PABLO GAROFALO
New Jersey City University, United States

JORGE M. STREB
Universidad del CEMA (UCEMA), Argentina

Broken Promises: Regime Announcements and Exchange Rates around Elections

ABSTRACT We study exchange rate dynamics around government changes conditional on the exchange rate regime, which we identify by combining the IMF de jure and the Reinhart and Rogoff de facto exchange rate regime classifications. This allows distinguishing whether the official exchange rate regime announcements match actual policy or are inconsistent with it. Using monthly data from Latin American democracies, we do not find significant exchange rate depreciations before the change of government in any of the regimes we identify. However, we do detect a gradual real exchange rate overvaluation when the de jure regime is fixed but the de facto policy is flexible, which is abruptly corrected after the change of government; this pattern of real exchange rate misalignments when the announcement does not match actual behavior is linked to incumbents that postpone devaluations until the successor steps in. This pattern of broken promises is typical until 1999, but it becomes exceptional thereafter.

JEL Codes: D72, D78, E00

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Reneging on exchange rate regime announcements occurs quite often. We track exchange rate regime announcements with the de jure exchange rate classification maintained by the International Monetary Fund (IMF), which reports what countries claim to be doing.[1] The IMF de jure classification has been criticized for representing words, not deeds (Levy-Yeyati and Sturzenegger, 2005; Reinhart and Rogoff, 2004). Among the de facto classifications proposed, Reinhart and Rogoff (2004) reclassify exchange rate arrangements by developing an algorithm based on the observed behavior of

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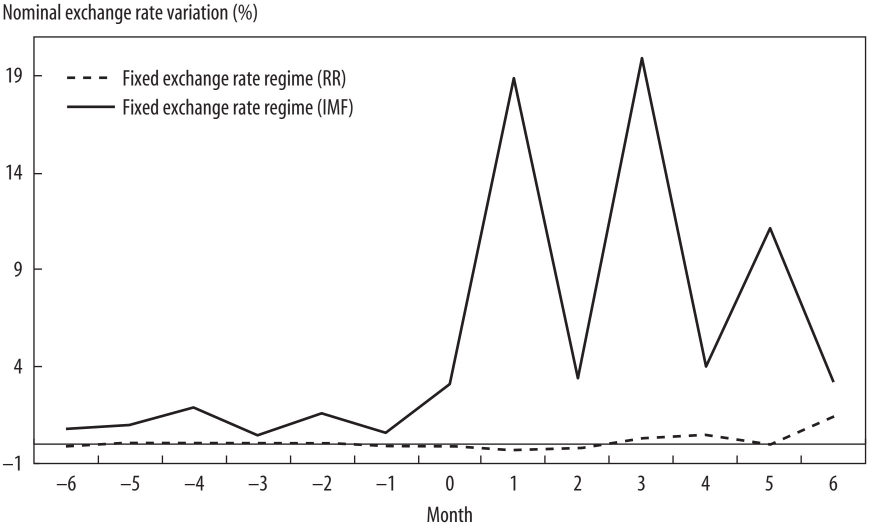
1. Exchange rate regime announcements are here distinguished from firmer monetary commitments like dollarization, in which the country relinquishes an independent currency, for example, Panama since 1904, Ecuador since 2000, and El Salvador since 2001. A regime announcement continues to hold until further notice.

FIGURE 1. Exchange Rate Devaluations around Government Changes



Note: Average exchange rate variation during nineteen and twenty-one complete episodes computed with the RR and IMF fixed exchange rate classifications, respectively, in twenty-one Latin American countries (Argentina, Barbados, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Guyana, Honduras, Jamaica, Mexico, Nicaragua, Paraguay, Peru, Trinidad and Tobago, Uruguay, and Venezuela) in the 1980–99 period. Both classifications are invariant throughout the twelve-month window. Month 0 is the government change month. Dollarization episodes are excluded.

exchange rates, while parallel exchange rates are used if multiple markets are present. While Reinhart and Rogoff (2004, p. 1) claim that the IMF exchange rate classification is "a little better than random," we have reasons to suspect otherwise.

Using the IMF and Reinhart and Rogoff (RR) classifications, figure 1 shows nominal exchange rate variations around government changes (when an incumbent's term ends and a new administration is inaugurated) in Latin American countries, conditional on a fixed exchange rate regime within the whole window. Devaluations are similar under both classifications up to the month of government change, but they increase considerably thereafter under the IMF de jure classification. This suggests that some exchange rate pegs are sustained in the prelude to elections and government changes, but not afterward.

Calvo and Reinhart (2002) show that many Latin American countries that claim to be floating are not doing so, a phenomenon known as fear of floating. This occurs, for instance, when a country classified as floating is in reality pegging its exchange rate to, say, the U.S. dollar. Conversely, Alesina and Wagner (2006) show that some countries often break commitments to pegging and end up floating more than they announce, a phenomenon they call fear of pegging. We analyze these mismatches around elections. Inspired by Alesina and Wagner (2006), we combine the de jure and de facto classifications to distinguish between "keeping" and "breaking" promises, that is, between regime announcements that are consistent with observed market-based exchange rate behavior and those that are not. As Genberg and Swoboda (2005) note, both announcements and actions may provide useful information about exchange rate policy. We thus explore the behavior of exchange rates conditional on the regimes that we identify based on the consistency of the de jure and de facto classifications. To the best of our knowledge, nobody has analyzed this issue before. While there is ample evidence on the delay of exchange rate adjustments when elections are coming up (for example, Cermeño, Grier, and Grier, 2010; Edwards, 1994; Stein and Streb, 2004; Stein, Streb, and Ghezzi, 2005), these earlier studies may suffer from downward bias because they do not control for either exchange rate regimes or the consistency between the de jure and de facto classifications: their results are a weighted average of devaluations in inconsistent de jure fixed exchange rate regimes, where we find that all the variability is concentrated, and all the other regimes, where no pattern is found. We henceforth focus on these inconsistent de jure fixed regimes, those exhibiting the so-called fear of pegging.[2] Observationally, de jure fixed regimes that are inconsistent with the de facto flexible behavior share identifiable underlying characteristics, namely, dual or multiple markets and high inflation before elections. These inconsistent fixed regimes, for short, always involve broken promises. In contrast, fear of floating need not imply broken promises: Genberg and Swoboda (2005) point out that a country that may seem to be pegging its currency to another country's might simply be following a similar monetary policy, so it is not breaking any commitment.

We first study the determinants of exchange rate regimes around elections using ordered logit models for both the IMF de jure and RR de facto regime classifications. As found by Klein and Marion (1997) and Gavin and Perotti (1997), there is no evidence that de jure regimes change before the government turnover, but the probability of abandoning a fixed exchange rate regime increases after the new administration is inaugurated. Additionally, we find that the probability of a de facto flexible regime increases before government

2. Inconsistent de jure fixed regimes are a slight modification of what Alesina and Wagner call fear of pegging. We develop the rationale for this modification later in the paper.

changes. We then rely on a multinomial logit model, which is widely used for unordered categories, to study the consistency between de jure and de facto regimes. This is something novel in the literature on exchange rate regimes. After government changes, we detect that the probability of inconsistent fixed regimes, which are de facto flexible, decreases in relation to the probability of consistent flexible regimes, with are both de jure and de facto flexible. In other words, though the market regime behavior already involves some degree of float, the authorities announce it after the inauguration of a new government, not before.

Second, we study the dynamics of the real exchange rate around the government change month, conditional on whether the de jure regime matches the de facto regime before the month of the election, by using a dynamic distributed lag model and a difference-in-differences strategy. We find that exchange rate behavior during consistent and inconsistent de jure fixed regimes is not statistically different until the month of the government turnover, but it differs significantly in the first quarter after that. Hence, although inconsistent fixed regimes might tend to be episodes of "poor macroeconomic performance and inability to maintain monetary and fiscal stability" (Alesina and Wagner, 2006, p. 774), official exchange rates are sustained until the government change date.

Third, the paper contributes to the literature on real exchange rate appreciations and their reversions. Goldfajn and Valdés (1999) show that real exchange rate appreciations are usually reverted by nominal exchange rate devaluations rather than by smooth inflation differentials. This nominal adjustment through sharp exchange rate devaluations causes overvaluation to last longer during the buildup stage than during the reversion stage. In our sample of Latin American countries, the overvaluation of the real exchange rate occurs only for inconsistent fixed regimes. Such overvaluation begins ten months before the government change date and lasts until two months after the government turnover (about one year of overvaluation), with a peak of 37 percent in the government change month. Reversion starts abruptly the next month and is completed in three months. This corroborates the findings by Goldfajn and Valdés (1999) on the asymmetry between the buildup and reversion stages due to sudden nominal exchange rate adjustments. While they do not characterize and describe the context in which these appreciation episodes take place, we identify one particular context where they occur: inconsistent fixed regimes before government changes.

Finally, we compare the 1980–99 period, to which the findings described above apply, with the 2000–16 period. Although the IMF changed to a de facto classification after 1999, we can use the fact that dual/multiple regime practices are an underlying characteristic of inconsistent de jure fixed regimes to study the recent period. We find that dual/multiple regimes are almost nonexistent in the 2000–16 period, but we provide a case study of the only election where the regime is likely to be classified as inconsistent fixed, namely, the Argentine general election of 2015.

The rest of the paper is organized as follows. The next section reviews the exchange rate classification literature. We then explain our methodology for identifying consistent and inconsistent exchange rate regime announcements. Subsequently we present the econometric models and results, analyze the appreciation of the real exchange rate and its reversion, and compare the 1980–99 period with the most recent period by relying on the underlying characteristics of inconsistent fixed regimes. The final section concludes.

Exchange Rate Regime Classifications

The IMF developed a traditional exchange rate regime classification, which it has published since 1950.[3] Until 1999, it asked "country members to self-declare their arrangement as belonging to one of four categories" (Alesina and Wagner, 2006, p. 775): float, managed, crawl, or fixed. If a country announced the adoption of a floating regime in a specific year, "the IMF classified this country-year as floating even if in practice this country pegged its currency to, say, the U.S. dollar" (Alesina and Wagner, 2006, p. 775). There are many reasons to seek other approaches to classifying exchange rate regimes. For instance, empirical work on the costs and benefits of alternative exchange rate arrangements can be misleading when actual behavior deviates significantly from the announced behavior; as Reinhart and Rogoff (2004) point out, Baxter and Stockman (1989) find that there are no significant differences in business cycles across exchange arrangements.

Reinhart and Rogoff (2004) provide a "natural classification" of exchange rate regimes that relies on a broad variety of descriptive statistics to group episodes into a grid of regimes based on market-determined exchange rate behavior. They provide detailed analyses to posit the importance of market-determined exchange rates as the best indicator of the underlying monetary policy. They first do so by showing that the market exchange rate consistently

3. *Annual Report on Exchange Arrangements and Exchange Restrictions*.

anticipates devaluation of the official rate, and not vice versa. Second, they find that the market-determined exchange rate keeps up with inflation, while the official rate sometimes does not. Additionally, they remark that "it is not unusual for dual or parallel markets (legal or otherwise) to account for the lion's share of transactions with the official rate being little more than symbolic" (Reinhart and Rogoff, 2004, p. 10).

To create the natural classification, they first check whether there is a unified rate instead of dual or parallel (black) markets. If there is a dual or parallel market, given the relevance of the market-determined rate explained above, they classify the regime as de facto using the market-determined exchange rate. If there is no parallel market, they examine summary statistics to verify the official de jure arrangement, if any, going forward from the date of the announcement. If the regime is verified, it is then classified as de jure. If the de jure regime fails verification, they seek a de facto statistical classification based on the behavior of the exchange rate if inflation is below 40 percent. When annual inflation is above 40 percent, the exchange rate is classified as "freely falling." A similar statistical classification is conducted when there is no preannounced path for the exchange rate. In all, they establish fourteen categories in what they call their fine grid, which they collapse into five categories for their coarse grid. We use the latter in our analysis.

Levy-Yeyati and Sturzenegger (2005) also provide a de facto classification of exchange rate regimes. Besides exchange rates, their algorithm uses base money and international reserves. While both classifications have their merits, the RR classification suits our analysis better because it provides a monthly classification that allows us to observe switching regimes, if any, around elections and government change dates, which is important to determine the endogeneity of the regime. Moreover, Levy-Yeyati and Sturzenegger (2005) use the official exchange rate in their de facto algorithm, rather than market rates. Like Alesina and Wagner (2006, p. 797), we are interested in how de facto behavior deviates from announced official policies, so this also points to the RR classification.

Consistency of De Jure and De Facto Exchange Rate Regimes

To identify consistent and inconsistent de jure regimes (that is, whether the official announced regime matches the actual policy), we follow an approach similar to Alesina and Wagner (2006), who quantify broken promises, which we call inconsistencies, as the difference between the coarse RR and IMF

FIGURE 2. Classification of De Jure Exchange Rate Regimes by Alesina and Wagner (2006)

RR de facto classification (actual policy)

	Fixed	Crawl	Managed	Float
Float	4, 1 (+3)	4, 2 (+2)	4, 3 (+1)	4, 4 (0)
Managed	3, 1 (+2)	3, 2 (+1)	3, 3 (0)	3, 4 (–1)
Crawl	2, 1 (+1)	2, 2 (0)	2, 3 (–1)	2, 4 (–2)
Fixed	1, 1 (0)	1, 2 (–1)	1, 3 (–2)	1, 4 (–3)

IMF de jure classification (announcement)

Source: Alesina and Wagner (2006).

Note: Each cell contains three numbers: X, Y (Z). X represents the RR classification and Y the IMF classification (where 4 = float, 3 = managed, 2 = crawl, and 1 = fixed), and $Z = X - Y$. Dark gray: Fear of floating, with more management than announced ($Z < 0$). Light gray: Fear of pegging, with more floating than announced ($Z > 0$).

classifications. They assign a value from one to four to identify the regime (fixed: 1; crawl: 2; managed: 3; float: 4) and then subtract the de jure value from the de facto value. For example, if the RR natural classification of the regime is a float (with a value of 4) while the IMF de jure classification is managed (value of 3), then the difference (denoted Z) is positive and called fear of pegging. A negative value for Z, in turn, indicates fear of floating, while a value of zero represents a consistent regime. Figure 2 shows all the possible combinations.

This classification does not control for the intensity of the differences between the RR and IMF classifications. It applies equally to $Z = –3$ and $Z = –1$, without distinguishing between strong and weak fear of floating (an analogous observation holds for $Z > 0$ regarding the different intensities of fear of pegging). This issue is the starting point for our regime classification below. Our main innovation lies in dividing consistent de jure regimes into fixed (fixed or crawl) and flexible (managed or float). We create the categories using a two-dimensional classification system: fixed versus flexible and consistent versus inconsistent. Our approach is depicted in figure 3, which presents four categories of de jure regimes: (1) consistent fixed (intermediate gray), (2) consistent flexible (unshaded), (3) inconsistent fixed (strong fear

FIGURE 3. Alternative Classification of De Jure Exchange Rate Regimes

RR de facto classification (actual policy)

Float	4, 1 (+3)	4, 2 (+2)	4, 3 (+1)	4, 4 (0)
Managed	3, 1 (+2)	3, 2 (+1)	3, 3 (0)	3, 4 (–1)
Crawl	2, 1 (+1)	2, 2 (0)	2, 3 (–1)	2, 4 (–2)
Fixed	1, 1 (0)	1, 2 (–1)	1, 3 (–2)	1, 4 (–3)
	Fixed	Crawl	Managed	Float

IMF de jure classification (announcement)

Note: Each cell contains three numbers: *X*, *Y* (*Z*). *X* represents the RR classification and *Y* the IMF classification (where 4 = float, 3 = managed, 2 = crawl, and 1 = fixed), and *Z* = *X* – *Y*. Dark gray: inconsistent flexible (*Z* < –1). Light gray: inconsistent fixed (*Z* > 1). Intermediate gray: consistent fixed. Unshaded: consistent flexible.

of pegging; light gray), and (4) inconsistent flexible (strong fear of floating; dark gray). The consistent de jure regime categories correspond to $Z = 0$ and $Z = \pm 1$ when there is either a match between the actual policy and the de jure regime ($Z = 0$) or a weak departure ($Z = \pm 1$). This is how we differentiate the intensity of the episodes in our analysis; that is, $Z \leq \text{abs}(1)$ belongs to consistent de jure regimes, while $Z > \text{abs}(1)$ belongs to inconsistent ones.

Data, Econometric Specifications, and Results

Our main focus is real exchange rate dynamics around government change dates conditional on the consistency of the de jure exchange rate regimes. We first study the determinants of the exchange rate regime policies and the extent to which they are sensitive to the electoral window. This is an important issue to address since regime types are used as controls in the study of exchange rate dynamics. Therefore, netting out covariates, we would like to see how endogenous regimes are around government changes, if at all. We can only carry out these econometric analyses for the 1980–99 period because the IMF abandoned its de jure classification after that.

We collected monthly data on exchange rates and inflation from twenty-one Latin American countries from the IMF International Financial Statistics (IFS) database over the 1980–99 period. The countries are Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Guyana, Honduras, Jamaica, Mexico, Nicaragua, Panama, Paraguay, Peru, Trinidad and Tobago, Uruguay, and Venezuela.[4] We constructed the series of the multilateral real exchange rate, which is a trade-weighted average of bilateral real exchange rates. We follow Goldfajn and Valdés (1999) in using only trading partners above 4 percent of overall trade. Also as in Goldfajn and Valdés (1999), we fixed the trade weights using trade flows of an intermediate year (1995 in our case) from the United Nations *International Trade Statistics Yearbook*.[5] Monthly observations of the RR natural exchange rate regime classification are from Ethan Ilzetzki's website, an updated version of the original data from Carmen M. Reinhart's website.[6] The traditional IMF annual exchange rate regime classification comes from the IMF *Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER)*.[7] We conducted a country-by-country study to transform the IMF annual classification into monthly series by reviewing all AREAER manuals from 1980 to 1999 (see details, methodology, and sources in online appendix A).[8]

Duration of Exchange Rate Regimes

We compare regime duration inside and outside the electoral window, because exchange rate estimations controlling for regime at election time may be biased

4. Chile, El Salvador, Guyana, Jamaica, Paraguay, and Trinidad and Tobago are dropped from the sample when the full set of covariates is used owing to missing observations in control variables for these countries. We thus work with two samples: a reduced sample that excludes these countries and an extended sample that includes them. Results for the reduced sample are very similar with and without covariates. Results for the extended sample are only available without covariates. While results without controls are somewhat smaller in magnitude for the extended sample, they are significant in both cases.

5. Identical qualitative results were found using only the bilateral real exchange rate with the United States. This may be because the United States is the main trading partner for almost all Latin American countries. We therefore conclude that our results should not be sensitive to the year of weights used. These alternative results are available on request.

6. Ethan Ilzetzki: www.ilzetzki.com/irr-data; Carmen M. Reinhart: www.carmenreinhart.com. See Ilzetzki, Reinhart, and Rogoff (2019).

7. Available online at 0-www-elibrary-imf-org.library.svsu.edu/subject/012?t=F&type_0=book&type_1=journalissue.

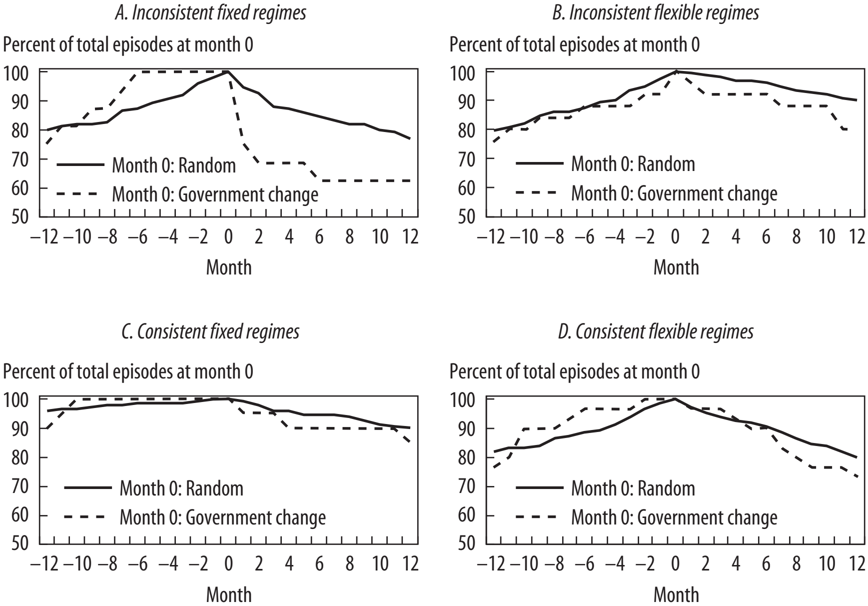
8. Supplementary material for this paper is available online at http://economia.lacea.org/contents.htm.

if regime duration is sensitive to the electoral window. We proceed to report summary statistics for our four-category regime classification: (1) inconsistent fixed (that is, de jure fixed in the IMF classification and de facto flexible in the RR classification); (2) consistent fixed (both de jure and de facto fixed); (3) inconsistent flexible (de jure flexible and de facto fixed); and (4) consistent flexible (both de jure and de facto flexible).

To compare regime duration inside and outside the electoral window, we generate, with a uniform distribution, a random number between one and 240 in order to select a month in the 1980–99 period for a given country. We then observe the regime classification for that month and construct a twenty-four-month window around the observation (twelve months on either side) to identify the duration of that regime. For example, an episode might change at month –5 (that is, the classification at month 0 started five months earlier) or at month +5 (that is, the classification at month 0 ended five months later). We repeat this randomization fifty times for each country and then, for each month, calculate the percentage of episodes in which the regime continues to equal that of month 0. This randomization process allows us to have an idea of the duration of regimes, independently of the covariates. We conduct exactly the same exercise around government change months. The percentage of episodes in each category is very similar for the two windows: 15.4 percent (17.6 percent) are inconsistent fixed regimes for the random (government change) month 0; 28.3 percent (27.5 percent) are inconsistent flexible regimes; 22.9 percent (22.0 percent) are consistent fixed regimes; and 33.5 percent (33.0 percent) are consistent flexible regimes. This similarity suggests that the distribution of regimes is not sensitive to the electoral window. Results are displayed in figure 4.

When looking at the duration of the regimes in figure 4, about 95 percent of the episodes of consistent fixed regimes (panel C) were already in that category eleven months before month 0 for both windows, and about 90 percent of the episodes continue to belong to this category eleven months later. Distributions are also similar for inconsistent flexible (panel B) and consistent flexible regimes (panel D). The exception is inconsistent fixed regimes (panel A). Cases in this category appear to have a longer duration before month 0 for the government change window: for instance, 100 percent of the government change episodes in this category had already started by month –6, versus only 87 percent for the random data window. Although this sounds problematic, in the next subsection we show that, after netting out covariates, the probability of inconsistent fixed regimes does not increase when government change approaches. In contrast, the abrupt decrease after government changes

FIGURE 4. Duration of Regime Classifications around Random and Government Change Months



Notes: Number of regime episodes at government change month 0: inconsistent fixed, 16; inconsistent flexible, 25; consistent flexible, 30; consistent fixed, 20. Number of regime episodes at random month 0: inconsistent fixed, 169; inconsistent flexible, 311; consistent flexible, 368; consistent fixed, 252.

anticipates a finding below: some incumbents switch categories from inconsistent fixed to consistent flexible.

Determinants of Exchange Rate Regimes

We use two estimation methods in this subsection. We first study the IMF de jure and the RR de facto regime classification separately for the period 1980–99 to compare with the findings of previous studies. Since both classifications represent clear ranks with a meaningful order (that is, fixed, crawl, managed, and float), ordered logit models are the appropriate option. We then study the determinants of our novel regime classification, that is, de jure fixed and flexible regimes that are either consistent or inconsistent, as shown in figure 4. Since it is hard to construct a meaningful order, the most suitable option is to adopt the multinomial logit model, where the probability of a category is computed in relation to a selected base category. We use inconsistent de jure fixed regimes as the base category. For both estimation methods, we use the same set of covariates to identify regime determinants, based on the following set of conditional probabilities:

$$\text{(1)} \qquad \mathrm{P}\left[Y_{it} = y \,\middle|\, \mathbf{X}_{it}, \mathbf{GovCh}\left(\mathbf{q}^{-}\right)_{\mathbf{it}}^{\mathbf{q}}, \mathbf{GovCh}\left(\mathbf{q}^{+}\right)_{\mathbf{it}}^{\mathbf{q}}\right],$$

where i and t stand for country and month, respectively. For the ordered logit model, where the categories present a clear rank order, the dependent variable is either the de jure or the de facto exchange rate regime, and y takes a value of 1, 2, 3, or 4 if the regime is fixed, crawl, managed, or float, respectively. For the multinomial logit model, where the categories are not ruled by any apparent rank order, the dependent variable y takes a value of 1, 2, 3, or 4 if the de jure regime is consistent fixed, inconsistent flexible, inconsistent fixed, or consistent flexible, respectively. Government change is represented by two matrices of four dummy variables each, accounting for the year before and after the change, which occurs in month 0. Although the data are monthly, we define the dummy variables by quarters (the superscript q stands for quarter). Thus,

$$\mathbf{GovCh}\left(\mathbf{q}^{-}\right)_{\mathbf{it}}^{\mathbf{q}} = \left[GovCh(-3)_{it}^{q}\ GovCh(-2)_{it}^{q}\ GovCh(-1)_{it}^{q}\ GovCh(0)_{it}^{q}\right],$$

where $GovCh(0)^q$ takes a value of one in the months 0 to 2 before the government change month, $GovCh(-1)^q$ takes a value of one in the months 3 to 5 before the government change month, and so on for $GovCh(-2)^q$ and $GovCh(-3)^q$. Analogously,

$$\mathbf{GovCh}\left(\mathbf{q}^{+}\right)_{\mathbf{it}}^{\mathbf{q}} = \left[GovCh(+1)_{it}^{q}\ GovCh(+2)_{it}^{q}\ GovCh(+3)_{it}^{q}\ GovCh(+4)_{it}^{q}\right]$$

is constructed for the twelve months following the change in government using four quarterly dummy variables.

X is a matrix composed of seven time-varying controls: (1) *Portfolio*: the sum of the absolute value of inward and outward flows of portfolio investment and financial derivatives as a percentage of GDP, from the IMF International Financial Statistics (IFS) database. Levy-Yeyati, Sturzenegger, and Reggio (2010) use this variable as a proxy for capital mobility. Given the impossible trinity, policymakers must give up on either monetary policy or exchange rate policy in environments with high capital mobility, which makes intermediate regimes less viable. Alternatively, under a currency mismatch argument, we should expect more commitments to pegging. (2) *Foreign.Liab.pc*:

foreign liabilities per capita, from the IMF IFS database. Countries with substantial foreign liabilities may be prone to fix their currency, since sharp nominal depreciation affects the solvency of the nontradable sector's balance sheets. Alesina and Wagner (2006) and Levy-Yeyati, Sturzenegger, and Reggio (2010) use foreign liabilities over monetary aggregates instead. However, the problem with this variable is that for Latin American countries, money demand was extremely unstable in the 1980s and early 1990s owing to high inflation. In crisis episodes with high inflation, money demand falls while the monetary authority lets the exchange rate float, creating a positive relation between foreign liabilities and flexible regimes, totally opposite to the currency mismatch hypothesis.[9] (3) *Size*: real GDP in dollars, from the IMF IFS database. As noted by Levy-Yeyati, Sturzenegger, and Reggio (2010), smallness favors a more stable exchange rate because small economies are more likely to trade internationally than economies with a large domestic market and because it limits the scope for the use of a national unit of account. (4) *ToT:* terms of trade. When the terms of trade are high, Latin American countries tend to fix their exchange rates as a device for accumulating international reserves in their central banks, probably to insure against sudden stops (Jeanne, 2007; Jeanne and Rancière, 2011). (5) *U.S.Interest*: the U.S. interest rate in real terms, from the IMF IFS database. Calvo, Leiderman, and Reinhart (1993) and Fernández-Arias and Montiel (1996) find that the U.S. interest rate is a determinant of capital inflows in Latin America.[10] When the U.S. interest rate increases, capital outflows may be stopped by letting the exchange rate float. This effect should be exacerbated when economies keep more open capital accounts. (6) *Openness*: exports plus imports over GDP, from the IMF IFS database. The decision to peg could be correlated with trade openness since highly open economies are in favor of a more stable exchange rate, as Levy-Yeyati, Sturzenegger, and Reggio (2010) note. (7) *Default*: a dummy variable that takes a value of one if the country has defaulted on its external debt and zero otherwise, from Carmen M. Reinhart's website. This variable is used to control for the fact that economies characterized by high macroeconomic instability cannot sustain their currency, so they let their currency float or, more precisely, freely fall.

9. We indeed find a significant positive coefficient when we use foreign liabilities over money, so the probability of a flexible regime increases when foreign liabilities to money increase. In contrast, there is a negative coefficient with our transformation of foreign liabilities normalized by population. The latter is consistent with the currency mismatch hypothesis as found in Levy-Yeyati, Sturzenegger, and Reggio (2010) for their regime classification. Results are shown below.

10. When the U.S. Treasury bill rate is used instead, results are qualitatively the same.

Among the seven controls described above, five are available only at annual frequencies. These are *Portfolio*, *Foreign.Liab.pc*, *Size*, *ToT*, and *Default*. For the first four, we use the log differential method to construct within-year imputation with constant monthly percentage change within each year. *Default* is left at its annual frequency insofar as it is a dummy variable. The remaining two, *U.S.Interest* and *Openness,* are available at monthly frequencies, so interpolation is not necessary. Given the possibility of reverse causality, we decided to use one-month lagged values of the variables available at monthly frequency. For the variables available at annual frequency that were interpolated using log differences, we use twelve-month lagged values instead. All variables are expressed in natural logs except *Default* and dummy variables for government change.

The estimations of equation 1 under the IMF de jure and the RR de facto exchange rate regimes for ordered logit models are shown in table 1, together with the results for the multinomial logit model. Results of the ordered and multinomial logits have different interpretations. For the former, a positive (negative) estimator indicates that the probability of a more flexible (fixed) regime increases if the corresponding covariate shows a marginal increase, but the estimator does not predict at first sight what happens with the probabilities of the middle categories.[11] For the latter, each coefficient is understood as the increase in the probability of category $j = 1, 2, 4$ in relation to the base category (3) for a marginal increase of the independent variable, if its coefficient is positive.

Here we first focus on the econometric results in table 1 for each of the covariates and relate them to the well-known literature on the de jure and de facto regime determinants.[12] Our innovation in relation to the literature is the novel regime classification, which identifies consistent and inconsistent de jure regimes (results are displayed in columns 3–5 of table 1). We then focus on the issue of endogeneity of regimes around the government change date.

For the RR classification in column 2, the probability of observing de facto fixed regimes tends to increase as the de facto capital account openness increases (that is, *Portfolio* = –0.045** in column 2). This is consistent with the currency mismatch hypothesis of Levy-Yeyati, Sturzenegger, and Reggio (2010). At the same time, de jure flexible regimes also tend to increase as the de facto

11. See online appendix B for the full set of marginal effects for each of the categories using mean values of covariates for both the ordered and multinomial logit models.

12. The connection between regression coefficients and changes in probabilities is detailed in online appendix B.

TABLE 1. Determinants of Exchange Rate Regimes: Ordered and Multinomial Logit Models

	Ordered logit		*Multinomial logit relative to inconsistent de jure fixed*		
Explanatory variable	*IMF (de jure) (1)*	*RR (de facto) (2)*	*Consistent de jure fixed (3)*	*Inconsistent de jure flexible (4)*	*Consistent de jure flexible (5)*
ln $Portfolio_{t-12}$	0.046**	−0.045**	0.166***	0.168***	0.104***
	(0.021)	(0.019)	(0.033)	(0.043)	(0.030)
ln $Foreign.Liab.pc_{t-12}$	0.012	−0.049***	0.265***	0.201***	0.297***
	(0.009)	(0.009)	(0.016)	(0.022)	(0.018)
ln $Size_{t-12}$	−0.006	−0.041	1.403***	1.172***	1.576***
	(0.041)	(0.040)	(0.105)	(0.110)	(0.104)
ln ToT_{t-12}	−0.656	−2.564***	4.141***	10.357***	3.193**
	(0.485)	(0.438)	(1.383)	(1.379)	(1.310)
ln $U.S.Interest_{t-1}$	−2.333***	−0.017	−0.897***	−4.550***	−2.517***
	(0.135)	(0.124)	(0.219)	(0.248)	(0.237)
ln $Openness_{t-1}$	0.464***	−0.297***	0.883***	0.665***	1.331***
	(0.109)	(0.105)	(0.222)	(0.232)	(0.227)
$Default_{t-12}$	0.588***	1.573***	−0.660***	−0.493*	1.604***
	(0.100)	(0.099)	(0.227)	(0.273)	(0.221)
$GovCh(-3)^{q}$	0.048	0.297*	−0.025	0.040	0.481
	(0.166)	(0.160)	(0.383)	(0.414)	(0.395)
$GovCh(-2)^{q}$	0.125	0.434***	−0.434	−0.176	0.288
	(0.169)	(0.163)	(0.341)	(0.379)	(0.351)
$GovCh(-1)^{q}$	0.163	0.449***	−0.503	−0.326	0.228
	(0.170)	(0.163)	(0.321)	(0.372)	(0.327)
$GovCh(0)^{q}$	0.201	0.464***	−0.530	−0.227	0.213
	(0.172)	(0.166)	(0.325)	(0.366)	(0.326)
$GovCh(+1)^{q}$	0.595***	0.630***	0.183	0.229	0.985**
	(0.172)	(0.167)	(0.397)	(0.431)	(0.387)
$GovCh(+2)^{q}$	0.483***	0.496***	0.246	0.341	0.838**
	(0.169)	(0.162)	(0.399)	(0.421)	(0.392)
$GovCh(+3)^{q}$	0.346**	0.322**	0.249	0.020	0.411
	(0.169)	(0.158)	(0.378)	(0.414)	(0.386)
$GovCh(+4)^{q}$	0.396**	0.255	0.106	0.080	0.080
	(0.172)	(0.160)	(0.375)	(0.410)	(0.376)
No. observations	2,557	2,592	2,662	2,662	2,662

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: Columns 1 and 2 show the results of the estimation of equation 1 with ordered logit models, where the dependent variable equals 1, 2, 3, or 4 if the regime is fixed, crawl, managed, or float, respectively. Columns 3, 4, and 5 show the results of the estimation of equation 1 with multinomial logit, where the dependent variable equals 1, 2, 3, or 4 if the de jure regime is consistent fixed, inconsistent flexible, inconsistent fixed, or consistent flexible, respectively; the results are relative to the inconsistent fixed category. Reduced sample includes Argentina, Bolivia, Brazil, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Mexico, Nicaragua, Peru, Uruguay, and Venezuela over the 1980–99 period. Nondemocratic episodes are excluded, based on the Polity IV Project. Dollarization episodes are also excluded. Robust standard errors are in parentheses.

capital account openness increases (that is, *Portfolio* = 0.046** in column 1). Taken together, these results may be suggestive of an increase in inconsistent flexible regimes. In columns 3 to 5, we find that the facto fixed regimes, including both the consistent fixed and inconsistent flexible versions, are more likely relative to inconsistent fixed regimes (that is, 0.166*** and 0.168***, respectively). Moreover, inconsistent fixed regimes are less likely since all three coefficients are positive (that is, all regimes are more likely in relation to the base category).

The coefficient for foreign liabilities per capita (*Foreign.Liab.pc*) is close to zero and insignificant for the IMF de jure classification in column 1, but it is significantly negative in the de facto classification in column 2 (that is, *Foreign.Liab.pc* = –0.049***), which is consistent with the currency mismatch hypothesis. The multinomial logit model corroborates this finding: both types of fixed regimes—consistent fixed (column 3) and inconsistent flexible (column 4)—are more likely relative to inconsistent fixed regimes (that is, 0.265*** and 0.201***, respectively). Consistent flexible regimes are also more likely relative to inconsistent fixed regimes (that is, 0.297*** in column 5). This suggests that inconsistent fixed regimes do not go hand in hand with liability dollarization, mainly because those are episodes of high macroeconomic instability. *Size* is insignificant in both columns 1 and 2, while we would have expected it to be positive at least for the de facto classification. However, in line with Levy-Yeyati, Sturzenegger, and Reggio (2010), the multinomial logit model finds that the consistent flexible category increases its likelihood the most, insofar as its estimator is the greatest of the three. This indicates that flexible regimes are indeed more likely in bigger countries, while inconsistent fixed regimes become less likely because the three estimators are positive. *ToT* has the predicted negative sign in the de facto classification of column 2, while in the de jure classification it is close to zero and insignificant. In columns 3–5, the two de facto fixed regimes—that is, consistent fixed (column 3) and inconsistent flexible (column 4)—become more likely when terms of trade increase. This is consistent with the strategy of pegging the exchange rate to acquire international reserves as, probably, an insurance device, as found in Jeanne (2007) and Jeanne and Rancière (2011). In addition, since all three estimators are positive, it indicates that an inconsistent fixed regime becomes less likely. This is probably because the increase in terms of trade tends to create a trade balance surplus that increases the supply of foreign currency, which may alleviate exchange rate pressures during high macroeconomic instability.

U.S.Interest is close to zero and insignificant in the de facto regime (column 2). For the de jure regime (column 1), the likelihood of a peg increases strongly as *U.S.Interest* increases, since the estimator is significant and negative. Altogether, this evidence might indicate the increase of de jure fixed regimes that cannot be sustained in the medium to short run, that is to say, inconsistent fixed regimes. This seems to occur since an increase in the U.S. interest rate produces capital outflows from the Latin American region, as found in Calvo, Leiderman, and Reinhart (1993) and Fernández-Arias and Montiel (1996). The de jure regime may try to signal stability as an attempt to control the market instability with mere words. The multinomial logit model corroborates this view: all three estimators are significantly negative, indicating that the likelihood of inconsistent fixed regimes increases when the U.S. interest rate increases.

Openness has the predicted negative sign in the de facto classification (column 2) (that is, economies that are more open prefer a more stable exchange rate). However, the de jure regime (column 1) is significantly positive. According to the multinomial logit results, inconsistent fixed regimes are less likely when *Openness* increases, while consistent fixed and flexible regimes become much more likely. This suggests that open economies are more compatible with macroeconomic strength. The market-based exchange rate tends to float when economies default on their debt (that is, *Default* = 1.573*** in column 2), while the de jure regime keeps pace with the market behavior (that is, *Default* = 0.588*** in column 1). Default definitely decreases the probability of de facto fixed regimes, including both the consistent fixed and inconsistent flexible versions, in relation to inconsistent fixed regimes (that is, *Default* = –0.660*** in column 3; *Default* = –0.493*** in column 4). Consistent flexible regimes become more likely (column 5 is the only positive coefficient), which is congruent with the findings in columns 1 and 2 (that is, de jure and de facto regimes become more flexible and flexible regimes become more consistent).

Now we move on to the issue of endogeneity of regimes around the government change date. The de jure regime does not seem to change in the four quarters leading up to a government change since $GovCh(-3)^q$, $GovCh(-2)^q$, $GovCh(-1)^q$, and $GovCh(0)^q$ are not significant in column 1. This is important since it indicates that de jure regimes are not likely to be strongly affected by the endogeneity of regime announcements.[13] After government changes,

13. In online appendix B, the marginal effects of $GovCh(-3)^q$, $GovCh(-2)^q$, $GovCh(-1)^q$, and $GovCh(0)^q$ are small and insignificant as well.

estimators $GovCh(+1)^q$, $GovCh(+2)^q$, and $GovCh(+3)^q$ in column 1 are significantly positive, indicating that the new government tends to announce more floating.[14] As to the de facto classification, the exchange rate tends to be more flexible both before and after the government change date, since all the government change estimators are significantly positive, with the exception of $GovCh(+4)^q$.

When we use our novel regime classification, the results indicate that no regime is more likely than the baseline (inconsistent de jure fixed regimes) before government changes. This is consistent with the analysis in figure 4 above, in which the regime duration distribution is quite similar for all the regimes in the twelve months before either government changes or a randomly generated month 0. After government changes, the probability of consistent flexible regimes increases relative to inconsistent fixed regimes in the first two quarters: $GovCh(+1)^q = 0.985$*** and $GovCh(+2)^q = 0.838$**. This indicates that the monetary authority announces a flexible regime in the first few months after a government change in an already de facto flexible environment. This is in line with the sharp, sudden drop of inconsistent de jure fixed regimes right after government changes in panel A of figure 4, which contrasts with the behavior after a randomly selected month 0.

The Dynamics of the Real Exchange Rate

Having found no statistical evidence that exchange rate regime announcements vary before government changes, we study the dynamics of the real exchange rate around government changes conditional on consistent/inconsistent de jure regimes. We use a dynamic distributed lag model of the following form:

$$
\begin{aligned}
(2)\qquad \Delta \ln\left(\mathrm{RER}_{it}\right) = {} & \sum\nolimits_{k=1}^{3} a_k \Delta \ln\left(\mathrm{RER}_{i,t-k}\right) + \Delta\mathbf{W}_{\mathbf{it}}\boldsymbol{\beta} + \mathbf{GovCh}_{\mathbf{it}}\boldsymbol{\delta} \\
& + \mathbf{GovChFI}_{\mathbf{it}}\boldsymbol{\delta}_{\mathbf{FI}} + \mathbf{GovChFEI}_{\mathbf{it}}\boldsymbol{\delta}_{\mathbf{FEI}} \\
& + \mathbf{GovChFEC}_{\mathbf{it}}\boldsymbol{\delta}_{\mathbf{FEC}} + \varepsilon_{it},
\end{aligned}
$$

where i and t stand for country and month. The dependent variable is the log difference of the real exchange rate. We control for three distributed lags to

14. These results are in line with Klein and Marion (1997) and Gavin and Perotti (1997) and are also consistent with the pattern found by Blomberg, Frieden, and Stein (2005), where the duration of pegs increases before elections and decreases afterward.

capture persistency.[15] Government change is represented as a matrix of quarterly dummy variables, where $GovCh(\pm l)^q$ takes a value of one if the government change is $\pm l$ quarters away:

$$\mathbf{GovCh_{it}} = \Big[GovCh(-3)^q_{it}\; GovCh(-2)^q_{it}\; GovCh(-1)^q_{it}\; GovCh(0)^q_{it}$$
$$GovCh(+1)^q_{it}\; GovCh(+2)^q_{it}\; GovCh(+3)^q_{it}\; GovCh(+4)^q_{it} \Big].$$

The analogous matrices **GovChFI**, **GovChFEI**, and **GovChFEC** capture the interaction between **GovCh** and inconsistent de jure fixed, inconsistent de jure flexible, and consistent de jure flexible regimes, respectively; the omitted category is consistent de jure fixed regimes. The regime classification used for the entire electoral window is invariant and equal to the classification at the month before the elections, which is typically two or three months before the government change.[16] **W** is a matrix of time-varying controls that attempt to control for the determinants of both exchange rate dynamics and regime announcement. In that regard, we use the same set of variables employed in the estimation of equation 1 to control for the determinants of regime announcement, namely, *Portfolio*, *Foreign.Liab.pc*, *Size, ToT*, *U.S.Interest*, *Openness*, and *Default*. Insofar as an expansion in the size of government will induce an appreciation of the real exchange rate when government demand is biased toward nontradable goods, as stressed by Goldfajn and Valdés (1999), we add government expenditure as a ratio of GDP, *GovSize*. Because we could not corroborate that our regressors produce a cointegrating vector, we estimate the model in first differences, following Cermeño, Grier, and Grier (2010).[17] However, our results do not change significantly when we study equation 2 in levels.[18] Finally, given the possibility of reverse causality, we use one-month lagged values of the variables in **W**. For the variables available at annual frequency that were interpolated using log differences, we use twelve-month lagged

15. Results are totally invariant to the inclusion of one lag. Results with one lag are available on request.

16. Results are virtually unchanged when we use the value six months before elections instead. Results under the latter are available on request.

17. We ran Engle-Granger tests for each country, and in almost all the countries the hypothesis of cointegration was rejected. Only Argentina, Bolivia, Guatemala, Honduras, and Uruguay showed evidence of cointegration at 5 percent significance or higher. Test results are available on request.

18. Results are available on request.

values instead. The rest of the variables are in natural logs except for *Default* and the dummy variables for government change. Results are displayed in table 2. Column 1 shows the results of the estimation of equation 2 for a set of quarterly dummy variables that indicates the proximity of government change date up to four quarters before and after the change. Column 2 replicates column 1 without using covariates, while column 3 replicates column 2 for the extended sample, which includes those countries that we lost owing to covariate limitations.

In column 1 of table 2, the real exchange rate decreases (that is, appreciates) moderately during the last quarter before the government change for inconsistent de jure fixed regimes, but the result is not significant, with $GovChFI(0)^q = -4.832$; after the government change, the real exchange rate depreciates 17 percent in the first quarter, with $GovChFI(+1)^q = 17.312$*. Linear combination 1, which shows the difference estimator for the two quarters, is not statistically significant, but linear combinations 2, 3, and 4, which progressively extend the window to cover the second, third, and fourth quarters around the government change date, do capture statistically significant depreciation differentials of 16, 11, and 9 percent, respectively; this undoubtedly reflects the significant depreciation during the second quarter, when $GovChFI(+2)^q =$ 9.469*. Column 2, which excludes time-varying covariates, is almost the same as column 1. This indicates that the empirical design produces a plausible exogenous variation of regime adoptions, as shown in figure 4.[19] When we use the extended sample in column 3, the effects drop substantially, but they are all statistically significant.[20]

Real Exchange Rate Misalignments around Government Changes

In the previous section we studied the short- and medium-term dynamics of the real exchange rate (RER) and found that there is a slight and insignificant appreciation quarter to quarter during the year leading up to the government change under inconsistent de jure fixed regimes, followed by a strong and significant depreciation after the change of government. In this section, we explicitly

19. Appendix D1 shows that the results are also robust to allowing for conditional heteroskedasticity.

20. The linear combinations for the countries that are only in the extended sample, though smaller in magnitude, are also positive and statistically significant. For example, linear combination 4 is 9.371** for the reduced sample, 4.932** for the extended sample, and 2.268*** for the extra countries. These regression results are available on request.

TABLE 2. Exchange Rate Variation around Government Changes

Explanatory variable	*Reduced sample, with covariates (1)*		*Reduced sample, no covariates (2)*		*Extended sample, no covariates (3)*	
$GovCh(-3)^q$	−0.358	(0.392)	−0.208	(0.374)	−0.116	(0.286)
$GovCh(-2)^q$	0.417	(0.277)	0.586*	(0.289)	0.362	(0.264)
$GovCh(-1)^q$	−0.540*	(0.281)	−0.328	(0.295)	−0.508**	(0.225)
$GovCh(0)^q$	−0.070	(0.649)	0.222	(0.583)	0.911	(0.790)
$GovCh(+1)^q$	0.697	(0.624)	0.706	(0.642)	0.019	(0.402)
$GovCh(+2)^q$	0.148	(0.281)	0.109	(0.190)	−0.271	(0.283)
$GovCh(+3)^q$	0.028	(0.357)	0.233	(0.352)	0.272	(0.250)
$GovCh(+4)^q$	−0.376	(0.441)	−0.322	(0.419)	0.221	(0.413)
$GovChFI(-3)^q$	−2.647	(2.526)	−2.757	(2.705)	−1.181	(1.512)
$GovChFI(-2)^q$	−2.754	(2.960)	−3.027	(2.592)	−2.015*	(1.113)
$GovChFI(-1)^q$	0.470	(0.833)	−0.130	(0.969)	−0.964	(1.084)
$GovChFI(0)^q$	−4.832	(6.787)	−5.060	(6.503)	−3.416	(2.405)
$GovChFI(+1)^q$	17.312*	(9.345)	16.899*	(8.908)	9.054*	(4.837)
$GovChFI(+2)^q$	9.469*	(5.188)	8.665	(5.227)	2.499	(2.771)
$GovChFI(+3)^q$	0.525	(1.035)	0.389	(1.003)	−0.413	(0.916)
$GovChFI(+4)^q$	0.663	(0.942)	0.556	(0.726)	1.013	(0.996)
$GovChFEI(-3)^q$	0.818	(0.592)	0.897	(0.517)	−0.112	(0.500)
$GovChFEI(-2)^q$	−0.367	(0.447)	−0.360	(0.406)	−0.727***	(0.217)
$GovChFEI(-1)^q$	1.331**	(0.507)	1.061**	(0.373)	0.595*	(0.310)
$GovChFEI(0)^q$	−0.375	(0.669)	−0.613	(0.592)	−1.229	(0.789)
$GovChFEI(+1)^q$	−0.623	(0.694)	−0.554	(0.693)	−0.002	(0.455)
$GovChFEI(+2)^q$	−0.891	(0.957)	−0.300	(0.350)	0.119	(0.324)
$GovChFEI(+3)^q$	−0.636	(0.431)	−0.510	(0.474)	−0.142	(0.357)
$GovChFEI(+4)^q$	0.364	(0.598)	0.529	(0.513)	−0.095	(0.487)
$GovChFEC(-3)^q$	−0.814	(1.068)	−1.020	(1.083)	−0.921	(0.782)
$GovChFEC(-2)^q$	−0.626	(1.328)	−0.813	(1.437)	−0.360	(0.933)
$GovChFEC(-1)^q$	1.545	(1.637)	1.426	(1.579)	1.314	(1.187)
$GovChFEC(0)^q$	0.151	(1.038)	−0.241	(1.077)	−0.203	(1.235)
$GovChFEC(+1)^q$	0.543	(1.572)	0.509	(1.575)	0.897	(1.068)
$GovChFEC(+2)^q$	−0.425	(0.643)	−0.524	(0.606)	−0.522	(0.559)
$GovChFEC(+3)^q$	0.293	(0.652)	−0.112	(0.672)	−0.394	(0.587)
$GovChFEC(+4)^q$	0.843	(0.636)	0.463	(0.684)	−0.520	(0.546)
No. observations	2,236		2,236		4,008	
R^2	0.083		0.056		0.028	
Linear combination 1	22.14	(15.150)	21.96	(14.35)	12.47**	(6.498)
Linear combination 2	15.57**	(7.190)	15.38**	(7.070)	7.967**	(3.650)
Linear combination 3	11.47**	(5.766)	11.39**	(5.605)	5.845**	(2.775)
Linear combination 4	9.433**	(4.761)	9.371**	(4.766)	4.932**	(2.226)

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: OLS estimation of equation 2. The dependent variable is $\Delta \ln$ RER. Reduced sample includes Argentina, Bolivia, Brazil, Colombia, Costa Rica, Dominican Republic, Ecuador, Guatemala, Honduras, Mexico, Nicaragua, Panama, Peru, Uruguay, and Venezuela over 1980–99 period. Extended sample also includes Barbados, Chile, El Salvador, Guyana, Jamaica, and Paraguay. *FI*, *FEI*, and *FEC* stand for inconsistent fixed, inconsistent flexible, and consistent flexible regimes; consistent flexible is the omitted category. Nondemocratic episodes are excluded, based on Polity IV Project. Dollarization episodes are also excluded. Controls are used, but not reported (see text). Linear combination $k = 1, 2, 3, 4$: $GovChFI(+1)^q - GovChFI(0)^q$, $\frac{1}{2}(GovChFI(+2)^q + GovChFI(+1)^q - GovChFI(0)^q - GovChFI(-1)^q)$, $\frac{1}{3}(GovChFI(+3)^q + GovChFI(+2)^q + GovChFI(+1)^q - GovChFI(0)^q - GovChFI(-1)^q - GovChFI(-2)^q)$, $\frac{1}{4}(GovChFI(+4)^q + GovChFI(+3)^q + GovChFI(+2)^q + GovChhFI(+1)^q - GovChFI(0)^q - GovChFI(-1)^q - GovChFI(-2)^q - GovChFI(-3)^q)$. Robust standard errors are in parentheses to the right of each estimator.

study the RER misalignment caused by pegging the exchange rate when it is not consistent with the market exchange rate, following the analysis in Goldfajn and Valdés (1999). We control for the stochastic trends of the exchange rate by applying the Hodrick-Prescott filter country by country to each series.[21] The series are then decomposed into two components:

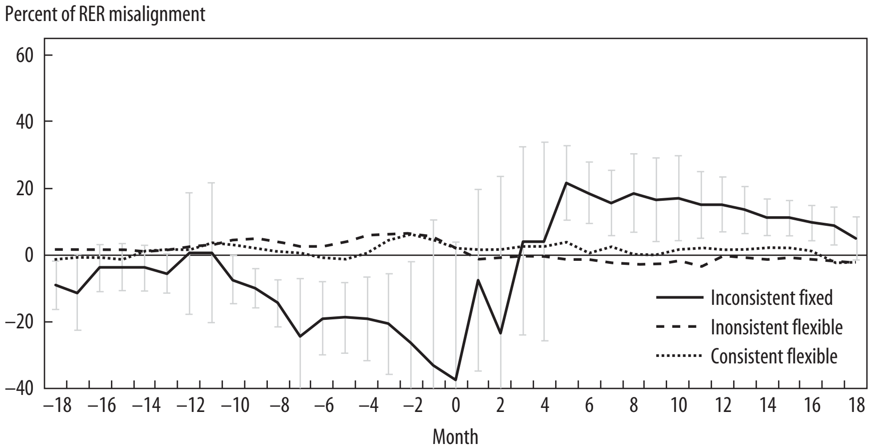
$$\ln\left(\mathrm{RER}_{it}\right) = \ln\left(\mathrm{RER}_{it}\right)_{cycle} + \ln\left(\mathrm{RER}_{it}\right)_{trend}. \tag{3}$$

We identify the trend component as the long-run equilibrium RER and the cycle as departures from that equilibrium. When the cyclical component is negative, the RER is overvalued; when it is positive, it is undervalued. Goldfajn and Valdés (1999) identify four appreciation phases of the RER: history, when the appreciation hits 5 percent; start, when the appreciation hits a threshold (for example, 10 percent, 15 percent); peak, when the appreciation reaches its highest value; and end, when the appreciation is back to the 5 percent history level, which is considered a statistical reversion of the appreciation process. We use this classification to identify when an appreciation represents a significant overvaluation of the exchange rate, in this case, 5 percent and above. The advantage of using logs is that $\ln(\mathrm{RER}_{it})_{cycle}$ already represents the percentage of overvaluation (below the trend) or undervaluation (above the trend). We then estimate the following equation using ordinary least squares (OLS):

$$\begin{aligned} \ln\left(\mathrm{RER}_{it}\right)_{cycle} = \alpha &+ \sum_{i=0}^{18} GovCh(-i)_{it}^{m} \bullet \beta_{(-i)} + \sum_{i=1}^{18} GovCh(+i)_{it}^{m} \bullet \beta_{(+i)} \\ &+ \sum_{i=0}^{18} GovChFI(-i)_{it}^{m} \bullet \beta_{FI(-i)} \\ &+ \sum_{i=1}^{18} GovChFI(+i)_{it}^{m} \bullet \beta_{FI(+i)} \\ &+ \sum_{i=0}^{18} GovChFEI(-i)_{it}^{m} \bullet \beta_{FEI(-i)} \\ &+ \sum_{i=1}^{18} GovChFEI(+i)_{it}^{m} \bullet \beta_{FEI(+i)} \\ &+ \sum_{i=0}^{18} GovChFEC(-i)_{it}^{m} \bullet \beta_{FEC(-i)} \\ &+ \sum_{i=1}^{18} GovChFEC(+i)_{it}^{m} \bullet \beta_{FEC(+i)} + \varepsilon_{it}. \end{aligned} \tag{4}$$

21. To filter the RER series, we use a smoothing parameter of 129,600, which is the value Ravn and Uhlig (2002) suggest for monthly data. Since the Hodrick-Prescott filter usually introduces a spurious dynamic relation into the series, we also use the Hamilton (2018) filtering technique. The results are qualitatively the same; see online appendix C.

FIGURE 5. Real Exchange Rate Misalignments around Government Changes



Notes: Graphic representation of the estimators of equation 5 for the cyclical component of the RER for the reduced sample, where the RER series is detrended using the Hodrick-Prescott filter (smoothing parameter of 129,600). Results are relative to consistent fixed episodes. Month 0 is the government change month. Vertical bars represent 95 percent confidence intervals of estimators based on robust standard errors for inconsistent fixed estimators around government change date.

In this particular case, we use monthly, rather than quarterly, dummy variables to identify precisely the months at which the overvaluation begins and when it reverts. Afterward, we collapse these monthly dummy variables into six-month periods to observe medium-run misalignments. For the sake of presentation, results of the monthly dummy variables are shown in figure 5, while the half-yearly dummy variables are shown in table 3.

Figure 5 shows that a significant overvaluation occurs only for the inconsistent de jure fixed regime announcement. The 5 percent history threshold is reached at month 10 before the government change, and a peak of 37 percent is reached at the government change month, that is, the history/peak stage lasts ten months, while the peak/end period lasts only three and is mostly completed in the first month. After the government change date, there is a process of undervaluation, which becomes significant at month 5 (undervaluation of 22 percent), but the process reverts smoothly in 14 months, when the RER reaches its equilibrium (that is, back to below 5 percent of undervaluation). Notably, when the exchange rate is overvalued, a quick one-month correction is observed, which indicates that this is achieved through a strong nominal devaluation, as highlighted in Goldfajn and Valdés (1999).

TABLE 3. Real Exchange Rate Misalignments Using Six-Month Dummy Variables

Explanatory variable	*(1)*	
$GovCh(-2)^s$	−1.513	(1.008)
$GovCh(-1)^s$	−2.688***	(0.962)
$GovCh(0)^s$	−3.041***	(0.934)
$GovCh(+1)^s$	0.975	(1.070)
$GovCh(+2)^s$	1.002	(1.035)
$GovCh(+3)^s$	−0.407	(0.926)
$GovChFI(-2)^s$	−4.551**	(2.191)
$GovChFI(-1)^s$	−12.344***	(3.038)
$GovChFI(0)^s$	−25.498***	(5.556)
$GovChFI(+1)^s$	2.981	(6.359)
$GovChFI(+2)^s$	16.292***	(2.226)
$GovChFI(+3)^s$	9.765***	(1.305)
$GovChFEI(-2)^s$	1.733	(1.090)
$GovChFEI(-1)^s$	3.734***	(1.004)
$GovChFEI(0)^s$	5.086***	(1.012)
$GovChFEI(+1)^s$	−0.943	(1.046)
$GovChFEI(+2)^s$	−2.255**	(1.128)
$GovChFEI(+3)^s$	−1.233	(1.021)
$GovChFEC(-2)^s$	0.429	(1.283)
$GovChFEC(-1)^s$	1.664	(1.582)
$GovChFEC(0)^s$	2.831	(2.096)
$GovChFEC(+1)^s$	2.111	(1.710)
$GovChFEC(+2)^s$	1.381	(1.474)
$GovChFEC(+3)^s$	0.449	(1.295)
No. observations	2,127	
R^2	0.119	

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: OLS estimation of equation 5 for the cyclical component of RER using six-month dummy variables and the reduced sample, where the RER series is detrended using the Hodrick-Prescott filter (smoothing parameter of 129,600). The dependent variable is $\ln(RER_{it})_{cycle}$. Results are relative to consistent fixed episodes. Nondemocratic episodes are excluded, based on Polity IV Project. Dollarization episodes also excluded. Robust standard errors are in parentheses to the right of each estimator.

However, when the exchange rate is undervalued, as in month +5, the correction takes place smoothly through either a gradual correction of the nominal exchange rate, which corrects the initial overshooting that brought about the undervaluation, or an organized correction of inflation differentials. This difference in the RER reversion in the appreciation and depreciation phases is not treated in Goldfajn and Valdés (1999). Hence, our paper highlights the large asymmetries of the reversions that occur in the overvaluation and undervaluation phases. Furthermore, Goldfajn and Valdés (1999) identify appreciation dynamics without characterizing and describing the context in which these appreciations take place. We identify one particular context where these appreciations occur: elections that coincide with a poor macroeconomic performance.

Finally, table 3 shows the collapsed six-month dummy variables to capture average medium-term behavior. For the inconsistent de jure fixed regimes, the six-month average overvaluation reaches 25 percent in the six months before the government change, which reverts in the first six months after government change, followed by a 16 percent undervaluation in the second six-month period, which then declines to 9 percent in the third period.[22]

Comparison of the 1980–99 and 2000–16 Periods

In this section, we first compare the characteristics of exchange rate regimes for the 1980–99 and 2000–16 periods. We then examine the 2015 Argentine general elections, which share the underlying characteristic of inconsistent de jure fixed regimes, and compare the case study with the econometric findings of the earlier period.

There are eighty-one changes of government in the 1980–99 period and eighty-four in the 2000–16 period. Table 4 shows that de facto flexible regimes are much less common in the recent period, falling from 57 to 30 percent of total cases from one period to the next. While both de jure and de facto classifications are available for the earlier period, only the de facto classification is available for the later period. Nevertheless, we can use the typical characteristics of inconsistent de jure fixed regimes in the earlier period to draw parallels for the more recent period. For 1980–99, two common shared characteristics of inconsistent fixed regimes, where the announced fixed regime does not coincide with the actual policy, are dual/multiple markets and high inflation (more than 10 percent a year). Dual markets and high inflation characterize 81 percent of the inconsistent fixed regime cases in the earlier period, as well as 73 percent of the consistent flexible regime cases, while within the total

22. We also produce both the figure and the table of RER misalignments using the Hamilton (2018) filter. The results, shown in online appendix C, do not change significantly. The results are also robust to allowing for conditional heteroskedasticity; see appendix D2.

TABLE 4. Regime Classification, 1980–2016: Number of Episodes at Government Change Date

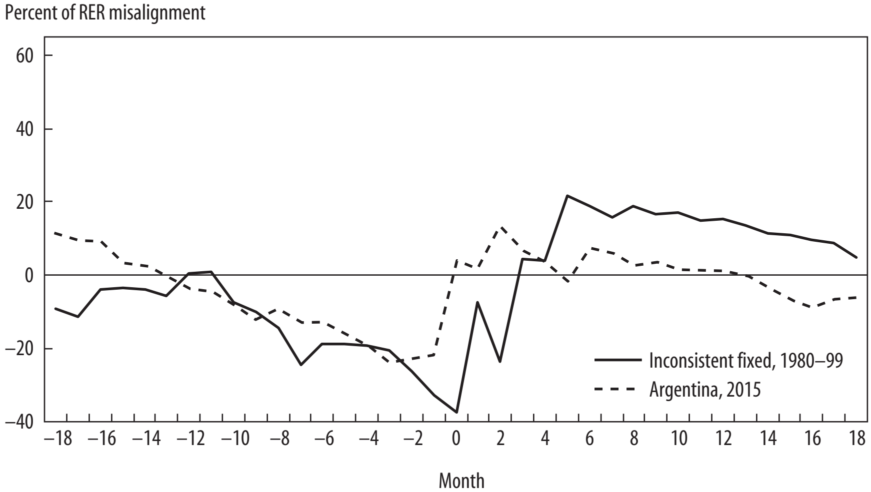
Period and regime	*Total cases*	*Cases with dual markets*	*Cases with annual inflation > 10%*	*Cases with dual markets and annual inflation > 10%*
1980–99 period				
De facto flexible				
Inconsistent de jure fixed	16	16	13	13
Consistent de jure flexible	30	23	27	22
Total de facto flexible	46	39	40	35
De facto fixed				
Inconsistent de jure flexible	24	8	14	6
Consistent de jure fixed	7	2	3	2
Total de facto fixed	31	10	17	8
2000–16 period				
Total de facto flexible	25	1	20	1
Total de facto fixed	49	3	33	2

Note: Dollarization episodes are excluded.

de facto flexible cases, 76 percent involve dual markets and high inflation. This proportion drops to only 4 percent in the later period, and 4 percent represents a single case. Hence, although we cannot observe the IMF de jure classification after 1999, the fact that dual/multiple markets and high inflation become exceptional after 1999 provides indirect evidence that inconsistent fixed regimes are no longer likely.

The only case that we can identify as likely to have been classified as an inconsistent de jure fixed regime—because it shows de facto flexible behavior, dual exchange rates, and high inflation—is the 2015 presidential election of Argentina. The Central Bank of Argentina was applying a crawling peg, devaluing around 1 percent per month from January to November 2015. The incumbent government also announced that the policy would continue the following year, selling huge amounts of future dollar contracts at prices consistent with that crawling peg. Figure 6 compares the real exchange rate misalignment of inconsistent fixed regimes in 1980–99, shown in figure 5, with the real exchange rate misalignment around the government change of Argentina in 2015. As the figure shows, the Argentine currency had an overvaluation of 22 percent in the months leading up to the election, which was corrected suddenly in the election month. The pattern is quite similar to the average trend for inconsistent fixed regimes in the earlier period, where overvaluation peaks at 37 percent in the government change month and is then corrected sharply in about two to three months. We also observe a post-government-change

FIGURE 6. Real Exchange Rate Misalignments: 1980–99 Period versus Argentina 2015



Notes: Graphic representation of the estimators of equation 5 for the cyclical component of the RER for the reduced sample of inconsistent de jure fixed regimes in the 1980–99 period, where the RER series is detrended using the Hodrick-Prescott filter (smoothing parameter of 129,600), and for the Argentina 2015 general elections. Month 0 is the government change month.

undervaluation that is corrected smoothly; for the 1980–99 period, the undervaluation peaks in month +5, while for Argentina 2015, it peaks in month +2.[23]

Conclusion

To explore the behavior of exchange rate policy around elections, we first classified regime announcements using the IMF de jure classification, identifying a regime as inconsistent (broken promises) when it differs from the corresponding Reinhart and Rogoff (2004) de facto classification. We then used ordered logit regressions to study the determinants of both de jure and de facto exchange rate regimes, employing several time-varying controls used in the literature to isolate the impact of dummy variables for government changes (for example, Alesina and Wagner, 2006; Juhn and Mauro, 2002;

23. Online appendix C presents the comparison of the 1980–99 period with Argentina 2015 using the Hamilton (2018) filter rather than the Hodrick-Prescott filter. The results are qualitatively the same.

Levy-Yeyati, Sturzenegger, and Reggio, 2010). We found that de jure regimes do not change in the four quarters leading up to a government change. This is important since it indicates that de jure regimes are not likely to be strongly affected by the endogeneity of regime announcements.[24] Next, we combined the two classifications to study the consistency of the announcement, rather than that of either the announcement or the de facto regime independently from each other. This is something new in the literature on exchange rate regimes.[25] We found that in the first two quarters after government changes, the probability of consistent de jure flexible regimes increases in relation to inconsistent de jure fixed regimes. This indicates that the monetary authority announces flexible regimes the first few months after a government change in an already de facto flexible environment.

We then used this classification to study the dynamics of the real exchange rate around elections conditional on consistent and inconsistent exchange rate regime announcements. We employed a dynamic distributed lag model and a difference-in-differences strategy. This revealed that the pattern found in the earlier political economy literature regarding incumbents who postponed depreciations until the inauguration of the new administration (for example, Cermeño, Grier, and Grier, 2010; Edwards, 1994; Stein and Streb, 2004; Stein, Streb, and Ghezzi, 2005) is specifically due to inconsistent fixed regimes. We found that during inconsistent fixed exchange rate announcements, the devaluation rate is not statistically different from consistent fixed announcements until the government change date, after which it increases and differs from the latter significantly. Although what Alesina and Wagner (2006) call fear of pegging (breaking commitments to pegging and floating more than announced) already shows up in our sample before the end of the incumbent's term, the adjustment of the official exchange rate takes place only *after* the change of government. In other words, part of the broken promise—the devaluation of the official exchange rate—shows up only afterward. One possible interpretation is that sustaining a peg before the government change date can be used as a signal of macroeconomic strength that could increase the probability of being reelected. Exchange rates can be stabilized in the short run by using international reserves and debt. Some incompetent incumbents may

24. In online appendix B, the marginal effects of $GovCh(-3)^q$, $GovCh(-2)^q$, $GovCh(-1)^q$, and $GovCh(0)^q$ are small and insignificant as well.

25. Alesina and Wagner (2006) provide a specific study of inconsistent fixed regimes, which they call fear of pegging. We develop a slightly different classification of the consistency of the announcement and also control for elections.

attempt to mimic competent ones by sustaining the peg announcement before elections (Stein and Streb, 1998, 2004).[26] However, in our sample the postponement of exchange rate adjustments is specifically linked to inconsistent fixed regimes. Hence an additional mechanism is at play: dual markets. Our results on inconsistent fixed regimes thus also suggest the presence of a channel of distributive politics. Specifically, maintaining an "official" appreciated exchange rate before elections hurts the concentrated export sectors to the benefit of the general population that consumes those goods, in particular the median voter. Afterward, the new administration devalues the exchange rate owing to the impossibility (or inconvenience) of sustaining it any longer. This resembles the logic behind the Bonomo and Terra (2005) model, which emphasizes the distributive consequences of appreciated exchange rates, though they do not consider the channel of dual markets. This could be an interesting topic for further research.

Finally, our paper contributes to the literature on real exchange rate appreciations and their reversions. Goldfajn and Valdés (1999) show that real exchange rate appreciations are usually reverted by nominal devaluations rather than through smooth inflation differentials. We identified the episodes of real exchange rate overvaluation as corresponding to inconsistent fixed regime announcements. This starts ten months before the government change date and peaks in the month of government change, with an overvaluation of 37 percent. The overvaluation is mostly reverted in one month through a sudden nominal devaluation. This process leads to a sharp undervaluation of the exchange rate, which is gradually corrected over the course of more than a year. We thus identified a precise timing for the macroeconomic scenario in which exchange rate overvaluation occurs: before the change of government. Additionally, a significant undervaluation takes place in its aftermath, in line with exchange rate overshooting.

26. Following the approach to political budget cycles under asymmetric information in Rogoff and Sibert (1988) and Rogoff (1990), Stein and Streb (1998, 2004) show that a low rate of devaluation can be used before elections by office-motivated incumbents to signal higher competence. In a two-sector model, the postponement of devaluations provokes an appreciated exchange rate (Stein, Streb, and Ghezzi, 2005). In these models, where nominal devaluation acts as a tax on consumption, tax smoothing is optimal from a welfare perspective, but some incumbents are tempted to exploit the trade-off between present and future devaluation for electoral reasons. In a setting with adaptive expectations, van der Ploeg (1989) derives a similar pattern, where the government appreciates the exchange rate before an election, to increase the real income of voters and boost its popularity, and depreciates it afterward. However, his prediction that all incumbents engage in this electoral manipulation is at odds with the evidence.

As to future lines of research, it may be interesting to study how the institutional setup affects the consistency of exchange rate regime announcements. The literature on central bank independence mainly focuses on outcomes like inflation and economic performance (for example, Alesina and Summers, 1993; Garriga and Rodriguez, 2020) or on exchange rate manipulation and volatility (for example, Cermeño, Grier, and Grier, 2010). Higher degrees of central bank independence might increase the likelihood of consistent exchange rate regime announcements (fixed and flexible) during the electoral window and beyond. It may also be interesting to study whether inconsistent fixed regimes lead to a lower probability of reelection and, more generally, whether multiple exchange rate markets affect electoral results.

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