \underset{(2.0)}{8.32}\, Post{*}Reform + \underset{(3.1)}{3.75}\, Reform + \underset{(.82)}{.96}\, Liberalize + \underset{(.44)}{3.76}\, Post{*}Liberalize$$

B. Conditional variance equation

$$h_{iit} = \underset{(15.2)}{1.6}\, \varepsilon^2_{iit-1} + \underset{(4.6)}{.12}\, h_{iit-1} + \underset{(0.9)}{474.9}\, Post + \underset{(1.0)}{518.9}\, Post{*}Reform - \underset{(2.1)}{63.7}\, Reform - \underset{(1.9)}{147.1}\, Liberalize + \underset{(3.95)}{1687.4}\, Post{*}Liberalize + country\ dummy\ variables$$

C. Conditional correlations

	Estimated ρ_{ik}							
	Ch	Col	Cr	Ec	Mex	Per	Ur	Ven
Br	.0	**.18**	**.15**	−.06	.07	.08	**.12**	.03
Ch		**.43**	**.33**	**.22**	**.37**	**.24**	**.24**	.05
Col			**.31**	.09	**.27**	**.20**	**.26**	−.001
Cr				.06	.12	**.30**	**.18**	.11
Ec					.08	.07	**.13**	−.01
Mex						**.21**	**.22**	.07
Per							**.20**	.05
Ur								**.16**

Log-likelihood function = −10,909.
N=9; T=239.
Part A gives the panel equation for the conditional mean of RER Growth. 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).
Part B gives the conditional variance panel equation. Nine country dummies were estimated here but are not reported for reasons of space.
Part C presents the estimated conditional error correlations for each country pair. Correlation coefficients in bold are significant at the .05 level or better.

Table 5
Comparison of political and reform coefficients across models.

	Table 2	Table 3	Table 4
Mean equation			
Post	−7.13 (2.5)	−9.68 (2.9)	−9.08 (2.5)
*Post*Reform*	–	13.6 (2.2)	8.32 (2.0)
Reform	–	4.19 (3.9)	3.75 (3.1)
Liberalize	–	–	.96 (.82)
*Post*Liberalize*	–	–	3.76 (.44)
Conditional variance equation			
Post	1177.5 (4.8)	464.8 (0.97)	474.9 (0.9)
*Post*Reform*	–	730.9 (1.26)	−518.5 (1.0)
Reform	–	−88.7 (3.0)	−63.7 (2.1)
Liberalize	–	–	−147.1 (1.9)
*Post*Liberalize*	–	–	1687.4 (3.95)

Numbers in parentheses are asymptotic t-statistics.

From the conditional variance equation reported in Panel B, we find that both financial liberalization and CBR significantly reduce uncertainty, but we now see a significant increase in post-election uncertainty in countries that have opened their financial markets.[16] These findings seem to lean towards CBR as the key to stopping the electoral cycle. However, we should point out again that while the dates of CBR and financial liberalization vary, their interactions with elections are collinear, so that these findings about the relative strengths of these two factors should be regarded as tentative.[17]

7. 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 statistical model which controls for potentially important economic variables and 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 should accompany democratic reforms in order to mitigate any tendency for elections to create exchange rate instability. It also suggests that political reversals of CBR, as recently demonstrated in Venezuela, may be cause for concern.

[16] The coefficient seems extremely large, so it is helpful to put it into context. We also estimate but do not report country fixed effects in the conditional variance equation. These range from around 108 up to 2900. So the coefficient of 1687 is within that range and not a massive outlier. It was suggested by a referee that the 1994 Mexican Crisis might be causing this result. We re-estimated the model dropping Mexico from the sample and the coefficient in question falls to 980, but remains positive and significant. We leave further analysis of this anomalous result for future research.

[17] There are only 5 elections where the variables *Post*Reform* and *Post*Liberalize* differ. When we drop CBR and only use *Liberalize* in the model, we find that *Post*Liberalize* is not significant in the conditional mean equation, which provides some further support to the findings reported in Table 4 that the disappearance of the RER election cycle is more closely correlated with CBR than with financial liberalization.

Appendix A. Presidential elections, central bank reform and financial liberalization

Country	Elections (runoff)	CB reform	Liberalization
Brazil	1/1985 11/1989 (12/89) 10/1994 10/1998	10/1988	5/1991
Chile	12/1989 12/1993 12/1999 (1/00)	10/1989	1/1990
Colombia	5/1982 5/1986 5/1990 5/1994 5/1998 (6/98)	8/1991	2/1991
Costa Rica	2/1982 2/1986 2/1990 2/1994 2/1998	11/1995	n.a.
Ecuador	1/1984 (5/84) 1/1988 (5/88) 4/1992 (7/92) 5/1996 (7/96) 5/1998 (7/98)	5/1992	9/1994
Mexico	7/1982 7/1988 8/1994 8/2000	12/1993	5/1989
Peru	5/1980 4/1985 4/1990 (6/90) 4/1995 4/2000 (5/00)	1/1993	1/1991
Uruguay	11/1984 11/1989 11/1994 10/1999	3/1995	n.a.
Venezuela	12/1983 12/1988 12/1993 12/1998 7/2000	12/1992	1/1990

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