Broken promises: regime announcements and exchange rates around elections

Pablo Garofalo
New Jersey City University

Jorge M. Streb[*]
Universidad del CEMA

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 afterwards.

JEL classification codes: D72, D78, E00
Key words: exchange rates, exchange-rate misalignment, exchange-rate regimes, electoral cycles

I. Introduction

Reneging on exchange-rate regime announcements occurs quite often. We track exchange-rate regime announcements with the IMF de jure exchange-rate classification that reports what countries claim to be doing.[1] The IMF de jure classification has been criticized for representing words, not deeds (Reinhart and Rogoff 2004, Levy-Yeyati and Sturzenegger 2005). Among the de facto classifications proposed, Reinhart

---

[*] Pablo Garofalo: New Jersey City University, School of Business, Department of Economics, 160 Hudson St # 205H, Harborside Plaza 2 – 2nd floor, Jersey City, New Jersey, 07311, USA; pgarofalo@njcu.edu. Jorge M. Streb: Universidad del Cema, Av. Córdoba 374, 1054 Buenos Aires, Argentina; jms@ucema.edu.ar. The editor, Veronica Rappoport, and two reviewers helped us to clarify the conceptual framework and sharpen the empirical results. We also appreciate the comments made by Juan Pereyra and participants at the Jornadas Anuales de Economía of the Banco Central del Uruguay at Montevideo, in 2017, the 45th Annual Conference of the Eastern Economic Association at New York City, NY, in 2019, the 56th Annual Meeting of the Public Choice Society at Louisville, KY, in 2019, and the 55th Annual Meeting of the Asociación Argentina de Economía Política in Buenos Aires, Argentina, in 2020.

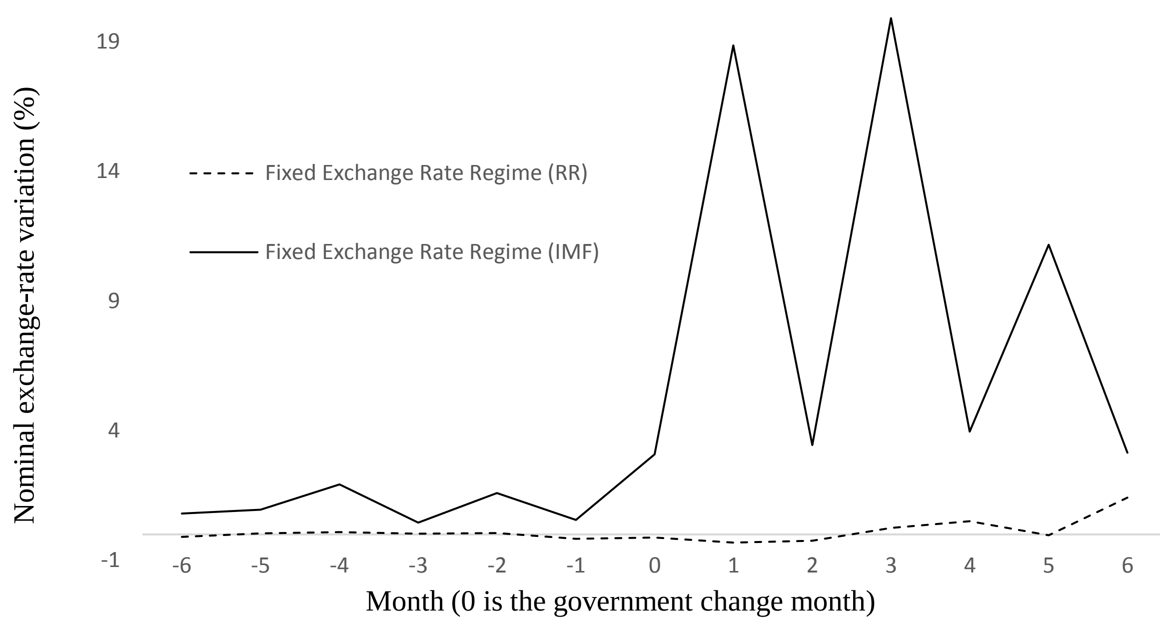
[1] Exchange-rate regime announcements are here distinguished from firmer monetary commitments like dollarization, in which the country relinquishes an independent currency, e.g., Panama since 1904, Ecuador since 2000, and El Salvador since 2001. A regime announcement continues to hold until further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

and Rogoff (2004) reclassify exchange-rate arrangements by developing an algorithm based on the observed behavior of exchange rates, while parallel exchange rates are used if multiple markets are present.[2] While Reinhart and Rogoff (2004: 1) claim that the IMF exchange-rate classification is "a little better than random", we have reasons to suspect otherwise.

Using the IMF and RR classifications (henceforth, RR refers to Reinhart and Rogoff 2004), 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 afterwards 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 afterwards.

[2] In the next section we discuss the Levy-Yeyati and Sturzenegger (2005) classification.

Figure 1. Exchange-rate devaluations around government changes



Note: Average exchange-rate variation during 19 [21] complete episodes in 21 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-Tobago, Uruguay and Venezuela) within 1980-1999 period computed with RR [IMF] fixed exchange-rate classification. Both classifications invariant throughout 12-month window. Dollarization episodes excluded.

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 US dollar. Conversely, Alesina and Wagner (2006) show that some countries often break commitments to pegging and end up floating more than what 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, i.e., 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 have 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 (e.g., Edwards, 1994; Stein and Streb, 2004; Stein, Streb, and Ghezzi, 2005; Cermeño, Grier, and Grier, 2010), we find that previous 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 that have been called "fear of pegging".[3] Observationally, de jure fixed regimes that are inconsistent with the de facto flexible behavior share identifiable underlying characteristics, namely, dual/multiple markets and high inflation before elections. These "fixed-inconsistent regimes", for short, always involve broken promises. On the contrary, "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 one might simply be following a similar monetary policy, so it is not breaking any commitment.

We first study the determinants of the exchange-rate regimes around elections using ordered logit models for both the IMF de jure and RR de facto regime classifications. As found, among others, by Klein and Marion (1997) and Gavin and Perotti (1997), there is no evidence that de jure regimes change before the government change date, 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 flexible de facto regime increases before government 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 fixed-inconsistent regimes, which are de facto flexible, decreases in relation to the probability of flexible-

[3] Inconsistent de jure fixed regimes are a slight modification of what Alesina and Wagner call "fear of pegging". We develop the rationale of this modification in section III.

consistent 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 so 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-difference strategy.[4] We find that exchange-rate behavior during consistent and inconsistent de jure fixed regimes is not statistically different until the month of the government change, but it differs significantly in the first quarter after that. Hence, although fixed-inconsistent regimes might tend to be episodes of "poor macroeconomic performance and inability to maintain monetary and fiscal stability" (Alesina and Wagner, 2006: 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 leads 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 fixed-inconsistent regimes. Such overvaluation begins ten months before the government change date and lasts until two months after the government change (about one year of overvaluation), with a peak of 37% in the government change month. Reversion starts abruptly the next month and is completed in three months. This corroborates the Goldfajn and Valdés (1999) findings 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: fixed-inconsistent regimes before government changes.

[4] In Section III we explain how we construct our four-category regime classification that identifies de jure regime consistency, and in Section IV we tackle the endogeneity problem of these categories.

Finally, we compare the 1980-1999 period, to which the findings described above apply, with the 2000-2016 period. Though the IMF changed its traditional de jure classification for a de facto one 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 non-existent during the 2000-2016 period, but we provide a case study for the only election where the regime is likely to be classified as fixed-inconsistent, the Argentine general election of 2015.

The rest of the paper is organized as follows. In Section II, we review the exchange-rate classification literature. In section III, we explain the methodology followed to identify consistent and inconsistent exchange-rate regime announcements. In Section IV, we present the econometric models and results. In Section V, we analyze the appreciation of the real exchange rate and its reversion. Section VI compares the 1980-1999 period with the most recent period by relying on the underlying characteristics of fixed-inconsistent regimes. Section VII concludes.

II. Exchange-rate regime classifications

The IMF developed a traditional exchange-rate regime classification.[5] Until 1999, it asked "country members to self-declare their arrangement as belonging to one of four categories" (Alesina and Wagner, 2006: 775): float, manage, crawl and fix. 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 US dollar" (Alesina and Wagner, 2006: 775). There are many reasons to seek other approaches to classifying exchange-rate regimes. For instance, empirical work on the cost and benefits of alternative exchange-rate arrangements can be misleading when there are significant deviations of the actual behavior from the pre-announced behavior; as pointed out by Reinhart and Rogoff (2004), Baxter and

[5] *Annual Report on Exchange Arrangements and Exchange Restrictions*.

Stockman (1989) found 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 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 does not as much. 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: 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 accordingly. If the de jure regime fails verification, they seek a de facto statistical classification based on the behavior of the exchange rate if the rate of 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.

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 that 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: 797), we are interested in how de facto behavior deviates from announced official policies, so this also points to the RR classification.

III. Consistency of de jure and de facto exchange-rate regimes

In order to identify consistent and inconsistent de jure regimes (whether the official pre-announced regime matches the actual policy or not) we follow an approach similar to Alesina and Wagner (2006). To quantify broken promises, which we call "inconsistencies", they take the difference between the coarse RR and IMF classifications. If the official pre-announced (de jure) regime is, say, manage [float], when the IMF classification equals 3 [4], while the natural classification is, say, float [manage], when the RR classification equals 4 [3], then the difference is positive [negative] and called "fear of pegging" ["fear of floating"]. In Figure 2 all the possible combinations that form either fear of pegging or fear of floating are depicted. The upper left shaded area represents fear of pegging, while the lower right, fear of floating. Each cell has three numbers $X, Y(Z)$, where X represents the RR classification, Y the IMF classification, and $Z = X - Y$. Note that $Z < 0$ represents fear of floating, $Z > 0$, fear of pegging, and $Z = 0$, consistent de jure regimes.

Figure 2. Classification of de jure exchange-rate regimes by Alesina and Wagner (2006)

RR de facto classification (actual policy)				
Float	4,1 (+3)	4,2 (+2)	4,3 (+1)	4,4 (0)
Manage	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)
Fix	1,1 (0)	1,2 (-1)	1,3 (-2)	1,4 (-3)
	Fix	Crawl	Manage	Float

IMF de jure classification (announcement)

Note: Each cell contains three numbers, $X, Y(Z)$. X represents RR classification and Y IMF classification, where 4 is float, 3 manage, 2 crawl, and 1 fix, while $Z = X - Y$. Fear of floating, with more managing than announced ($Z < 0$), in dark gray area. Fear of pegging, with more floating than announced ($Z > 0$), in light gray area. Source: Alesina and Wagner (2006).

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 (fix or crawl) and flexible (manage or float). We create the categories using a two-dimensional classification system: fix versus flexible and consistent versus inconsistent. Our approach is depicted in Figure 3. There we observe four categories of de jure regimes: (1) fixed-consistent, the striped area at the bottom left, (2) flexible-consistent, the unshaded area at the top right, (3) fixed-inconsistent (strong fear of pegging), the light gray area at the top left, and (4) flexible-inconsistent (strong fear of floating), the dark gray area at the bottom right. Note that in our case $Z = 0$ or $Z = \pm 1$ correspond to consistent de jure regimes categories, 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, i.e., $Z \leq abs(1)$ belongs to consistent de jure regimes, while $Z > abs(1)$ belongs to inconsistent ones.

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)
Manage	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)
Fix	**1,1 (0)**	**1,2 (-1)**	1,3 (-2)	1,4 (-3)
	Fix	Crawl	Manage	Float
				IMF de jure classification (announcement)

Note: Each cell contains three numbers, $X, Y(Z)$. X represents RR classification and Y IMF classification, where 4 is float, 3 manage, 2 crawl, and 1 fix. $Z = X - Y$. De jure regimes: flexible-inconsistent ($Z < -1$) in dark gray area, fixed-inconsistent ($Z > 1$) in light gray area, fixed-consistent in striped area, and flexible-consistent in unshaded area.

IV. 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 to what extent they are sensitive to the electoral window. This is an important question to answer since regime types will be 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 1980-1999 period because the IMF abandoned its de jure classification after that.

We collect monthly data on exchange rates and inflation from twenty-one Latin American countries from the *IMF International Financial Statistics* over the 1980-1999 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-Tobago, Uruguay, and Venezuela.[6] We construct the series of multilateral real exchange rate, which is a trade-weighted average of bilateral real exchange rates. We follow the Goldfajn and Valdés (1999) approach of using only trading partners above 4 percent of overall trade. Also, as in Goldfajn and Valdés (1999), we fixed trade weights using trade flows of an intermediate year (1995 in our case) from the UN *International Trade Statistics Yearbook*.[7] The RR "natural" monthly exchange-rate regime classification comes from Ethan Ilzetzki's website, an up-to-date version of the original data from Carmen M. Reinhart's webiste.[8] The "traditional" IMF annual exchange-rate regime classification comes from the IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER)*. 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 Appendix A).

A. Duration of exchange-rate regimes

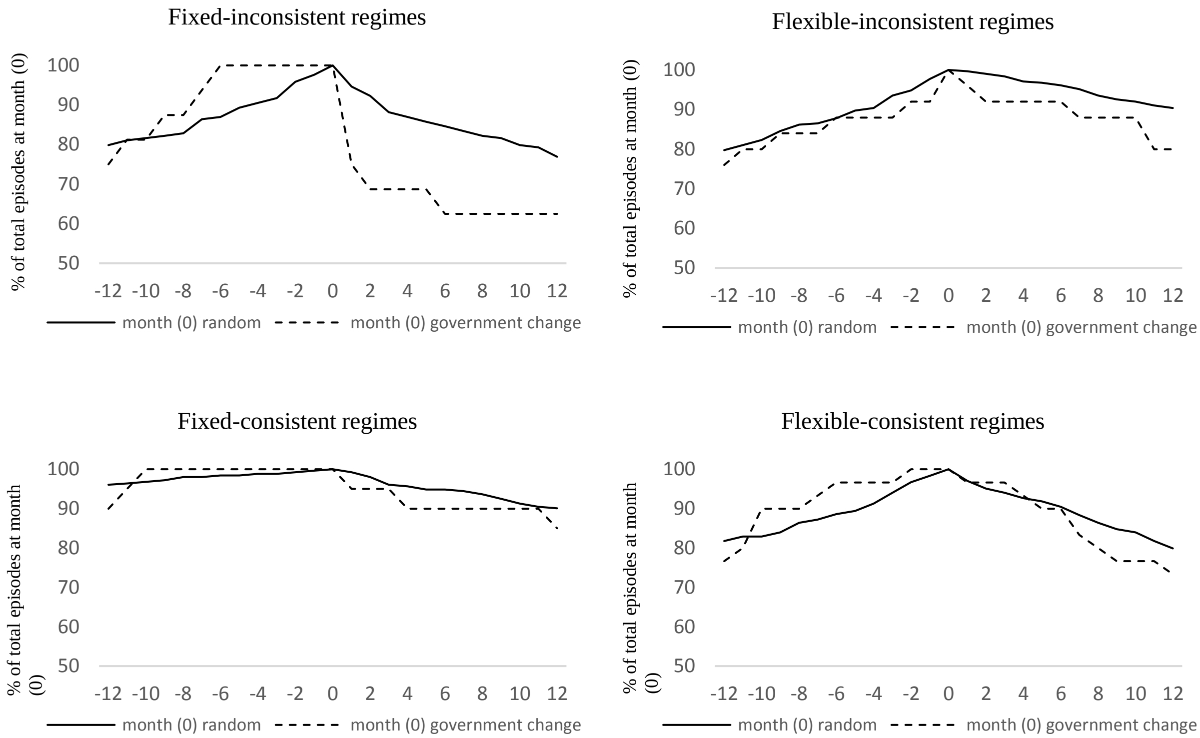
We compare regime duration inside and outside the electoral window, because exchange-rate estimations controlling for regime at election time may be biased if regime duration is sensitive to the electoral window. We proceed to report summary statistics for our four-category regime classification: (i) fixed-inconsistent, de jure fixed in the IMF classification and de facto flexible in the RR classification, (ii) fixed-consistent,

[6] Chile, El Salvador, Guyana, Jamaica, Paraguay, and Trinidad and Tobago are dropped from the sample when the full set of covariates is used due to missing observations in control variables for these countries. We thus work with two samples: a reduced sample where these countries are excluded, and an extended sample where they are included. 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.
[7] Identical qualitative results were found using only the bilateral real exchange rate with the U.S. This may be because the U.S is the main trade 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 upon request.
[8] https://www.ilzetzki.com/irr-data for Ethan Ilzetzki, and http://www.carmenreinhart.com for Carmen M. Reinhart.

both de jure and de facto fixed, (iii) flexible-inconsistent, de jure flexible and de facto fixed, and (iv) flexible-consistent, both de jure and de facto flexible.

In order to compare regime duration inside and outside the electoral window, we generate with a uniform distribution a random number between 1 and 240 in order to select a month in the 1980-1999 period for a given country. Then we observe which of the four regime classifications that month belongs to and construct a ± 12 month windows to see the duration of that regime. For example, some episodes change at, say, month -5 (that is, the classification at month 0 started on month -5), while some episodes change at, say, $+5$ (that is, the classification at month 0 ended on month +5). We repeat this randomization 50 times for each country and count afterwards for each month 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 both windows: 15.4% [17.6%] are fixed-inconsistent regimes for the random [government change] month 0, 28.3% [27.5%] are flexible-inconsistent regimes, 22.9% [22.0%] are fixed-consistent regimes, and 33.5% [33.0%] are flexible-consistent regimes. This similarity suggests that the distribution of regimes is not sensitive to the electoral window. Results are displayed in Figure 4.

Figure 4. Duration of regime classifications around random and government-change months



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

When looking at the duration of the regimes in Figure 4, about 95% of the episodes of fixed-consistent regimes (bottom-left graph) were already in that category 11 months before month 0 for both windows, and about 90% of the episodes continue to belong to this category 11 months later. Distributions are also similar for flexible-inconsistent (upper-right graph), and flexible-consistent regimes (lower-right graph). The exception is given by fixed-inconsistent regimes (upper-left graph). There appears to be a longer duration of cases in this category before month 0 for the government change window, for instance, at month -6 100% of the episodes already belong to this category in contrast to only 87% 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 anticipates a finding below: some incumbents switch categories from inconsistent fixed to consistent flexible.

B. 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 1980-1999 to compare with the findings in previous studies. Since both classifications represent clear ranks with a meaningful order (i.e., fix, crawl, manage and float), ordered logit models are the appropriate option. Then we study the determinants of our novel regime classification, i.e., 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:

$$\mathrm{P}(regime\ Y_{it} = y |\ \mathbf{X_{it}}, \mathbf{govch}(\mathbf{q}^{-})_{\mathbf{it}}^{\boldsymbol{q}}, \mathbf{govch}(\mathbf{q}^{+})_{\mathbf{it}}^{\boldsymbol{q}}), \tag{1}$$

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 value 1 (2) [3] {4} if the regime is fix (crawl) [manage] {float}. For the multinomial logit model, where the categories are not ruled by any apparent rank order, the dependent variable *y* takes value 1 (2) [3] {4} if the de jure regime is fixed-consistent (flexible-inconsistent) [fixed-inconsistent] {flexible-consistent}. $\mathbf{govch}(\mathbf{q}^{-})_{\mathbf{it}}^{\boldsymbol{q}} = \left(govch(-3)_{it}^{q}\ \ govch(-2)_{it}^{q}\ \ govch(-1)_{it}^{q}\ \ govch(0)_{it}^{q}\right)$ is a matrix of four dummy variables. $govch(0)^{q}$ takes value of 1 in the months 0 to 2 before the government change month (month 0 is when the government changes), $govch(-1)^{q}$ takes value of 1 in the months 3 to 5 before the government change month, while $govch(-2)^{q}$ and $govch(-3)^{q}$ are defined similarly. Note that although the data is monthly, we are defining dummy variables per quarters (the superscript "q" stands for quarter). Analogously, $\mathbf{govch}(\mathbf{q}^{+})_{\mathbf{it}}^{\boldsymbol{q}} = \left(govch(+1)_{it}^{q} \;\; govch(+2)_{it}^{q} \;\; govch(+3)_{it}^{q} \;\; govch(+4)_{it}^{q}\right)$ is constructed for the 12 months following the month of the change of government using four quarterly dummy variables.

$\mathbf{X}$ is a matrix composed by the following time-varying controls: (i) *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*. This variable is used by Levy-Yeyati, Sturzenegger, and Reggio (2010) as a proxy variable for capital mobility. Based on the "impossible trinity" argument, policymakers should give up on either monetary policy or exchange-rate policy in environments with high capital mobility. Thus, intermediate regimes are less viable. Alternatively, given the "currency mismatch" argument, we should expect more commitments to pegging; (ii) *Foreign.Liab.pc*, foreign liabilities per capita, from the IMF *International Financial Statistics*. Countries with important foreign liabilities may be more prone to fix their currency since sharp nominal depreciation of the currency impact on the solvency of the non-tradable sector's balance sheets. Alesina and Wagner (2006), and Levy-Yeyati, Sturzenegger, and Reggio (2010) use foreign liabilities over monetary aggregates instead. The problem with this variable is that for Latin American countries, money demand is extremely unstable during the 80's and beginning of the 90's due to high inflation regimes. In crisis episodes during 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 opposed to the currency mismatch hypothesis;[9] (iii) *Size*, real GDP in dollars, from the IMF *International Financial Statistics*. As noted in Levy-Yeyati, Sturzenegger, and Reggio (2010), smallness favors a more stable exchange rate through the higher propensity of small economies to trade internationally, and by limiting the scope for the use of a national unit of account; (iv) *ToT,* terms of trade. When 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 be insured against sudden stops

[9] We indeed find a significant positive coefficient when foreign liabilities over money are used, so the probability of a flexible regime increases when foreign liabilities to money increase. There is instead a negative coefficient with our transformation of foreign liabilities normalized with 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.

(Jeanne, 2007; Jeanne and Rancière, 2011); (v) *U.S.interest*, the U.S. interest rate in real terms, from the IMF *International Financial Statistics*. Calvo et al. (1993) and Fernandez-Arias and Montiel (1996) find that the U.S interest rate is a determinant for capital inflows in Latin America.[10] When 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; (vi) *Openness*, exports plus imports over GDP, from the IMF *International Financial Statistics*. The decision of pegging 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; finally, (vii) *Default*, a dummy variable that takes value 1 if the country has defaulted the external debt and 0 otherwise, from Carmen M. Reinhart's website, is used to control for the fact that these economies cannot sustain their currency given a high macroeconomic instability, so they let their currency float or, more precisely, freely fall.

Among the seven controls describe 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 given that it is a dummy variable. The rest, *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 1-month lagged values of the variables available at monthly frequency. For the variables available at annual frequency that were interpolated using log differences, we adopt 12-months lagged value instead. Except for *Default* and dummy variables for government change, the rest of the variables are expressed in natural logs.

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

Table 1. Determinants of exchange-rate regimes. Ordered and multinomial logit models

	(1)	(2)	(3)		
	IMF (de jure)	RR (de facto)	Consistent de jure fixed	Inconsistent de jure flexible	Consistent de jure flexible
			relative to inconsistent de jure fixed		
$ln\ Portfolio_{t-12}$	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]
Observations	2,557	2,592		2,662	

Notes: Estimation of Equation (1) with ordered logit models in Columns 1 and 2. Dependent variable equals 1 (2) [3] {4} if regime is fix (crawl) [manage] {float}. Estimation of Equation (1) with multinomial logit in Column 3. Dependent variable equals 1 (2) [3] {4} if de jure regime is fixed-consistent (flexible-inconsistent) [fixed-inconsistent] {flexible-consistent}; results relative to fixed-inconsistent category. Reduced sample includes Argentina, Bolivia, Brazil, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Mexico, Nicaragua, Peru, Uruguay, and Venezuela over 1980-1999 period. Non-democratic episodes excluded based on Polity IV Project. Dollarization episodes also excluded. Robust standard errors in brackets. (*) [**] {***} stands for significance at (10%) [5%] {1%}.

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, Columns 1 and 2. In Column 3, we present 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 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; in Appendix B, we present the full set of marginal effects for each of the categories using mean values of covariates. For the latter, each coefficient is understood as the increase in the probability of category $j = 1,2,4$ in relation to the category 3 (base) for a marginal increase of the independent variable, if its coefficient is positive; again, we provide the full set of marginal effects in Appendix B at mean values of covariates.

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.[11] Our innovation in relation to the literature is the study of our novel regime classification that is divided into consistent and inconsistent de jure regimes (results are displayed in Column 3 of Table 1). We then focus on the issue of endogeneity of regimes around the government change date.

For the RR de facto classification in Column 2, the probability of observing fixed regimes tends to increases as the de facto capital account openness increases (i.e., $Portfolio = -0.045^{**}$ in Column 2). This is consistent with the "currency mismatch" hypothesis of Levy-Yeyati, Sturzenegger, and Reggio (2010). By contrast, we observe that de jure flexible regimes tend to increase as the de facto capital account openness increases (i.e., $Portfolio = \ 0.046^{**}$ in Column 1). Altogether, these may be suggestive of an increase in flexible-inconsistent regimes. In Column 3, we find that the facto fixed regimes, in both the fixed-consistent and flexible-inconsistent versions, are the most likely outcomes compared to fixed-inconsistent regimes (i.e., 0.166^{***} and 0.168^{***}, respectively). Moreover, fixed-inconsistent regimes are less likely since all three coefficients are positive (i.e., all regimes are more likely in relation to the base category). While foreign liabilities per capita ($Foreign.Liab.pc$) is close to zero and insignificant for the IMF de jure classification in Column 1, it is significantly negative in de facto classification in Column 2 (i.e., $Foreign.Liab.pc = -0.049^{***}$), which is consistent with the currency mismatch hypothesis. We corroborate this finding in Column 3 by observing that both types of fixed regimes, consistent, in Column 3-i, and inconsistent, in Column 3-ii, are more likely in relation to inconsistent fixed regimes (i.e., 0.265^{***}

[11] The connection between regression coefficients and changes in probabilities is detailed in Appendix B.

and 0.201^{***}, respectively). Consistent flexible regimes are also more likely relative to inconsistent fixed regimes (i.e., 0.297^{***} in Column 3-iii). This corroborates that inconsistent fixed regimes do not go hand-in-hand with liability dollarization, mainly because those are episodes of highly 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), we observe in Column 3 that the consistent flexible category is the most likely outcome given that its estimator is the greatest among all three. This indicates that indeed flexible regimes are more likely in bigger countries, while inconsistent fixed is the less likely outcome given that 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 column 3, both de facto fixed regimes, the fixed-consistent of Column 3-i and the flexible-inconsistent of Column 3-ii, are the most likely outcomes 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, given that all three estimators are positive, it indicates that the less likely outcome when terms of trade increase is a fixed-inconsistent regime. This should occur 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 of Column 2. Regarding the de jure regime, in Column 1, we observe a strong increase in the likelihood of a peg 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/short run, that is to say, fixed-inconsistent regimes. This seems to occur given that an increase in the U.S. interest rate produces capital outflows from the Latin American region, as found in Calvo et al. (1993) and Fernandez-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 findings in column 3 corroborate this view, since all three estimators are significantly negative, which indicates that the likelihood of fixed-inconsistent regimes increases when the US interest rate increases. *Openness* possesses the predicted negative sign in the de facto classification of Column 2 (i.e., more open economies are in favor of a more stable exchange rate). However, the de jure regime on column 1 is significantly positive. Complementing this with the analysis of column 3, we observe that fixed-inconsistent 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. We observe that the market-based exchange rate tends to float when economies default on their debt (i.e., $Default = 1.573^{***}$ in Column 2), while the de jure keeps up with the market behavior (i.e., $Default = 0.588^{***}$ in Column 1). Column 3 reveals that defaults definitely decreases the probability of de facto fixed regimes, in both the consistent fixed and inconsistent flexible versions, in relation to inconsistent fixed regimes (i.e., in Column 3-i $Default = -0.660^{***}$, in Column 3-ii $Default = -0.493^{***}$). It also reveals that the most likely outcome is consistent flexible regimes (column 3-iii is the only positive coefficient), which is congruent with the findings in columns 1 and 2 (that is, de jure and the facto regime becomes more flexible, so does the consistency of flexible regimes).

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 the potential endogeneity of the de jure regimes are not likely to be strongly affected by the endogeneity of regime announcement.[12] After government changes, 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.[13] As to the de facto classification, the exchange rate tends to be more flexible both before and after the government change date, since the estimators $govch(-3)^q$, $govch(-2)^q$, $govch(-1)^q$ and $govch(0)^q$, $govch(+1)^q$, $govch(+2)^q$, $govch(+3)^q$, and $govch(+4)^q$ are significantly positive.

[12] In 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.

[13] These results resemble the findings in Klein and Marion (1997), and Gavin and Perotti (1997). The finding in Blomberg, Frieden, and Stein (2005) that the duration of pegs increases before elections and decreases afterwards is congruent with this pattern.

Results in Column 3, where we use our novel regime classification of Figure 3, indicates that no regime is more likely to occur before government changes in relation to inconsistent de jure fixed regimes. This is congruent with the analysis in Figure 4 above, in which regime duration distribution are quite similar for all the regimes in the 12 months before either government changes or a randomly generated month 0. After government changes, we observe in Column 3 that for the first two quarters the probability of flexible-consistent regimes increases in relation to fixed-inconsistent regimes ($govch(+1)^q = 0.985^{***}$ and $govch(+2)^q = 0.838^{**}$). This indicates that the monetary authority announces flexible regimes the first few months after a government change in an already flexible de facto environment. This is congruent with the sharp, sudden drop of inconsistent de jure fixed regimes right after government changes in the upper left graph of Figure 4, which contrasts with the behavior after a randomly selected month 0.

C. The dynamics of the real exchange rate

After finding that there is 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 form:

$$\Delta \ln(RER_{it}) = \sum_{k=1}^{3} a_k \, \Delta \ln\left(RER_{i,t-k}\right) + \Delta\mathbf{W}_{\mathbf{it}}'\boldsymbol{\beta} + \mathbf{govch}_{\mathbf{it}}' \, \boldsymbol{\delta} + \mathbf{govchFI}_{\mathbf{it}}' \, \boldsymbol{\delta}_{\mathbf{FI}} + \mathbf{govchFEI}_{\mathbf{it}}' \, \boldsymbol{\delta}_{\mathbf{FEI}} + \mathbf{govchFEC}_{\mathbf{it}}' \, \boldsymbol{\delta}_{\mathbf{FEC}} + e_{it}, \quad (2)$$

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 capture persistency.[14] $\mathbf{govch}_{\mathbf{it}}' = (govch(-3)^q \; govch(-2)^q \; govch(-1)^q \; govch(0)^q \; govch(+1)^q \; govch(+2)^q \; govch(+3)^q \; govch(+4)^q)'$ represents quarterly dummy variables, where $govch(\pm l)^q$ takes value one if the government change is $\pm l$ quarters away. ($\mathbf{govchFI}$)[$\mathbf{govchFEI}$]{$\mathbf{govchFEC}$} is the interaction between $\mathbf{govch}$ with an (inconsistent and fixed) [inconsistent and flexible] {consistent and flexible} de jure regime; the omitted category are consistent de jure fixed regimes. The regime classification used for the entire electoral window is invariant

[14] Results are totally invariant to the inclusion of one lag instead. Results with one lag are available upon request.

and equal to the classification at the month before the elections, which is typically two or three months before government changes.[15] $\mathbf{W}$ is a matrix of time-varying controls that attempt to control for both determinants of 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 determinants of regime announcement, namely, *Porfolio*, *Foreign.Liab.pc*, *Size*, *ToT*, *U.S.interest*, *Openness*, *Default*. Given that an expansion in the size of government will induce an appreciation of the real exchange rate when government demand is biased towards non-tradable goods, as stressed in Goldfajn and Valdés (1999), we add government expenditure as a ratio of GDP, *Govsize*. Given that we could not corroborate that our regressors produce a co-integrating vector,[16] we decided to estimate the model in first differences, as done in Cermeño, Grier, and Grier (2010). However, our results do not change significantly once we study Equation (2) in levels.[17] Finally, given the possibility of reverse causality, we decided to use 1-month lagged values of the variables in $\mathbf{W}$. For the variables available at annual frequency that were interpolated using log differences, we adopted 12-months lagged value instead. Except for *Default* and dummy variables for government change, the rest of the variables are in natural logs. 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 ± 4 quarters. 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 due to covariate limitations.

In Column 1 of Table 2, the real exchange rate decreases (i.e., appreciates) moderately during the last quarter up to the government change for inconsistent de jure fixed regimes, but the result is not significant ($govchFI(0)^q = -4.832$); after the government change, we observe a 17% depreciation of the real exchange rate during the first quarter ($govchFI(+1)^q = 17.312^*$). Linear Combination 1, which shows

[15] Results are virtually unchanged when we use the value six months before elections instead. Results under the latter are available upon request.
[16] We ran Engle-Granger tests for each country and in almost all the countries the hypothesis of co-integration was rejected. Only Argentina, Bolivia, Guatemala, Honduras and Uruguay showed evidence of co-integration at 5% significance or higher. Test results are available upon request.
[17] Results are available upon request.

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 two, three and four quarters around the government change date, do capture statistically significant depreciation differentials of 16, 11 and 9%; the significant depreciation during the second quarter ($govchFI(+2)^q = 9.469^*$) undoubtedly contributes to this. Column 2, without including any time-varying covariate, are almost the same as those in Column 1. This indicates that the empirical design produces a plausible exogenous variation of regime adoptions, as shown in Figure 4.[18] When we include the extended sample in Column 3 that we lost due to the lack for availability of covariates, the effects drop substantially, but they are all statistically significant.[19]

[18] Appendix D1 shows that the results are also robust to allowing for conditional heteroskedasticity.

[19] 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 upon request.

Table 2. Exchange rate variation around government changes

Dependent variable $\Delta\ln RER$	(1)		(2)		(3)	
Type of sample	Reduced		Reduced		Extended	
Covariates included?	Yes		No		No	
$govch(-3)^q$	-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]
Observations	2,236		2,236		4,008	
R-squared	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]

Notes: Estimation of Equation (2). Reduced sample includes Argentina, Bolivia, Brazil, Colombia, Costa Rica, Dominican Republic, Ecuador, Guatemala, Honduras, Mexico, Nicaragua, Panama, Peru, Uruguay, and Venezuela over 1980-1999 period. Extended sample also includes Barbados, Chile, El Salvador, Guyana, Jamaica, and Paraguay. *FI*, *FEI*, and *FEC* stand for fixed-inconsistent, flexible-inconsistent, and flexible-consistent regimes; flexible-consistent is omitted category. Non-democratic episodes excluded based on Polity IV Project. Dollarization episodes also excluded. Controls used, but not reported, listed under Equation (2). OLS used for estimations. Linear Combination k = 1,2,3,4: $govchFI(+1)^q - govchFI(0)^q$, $\frac{1}{2}(govchFI(+2)^q + govchFI(+1)^q - govchFI(0)^q - govchFI(-1)^q)$, $\frac{1}{3}(govchFI(+3)^q + govchFI(+2)^q + govchFI(+1)^q - govchFI(0)^q - govchFI(-1)^q - govchFI(-2)^q)$, $\frac{1}{4}(govchFI(+4)^q + govchFI(+3)^q + govchFI(+2)^q + govchFI(+1)^q - govchFI(0)^q - govchFI(-1)^q - govchFI(-2)^q - govchFI(-3)^q)$. Robust standard errors reported in brackets at right of each estimator. (*) [**] {***} stands for significance at (10%) [5%] {1%}.

V. Real exchange-rate misalignments around government changes

In the previous section we studied the short and medium-term dynamics of the real exchange rate and found that there is a slight and insignificant appreciation quarter to quarter during the year leading to the government change under inconsistent de jure fixed regimes, and a strong and significant depreciation after the change of government. In this section we explicitly study the real exchange-rate misalignment consequences of pegging the exchange rate when it is not consistent given the market exchange-rate behavior 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.[20,21] Then the series are decomposed in two components:

$$\ln(RER_{it}) = \ln(RER_{it})_{cycle} + \ln(RER_{it})_{trend}. \tag{4}$$

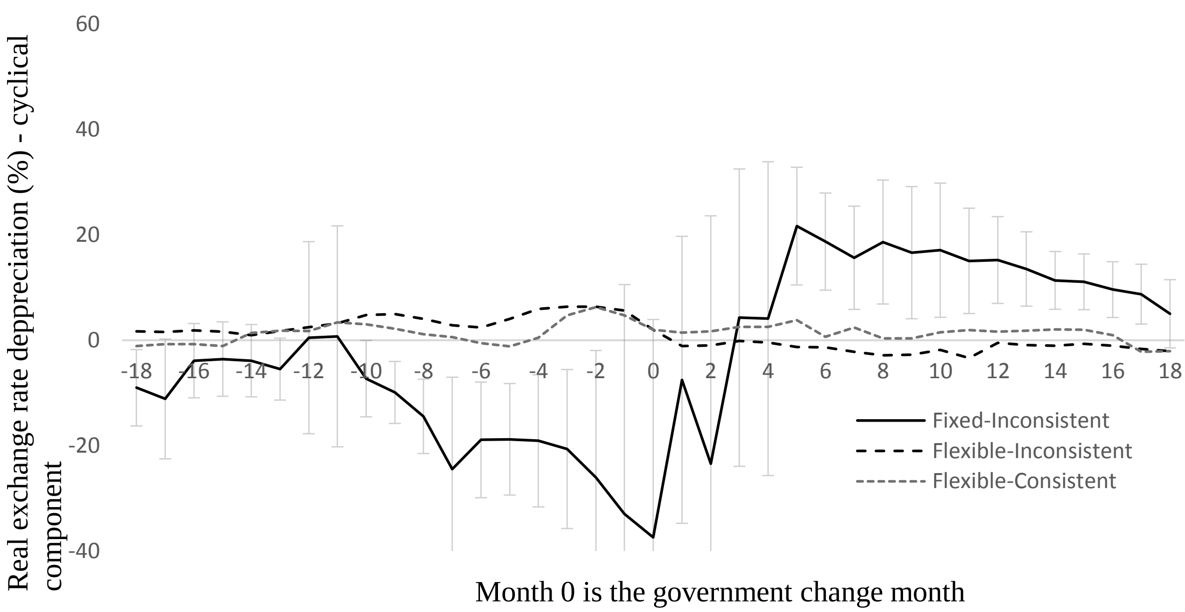
We identify the trend component as the long run RER equilibrium, and the cycle as departures from that equilibrium. When the cyclical component is negative, the RER is overvalued; when it is positive, it is undervalued. Goldfajn and Valdés (1999) identified four appreciation phases of the real exchange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

[20] To filter the RER series, we use a smoothing parameter of 129,600, which is the value Ravn and Uhlig (2002) suggest to use with monthly data.
[21] In Appendix C we use the Hamilton (2018) filtering technique, since the Hodrick-Prescott filter usually introduces a spurious dynamic relation into the series. Results are qualitatively the same.

$$\ln(RER_{it})_{cycle} = \alpha + \sum_{i=0}^{17} govch(-i)_{it}{}^{m} \cdot \beta_{(-i)} + \sum_{i=1}^{18} govch(+i)_{it}{}^{m} \cdot \beta_{(+i)} + \sum_{i=0}^{17} govchFI(-i)_{it}{}^{m} \cdot \beta_{FI(-i)} + \sum_{i=1}^{18} govchFI(+i)_{it}{}^{m} \cdot \beta_{FI(+i)} + \sum_{i=0}^{17} govchFEI(-i)_{it}{}^{m} \cdot \beta_{FEI(-i)} + \sum_{i=1}^{18} govchFEI(+i)_{it}{}^{m} \cdot \beta_{FEI(+i)} + \sum_{i=0}^{17} govchFEC(-i)_{it}{}^{m} \cdot \beta_{FEC(-i)} + \sum_{i=1}^{18} govchFEC(+i)_{it}{}^{m} \cdot \beta_{FEC(+i)} + \epsilon_{it}. \quad (5)$$

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 is reverted. Afterwards, we collapse these monthly dummy variables into semesters to observe medium run misalignments. For the sake of presentation, results of the monthly dummy variables are shown in Figure 5, while the semestral dummy variables are shown in Table 3.

Figure 5. Real exchange rate misalignments around government changes



Notes: Graphic representation of estimators of Equation (5) for cyclical component of RER, detrending RER series with Hodrick-Prescott filter (smoothing parameter of 129,600). Results relative to fixed-consistent episodes. Reduced sample used. Vertical bars represent 95% confidence intervals of estimators based on robust standard errors for fixed-inconsistent estimators around government change date.

In Figure 5, we observe that a significant overvaluation occurs only for the inconsistent de jure fixed regimes announcement. The 5% history threshold is hit at month 10 before the government change, and a peak of 37% is reached at the government change month, i.e., the history/peak stage last ten months, while the peak/end period lasts only three, being completed mostly in the first month. We observe that after the government change date, there is a process of undervaluation, which becomes significant at month 5 (undervaluation of 22%), but the process reverts smoothly in 14 months, when the real exchange rate reaches its equilibrium (i.e., back to below 5% of undervaluation). Here, there is something important to stress: when the exchange rate is overvalued, a quick one-month correction is observed, which indicates that this is done through a strong nominal devaluation, as highlighted in Goldfajn and Valdés (1999). However, when the exchange rate is undervalued, as we observe in month +5, the correction takes place smoothly through either a gradual correction of the nominal exchange rate that corrects the initial overshooting that brings about an undervaluation, or an organized correction of inflation differentials. This difference between real exchange-rate 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 during both 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 in which there is 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 hits a six-month average of 25% during the last semester before government change, which is reverted during the first semester after government change, followed by 16% undervaluation during the second semester, that then declines to 9% the third semester.[22] [23]

[22] We also produce both the month-to-month figure and the semester-to-semester table of the real exchange rate misalignments by employing the Hamilton (2018) filter. Results, shown in Appendix C, do not change significantly.
[23] Appendix D2 allowing for conditional heteroskedasticity shows that the results are robust.

Table 3. Real exchange rate misalignments using six-month dummy variables

Dependent variable $ln(RER_{it})_{cycle}$	(1)	
$govch(-2)^s$	-1.5133	[1.008]
$govch(-1)^s$	-2.6879***	[0.962]
$govch(0)^s$	-3.0408***	[0.934]
$govch(+1)^s$	0.9747	[1.070]
$govch(+2)^s$	1.0024	[1.035]
$govch(+3)^s$	-0.4071	[0.926]
$govchFI(-2)^s$	-4.5510**	[2.191]
$govchFI(-1)^s$	-12.3445***	[3.038]
$govchFI(0)^s$	-25.4982***	[5.556]
$govchFI(+1)^s$	2.9813	[6.359]
$govchFI(+2)^s$	16.2924***	[2.226]
$govchFI(+3)^s$	9.7655***	[1.305]
$govchFEI(-2)^s$	1.7333	[1.090]
$govchFEI(-1)^s$	3.7338***	[1.004]
$govchFEI(0)^s$	5.0865***	[1.012]
$govchFEI(+1)^s$	-0.9429	[1.046]
$govchFEI(+2)^s$	-2.2551**	[1.128]
$govchFEI(+3)^s$	-1.2326	[1.021]
$govchFEC(-2)^s$	0.4289	[1.283]
$govchFEC(-1)^s$	1.6642	[1.582]
$govchFEC(0)^s$	2.8312	[2.096]
$govchFEC(+1)^s$	2.1107	[1.710]
$govchFEC(+2)^s$	1.3806	[1.474]
$govchFEC(+3)^s$	0.4493	[1.295]
Observations	2,127	
R-squared	0.119	

Notes: Estimation of Equation (5) for cyclical component of RER using six-month dummy variables, detrending RER series with Hodrick-Prescott filter (smoothing parameter of 129,600). Results relative to fixed-consistent episodes. Reduced sample used. Non-democratic episodes excluded based on Polity IV Project. Dollarization episodes also excluded. OLS used for estimation. Robust standard errors reported in brackets at right of each estimator. (*) [**] {***} stands for significance at (10%) [5%] {1%}.

VI. Comparison of 1980-1999 and 2000-2016 periods

In this section, we first compare the characteristics of exchange-rate regimes for the 1980-1999 and 2000-2016 periods. Then we do a case study of the 2015 Argentine general elections, which share the underlying characteristic of inconsistent de jure fixed regimes, and compare it to the econometric findings of the earlier period.

There are 81 changes of government in the 1980-1999 period and 84 in the 2000-2016 period. Table 4 shows that de facto flexible regimes become much less common in the recent period, falling from 57% to 30% 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 the period 1980-1999, two common shared characteristics of fixed-inconsistent regimes, where the preannounced fixed regime does not coincide with the actual policy, are dual/multiple markets and high inflation (more than 10% a year). Dual markets and high inflation characterize 81% of the fixed-inconsistent regime cases in the earlier period, as well as 73% of the flexible-consistent regime cases, while within the total de facto flexible cases, the proportion that involves dual markets and high inflation is 76%. This proportion drops to only 4% in the later period. That 4% represents a single case, something to which we return shortly. Hence, although we cannot observe the IMF de jure classification after 1999, we can still observe that dual/multiple markets and high inflation become exceptional after 1999. This provides indirect evidence that fixed-inconsistent regimes are no longer likely.

Table 4. Regime classification, 1980-2016. Number of episodes at government change date

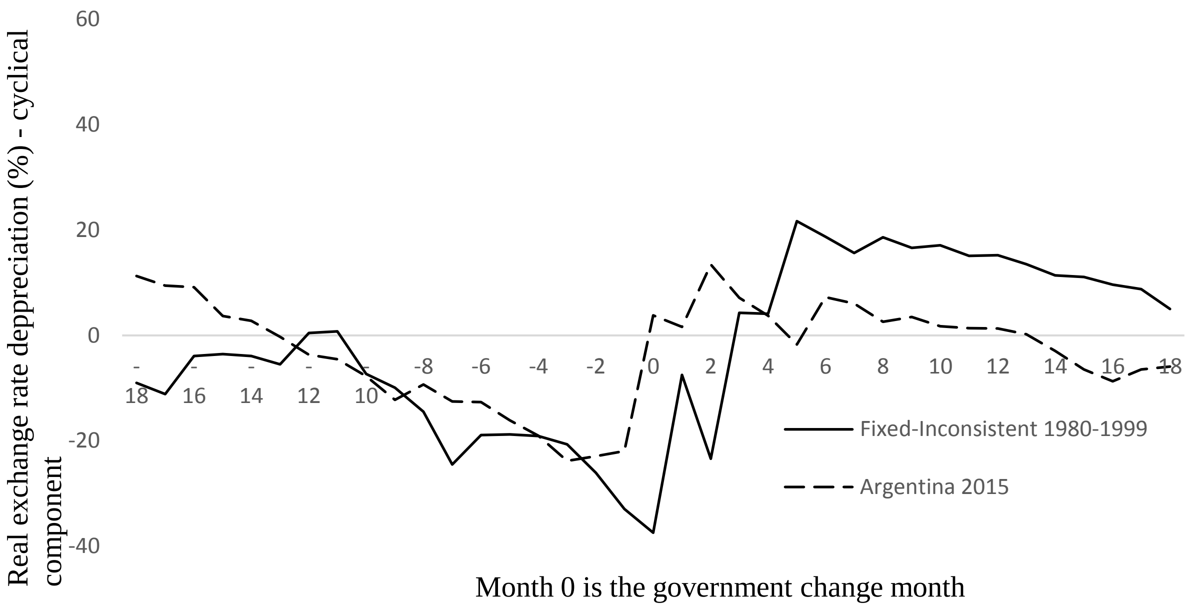
Regime	Total cases	Cases with dual markets	Cases with annual inflation > 10%	Cases with dual markets & annual inflation > 10%
		1980-1999 period		
Inconsistent de jure fixed	16	16	13	13
Consistent de jure flexible	30	23	27	22
Total de facto flexible	46	39	40	35
Inconsistent de jure flexible	24	8	14	6
Consistent de jure fixed	7	2	3	2
Total de facto fixed	31	10	17	8
		2000-2016 period		
Total de facto flexible	25	1	20	1
Total de facto fixed	49	3	33	2

Note: Dollarization episodes excluded.

The only case we can identify that is likely to have been classified as belonging to inconsistent de jure fixed regimes, because it shows flexible de facto 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% 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 fixed-inconsistent regimes for the period 1980-1999, shown in Figure 5, with the real exchange rate misalignment around the government change of Argentina in 2015: the misalignment of Argentina, an overvaluation of 22% that is corrected suddenly in a month, is pretty similar to the one found for the earlier period, an overvaluation that hits 37% and is corrected in about two to three months. We also observe an after-government-change undervaluation that is corrected smoothly; for the 1980-1999 period the undervaluation hits the peak at month +5, while for Argentina 2015 it is hit at month +2. [24]

[24] In Appendix C we construct the same figure for the comparison of the 1980-1999 period with Argentina 2015 using the Hamilton (2018) filter rather than the Hodrick-Prescott filter. Results are qualitatively the same.

Figure 6. Real exchange rate misalignments: 1980-1999 period versus Argentina 2015



Notes: Graphic representation of estimators of Equation (5) for cyclical component of RER for period 1980-1999, detrending the RER series with Hodrick-Prescott filter (smoothing parameter of 129,600). Results shown for inconsistent de jure fixed regime (solid line) and Argentina 2015 general elections (stripped line).

VII. Conclusions

We are interested in understanding the behavior of exchange-rate policy around elections. Our first step is to classify regime announcements using the IMF de jure classification, identifying a regime as inconsistent (broken promises) when it differs from the Reinhart and Rogoff (2004) de facto one.

We then study the determinants of both de jure and de facto exchange-rate regimes with ordered logit regressions, employing several time-varying controls used in the literature (among others, Juhn and Mauro, 2002; Alesina and Wagner, 2006; Levy-Yeyati, Sturzenegger, and Reggio, 2010) to isolate the impact of dummy variables for government changes. We identified that the de jure regime does not change in the four quarters leading up to a government change. This is important since it indicates that the potential endogeneity of the de jure regimes are not likely to be strongly affected by the endogeneity of regime announcement. Afterwards, we combine both classifications to study the consistency of the announcement, rather than of either the announcement or the de facto regime independently from each other. This is something new in the literature of exchange-rate regimes.[25] We find that 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 flexible de facto environment.

We then use this classification to study the dynamics of the real exchange rate around elections conditional on consistent and inconsistent exchange-rate regime announcements. We employ a dynamic distributed lag model and a difference-in-difference strategy. This allows us to pinpoint that the pattern found in the earlier political economy literature of incumbents that postponed depreciations until the inauguration of the new administration (e.g., Edwards, 1994; Stein and Streb, 2004; Stein, Streb, and Ghezzi, 2005; Cermeño, Grier, and Grier, 2010) is specifically due to fix-inconsistent regimes. We find that during fix-inconsistent exchange-rate announcements, the devaluation rate is not statistically different from fix-consistent announcements until the government change date, but it increases and differs from the latter significantly afterwards. 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 incumbent's term ends, the adjustment of the official exchange rate only takes place *after* the change of government. In other words, part of the broken promises –the devaluation of the official exchange rate– only show up afterwards. 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 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

[25] Alesina and Wagner (2006) provide a specific study of fixed inconsistent regimes, what they call "fear of pegging". We produce a slightly different classification of the consistency of the announcement in Section III, and also control for elections.

[26] Following the approach to political budget cycle 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

linked to fixed-inconsistent regimes. Hence, an additional mechanism is at play: dual markets. Thus, our results during fixed-inconsistent regimes also suggest the presence of a channel of distributive politics: an "official" appreciated exchange rate before elections hurts the concentrated export sectors to the benefit of the general population that consumes those goods, and hence the median voter. Afterwards, the new administration devalues given 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 of 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 identify that the episodes where the real exchange rate suffers an overvaluation correspond to regime announcements that are fixed inconsistent. This starts 10 months before the government change date, reaching its peak the month of government change with an overvaluation of 37%. The overvaluation is mostly reverted in a 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 identify a precise timing for the macroeconomic scenario where exchange-rate overvaluation occurs: before the change of government. Additionally, a significant undervaluation takes place in its aftermath, in line with exchange-rate overshooting.

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 (e.g., Alesina and Summers, 1993; Garriga and Rodriguez, 2020), or on exchange-rate manipulation and volatility (e.g., Cermeño, Grier, and Grier,

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 election, to increase the real income of voters and boost its popularity, and depreciates it afterwards. However, his prediction that all incumbents engage in this electoral manipulation is at odds with the evidence.

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 if fixed-inconsistent regimes lead to a lower probability of re-election, and more generally if multiple exchange-rate markets affect electoral results.

References

Alesina, Alberto, and Lawrence H. Summers. Central bank independence and macroeconomic performance: some comparative evidence. *Journal of Money, Credit and Banking.* 1993. 25 (2) 151–162.

Alesina, Alberto; Alexander Wagner. Choosing (and reneging on) exchange-rate regimes. *Journal of the European Economic Association*. 2006. 4 (4) 770–799.

Baxter, Marianne; Alan Stockman. Business cycles and the exchange-rate regime: Some international evidence. *Journal of Monetary Economics*. 1989. 23 (3) 377–400.

Blomberg, Brock; Jeffry Frieden; Ernesto Stein. Sustaining fixed rates: The political economy of currency pegs in Latin America. *Journal of Applied Economics.* 2005. 8 (2) 203–225.

Bonomo, Marco; Cristina Terra. Elections and exchange rate policy cycles. *Economics & Politics*. 2005. 17 (2) 151–176.

Calvo, Guillermo A.; Leonardo Leiderman; Carmen M. Reinhart. Capital inflows and real exchange rate appreciation in Latin America: the role of external factors. *IMF Staff Papers*. 1993. 40 (1) 108-151.

Calvo, Guillermo A.; Carmen Reinhart. Fear of floating. *Quarterly Journal of Economics*. 2002. 118 (2) 379–408.

Cermeño, Rodolfo; Robin Grier; Kevin Grier. Elections, exchange rates and reform in Latin America. *Journal of Development Economics*. 2010. 92 (2) 166–174.

Edwards, Sebastian. The political economy of inflation and stabilization in developing countries. *Economic Development and Cultural Change*. 1994. 42 (2) 235–266.

Fernandez-Arias, Eduardo; Peter J. Montiel. The surge in capital inflows to developing countries: an analytical overview. *World Bank Economic Review*. 1996. 10 (1) 51–77.

Garriga, Ana; Cesar Rodriguez. More effective than we thought: Central bank independence and inflation in developing countries. *Economic Modelling*. 2020. 85 87–105.

Gavin, Michael; Roberto Perotti. Fiscal policy in Latin America. In: Ben S. Bernanke and Julio Rotemberg, eds., *NBER Macroeconomics Annual*. 1997. 12. MIT Press, Cambridge, MA.

Genberg, Hans, and Alexander K. Swoboda. Exchange rate regimes: Does what countries say matter? *IMF Staff Papers*. 2005. 52 (special issue) 129–141.

Goldfajn, Ilan; Rodrigo Valdés. The aftermath of appreciations. *The Quarterly Journal of Economics*. 1999. 114 (1) 229–262.

Hamilton, James D. Why you should never use the Hodrick-Prescott filter. *Review of Economics and Statistics*. 2018. 100 (5) 831-843.

Ilzetzki, Ethan, Carmen M. Reinhart, and Kenneth S. Rogoff. Exchange arrangements entering the twenty-first century: Which anchor will hold? *The Quarterly Journal of Economics*. 2019. 134 (2) 599–646.

Jeanne, Olivier. International reserves in emerging market countries: too much of a good thing? *Brookings papers on Economic activity*. 2007. 1 (2007) 1–79.

Jeanne, Olivier; Romain Rancière. The optimal level of international reserves for emerging market countries: A new formula and some applications. *Economic Journal*. 2011. 121 (555) 905–930.

Juhn, Grace; Paolo Mauro. Long-Run determinants of exchange rate regimes: A simple sensitivity analysis. Unpublished manuscript. 2002. NBER.

Klein, Michael W.; Nancy P. Marion. Explaining the duration of exchange-rate pegs. *Journal of Development Economics.* 1997. 54 (2) 387–404.

Levy-Yeyati, Eduardo; Federico Sturzenegger. Classifying exchange rate regimes: Deeds vs. words. *European Economic Review.* 2005. 49 (6) 1603–1635.

Levy-Yeyati, Eduardo; Federico Sturzenegger; Illiana Reggio. On the endogeneity of exchange rate regimes. *European Economic Review.* 2010. 54 659–677.

Ploeg, Frederick van der. The political economy of overvaluation. *Economic Journal.* 1989. 99 (397) 850–855.

Reinhart, Carmen; Kenneth Rogoff. The modern history of exchange rate arrangements: A reinterpretation. *Quarterly Journal of Economics.* 2004. 119 (1) 1–48.

Ravn, Morten O.; Harald Uhlig. On adjusting the Hodrick-Prescott filter for the frequency of observations. *Review of economics and statistics.* 2002. 84 (2) 371–376.

Rogoff, Kenneth; Anne Sibert. Elections and macroeconomic policy cycles. *Review of Economic Studies.* 1988. 55 (1) 1–16.

Stein, Ernesto; Jorge M Streb. Political stabilization cycles in high-inflation economies. *Journal of Development Economics.* 1998. 56 (1) 159–180.

Stein, Ernesto; Jorge M Streb. Elections and the timing of devaluations. *Journal of International Economics.* 2004. 63 (1) 119–145.

Stein, Ernesto; Jorge M Streb; Piero Ghezzi. Real exchange rate cycles around elections. *Economics & Politics.* 2005. 17 (3) 297–330.

Appendix A. Monthly de jure exchange-rate regime classification

The IMF de jure exchange-rate regime classification is reported with an annual frequency. However, the IMF *Annual Report on Exchange Arrangements and Exchange Restrictions* (AREAER) provides details on the day a country member adopted a new de jure regime classification. Identifying the day a regime change is announced is crucial to our identification strategy around the month of the change of government.

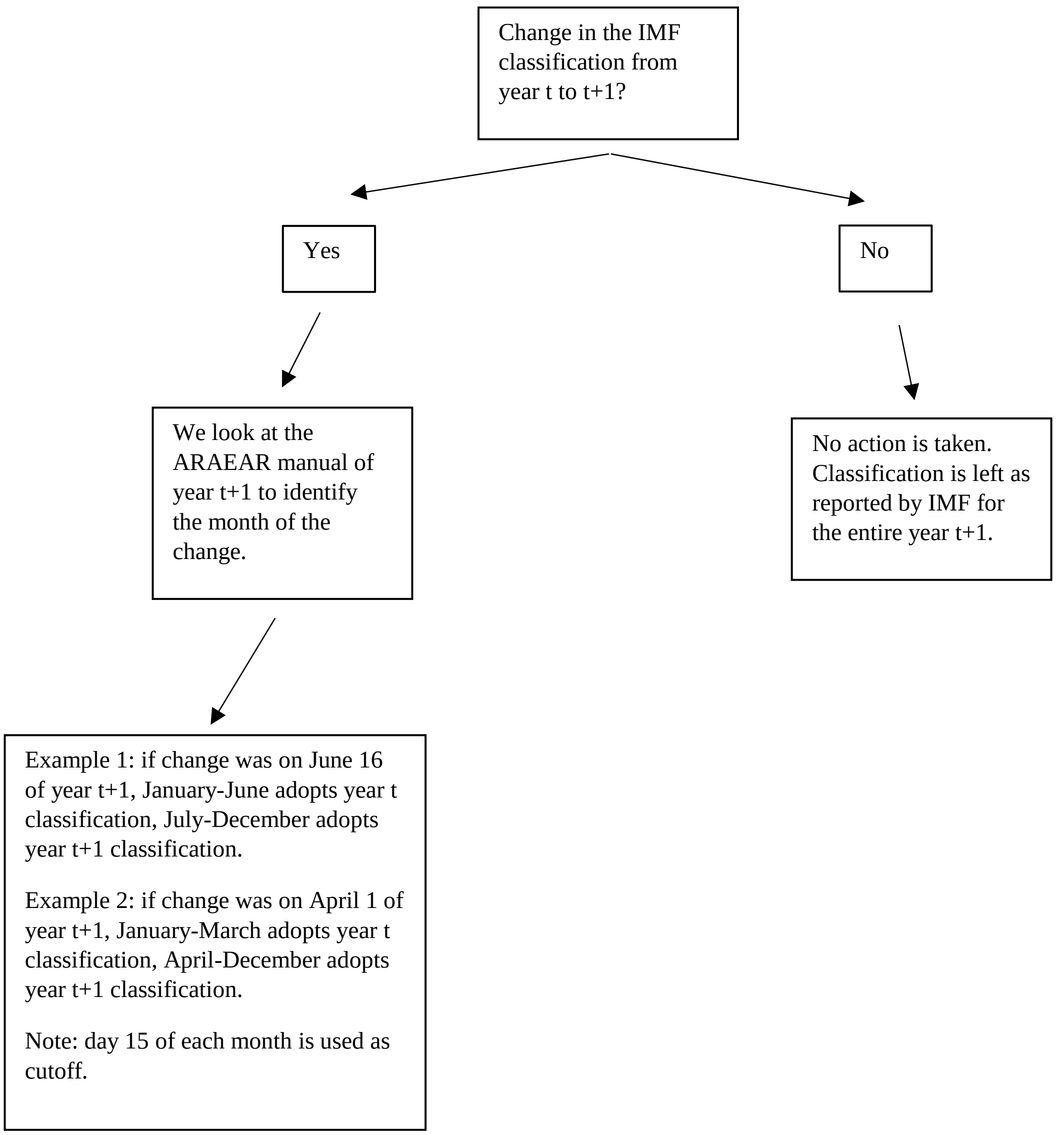
We apply the following criterion to transform the IMF classification into a monthly frequency. The IMF reports in the summary section of exchange-rate regimes the regime as of December 31st of the current year. The IMF regime classification for the entire year is based on this classification as of December 31st of each year. The first approach we follow to identify if there is a need to change the classification within the year is to look if there was a change from one year to another. If there is no change, then we assume there was no change within the year. But if there was a change, then we look for the month of the change.

If there is a change within the year, we use as the cutoff day the 15th of each month to decide what to do with the month of the change. For example, suppose a country announces a change in the regime from peg to crawl on, say, December 25th 1998. The yearly IMF classification would lead to consider the entire year as a crawl, although it corresponds only to the last week of that year, while we would rather consider that year as a peg. Suppose instead that the new classification is adopted on, say, December 14th. Then, we impute the old peg classification from January to November, leaving only December with the new crawl classification. This indicates how important it is to produce a monthly de jure classification by diving into the AERAER manuals. If the day of the change is not reported in the AREAER manual of a particular year, we then proceed to look at alternative sources. Figure A1 below describes the decision rule applied.

Finally, a terminological clarification. The month of the announcement of a regime is the month in which the country member informs the IMF of the adoption of a new official regime. However, since the regime announcement lasts until a new announcement is made, we use the term "announcement" to describe the de jure or preannounced exchange-rate regime officially in place. For example, if country X announces on, say, 1 June 1995 the adoption of a float, while it announces the adoption of a peg on, say, 1 September

1999, then the regime announcement or de jure regime from June 1995 to August 1999 is flexible, and from September 1999 on fixed.

Figure A1. Algorithm for turning annual IMF classification into monthly frequency



Details on the months of changes identified per country and year are displayed in Table A1 below.

Table A1. Monthly classification of exchange-rate regime announcements

Year	Country	Months that change	Category originally in place (replace with)	Description: text from ARAEAR unless alternate source or *authors' note* indicated
1980	Costa Rica	December	1(4)	Costa Rica informed the Fund in December that the dual exchange market established in September 1980 had been unified and that, as a temporary exchange arrangement, the colon had been allowed to float. *Authors' note: All 1980 is classified by the IMF as fixed (1), though in December a float (4) was adopted.*
1981	Costa Rica	January to March	3 (1)	During April-July, when the exchange markets were reunified on the basis of a flexible exchange rate system, the exchange arrangement belonged to the "More Flexible (Independently floating)" category. In July, although multiple exchange markets were introduced, the pegging arrangement was used with respect to two main exchange rates so that the exchange arrangement was reclassified to the category "Pegged (U.S. dollar)." In October, as a large proportion of exchange transactions began to be transacted in an exchange market where the rates were responsive to market forces, the exchange arrangement was reclassified as belonging to the category of "More Flexible (Independently floating)." *Authors' note: All 1981 is classified as managed float (3) by the IMF. Based on our review, it should have been either fix (1) or float (4).*
		April to July	3 (4)	
		August and September	3 (1)	
		October to December	3 (4)	
1982	Chile	January to June	3 (1)	In mid-June Chile changed the exchange arrangement for its peso from a U.S. dollar peg to an arrangement whereby the exchange rate for the peso would be determined on the basis of a currency basket, against which the peso was to be depreciated on a daily basis at a rate equivalent to 0.8 percent a month. The Chilean peso was therefore reclassified from the category "Pegged: U.S. Dollar" to the "Other Managed Floating" category.
	Costa Rica	January to July	3 (4)	With effect from August 9, the authorities of Costa Rica closed all authorized exchange houses and transferred all exchange transactions to the banking system. As the rate in the banking market, now the major exchange market for the Costa Rican colon, is adjusted periodically in the light of supply and demand developments, the Costa Rican colon was reclassified within the "More Flexible" category from "Independently Floating" to "Other Managed Floating."

	Uruguay	January to November	4 (3)	On November 29 the Central Bank of Uruguay discontinued its policy adopted in December 1978 of announcing exchange rates in advance and now permits the exchange rate of the new peso to be determined by supply and demand without intervention from the Central Bank. The Uruguayan new peso was accordingly reclassified within the "More Flexible" category from "Other Managed Floating" to "Independently Floating."
1983	Ecuador	January to March	3 (1)	Ecuador informed the Fund of several changes to its dual exchange market arrangements consisting of an official and a free market. On March 19, the official exchange rate of the sucre was devalued and Ecuador announced that with effect from March 23, the sucre would be depreciated on a daily basis against the U.S. dollar at a rate equivalent to 2.9 percent a month. Ecuador was therefore reclassified from the category "Pegged: U.S. Dollar" to the "More Flexible: Other Managed Floating" category.
	Jamaica	January to November	3 (1)	Jamaica informed the Fund that, with effect from November 24, the pre-existing official, CARICOM and parallel market rates had been unified into a single exchange rate. Since then the spot exchange rate has been fixed jointly by the commercial banks on a daily basis within a band prescribed by the Bank of Jamaica and reviewed on a fortnightly basis. Jamaica's exchange arrangement was reclassified from the category "Pegged: U.S. Dollar" to the "More Flexible: Other Managed Floating" category.
1984	Jamaica	January to November	4 (3)	The authorities of Jamaica notified the Fund that with effect from November 29, 1984 the exchange rate was to be permitted to float freely.
1985	Argentina	January to May	1 (3)	With effect from June 14, Argentina introduced a new monetary unit, the "austral" (A), to replace the peso argentine ($a) as the currency of Argentina. The austral has a value equal to 1,000 pesos.
	Bolivia	January to August	4 (1)	On August 29, 1985 the Government of Bolivia announced that the previous fixed exchange rate relationship between the Bolivian peso and the U.S. dollar was replaced by a system of foreign exchange auctions to be held at least twice a week by the Central Bank of Bolivia.
	Dom. Rep.	January	4 (1)	The authorities of the Dominican Republic informed the Fund that with effect from January 23, 1985, the previous dual exchange markets were unified, and all foreign exchange transactions now take place at market-determined exchange rates.
	El Salvador	January to June	3 (1)	The authorities of El Salvador informed the Fund that with effect from June 17, 1985, a number of categories of transactions have been transferred from the official to the parallel market, resulting in an effective depreciation of the colon. As a result of this change, the majority of external transactions are now conducted in the parallel market. El Salvador was thus reclassified from the category "Pegged: U.S. Dollar" to the category "More Flexible: managed Floating."

	Peru	January to July	1 (3)	Peru informed the Fund that, with effect from August 1, 1985, the Peruvian authorities depreciated the sol and fixed the exchange rate in terms of the U.S. dollar.
1986	Argentina	January to June	3 (1)	…while inflation was contained, the annual rate of inflation was still at about 100 percent by mid-1986, forcing the government to abandon the peg. Alternate source: Thirty Years of Currency Crises in Argentina. External Shocks or Domestic Fragility? Kaminsky, Mati, and Choueiri, 2009. https://home.gwu.edu/~graciela/HOME-PAGE/RESEARCH-WORK/WORKING-PAPERS/argentina.pdf
	El Salvador	January	1 (3)	El Salvador has notified the Fund that, with effect from January 22, the official exchange system has been unified, and the colon has been pegged to the U.S. dollar at a rate of C 5 = US$1.
1987	Dom. Rep.	July to October	4 (3)	June 17. The free foreign exchange market was suspended, and all foreign exchange was required to be surrendered to the Central Bank at the exchange rate fixed by the Bank. November 12. The official and unofficial parallel exchange markets were unified, with the official rate to be set at the rate prevailing in the parallel market.
1988	Dom. Rep	January	3 (4)	February 11. The freely floating exchange rate system was abolished, and a temporary exchange arrangement (to be effective for 90 days) was introduced. The Central Bank would set the exchange rate for the peso and appreciate it on a weekly basis. The Central Bank's buying and selling rates were initially set at RD$5.40 = US$1 and RD$5.45 = US$1, respectively.
		August to December	3 (1)	August 1. A new exchange system was established, under which the Central Bank would administer all foreign exchange transactions. All foreign exchange receipts were required to be surrendered, and most payments were subject to the prior approval of the Central Bank. Only foreign exchange purchases by private individuals for specified current invisibles up to US$5,000 for each transaction and US$10,000 a year would not require prior approval. The midpoint official exchange rate was established at RD$6.34 = US$1.... (Position on December 1989) The currency of the Dominican Republic is the Dominican Peso, which is pegged to the U.S. dollar, the intervention currency, at RD$6.34 = US$1. *Authors' note: The two paragraphs above (first from 1989 ARAEAR and second from 1990 ARAEAR) suggest that on August 1st, 1988, the monetary authority decided to peg the official exchange rate, and it was still sustained in December 1989.*
	Ecuador	January to August	3 (1)	August 31. The intervention market rate was changed from SI. 250 per US$1 to SI. 390 per US$1, and it was announced that the intervention rate would be depreciated by SI. 2.50 a week against the U.S. dollar.

1989	Argentina	January to December	4 (3)	December 20. The foreign exchange market was unified under a freely floating system. At the end of the month, the exchange rate depreciated to about A 2,000 per US$1.
	El Salvador	January to June	3 (1)	Until July 10, 1989, most authorized foreign exchange transactions took place in the official market, in which the Salvadoran colon was pegged to the U.S. dollar, the intervention currency, at C5.00 = US$1. Foreign exchange held in approved foreign currency deposit accounts with commercial banks, however, could be freely sold among registered account holders at a rate that diverged from the official exchange rate. On July 11, 1989, the Salvadoran Monetary Board modified the foreign exchange system by transferring the equivalent of 75-80 percent of current transactions to the free bank foreign exchange market operated by commercial banks. As a result, and until December 31,1989, transactions in the official market were limited to foreign exchange receipts from loans and transfers to the Government and foreign exchange sales for imports of petroleum products and public sector debt-service payments.
	Guatemala	January to October	4 (1)	November 3. A floating exchange rate system was introduced. The Bank of Guatemala ceased to provide exchange rate guarantees for prefinancing of exports.
	Paraguay	January and February	3 (1)	February 27. A unified floating exchange rate system applicable to all current and capital transactions was introduced, and the system of minimum surrender export prices was eliminated.
	Peru	January to June	3 (1)	June 8. A crawling peg system was introduced, and the Central Bank announced that, effective June 9, it would inform commercial banks about the prevailing exchange rate on a daily basis.
	Venezuela	January and February	4 (1)	March 13. The system of multiple exchange rates was replaced by a unified, interbank market system under which the exchange rate would be allowed to be determined by market forces. Virtually all forms of exchange control were eliminated.
1990	Brazil	January to March	4 (3)	March 16. The Central Bank introduced an official foreign exchange interbank market. The following transactions would take place in this market where the exchange rate is to be determined freely: import and export transactions; profit and divided remittances; capital repatriation; debt-service payments; and approved foreign investments.
	El Salvador	January to May	4 (3)	June 1. The process of unification of the exchange system, which began in May, was completed. The Central Reserve Bank's official exchange rate would henceforth be determined on a weekly basis as a simple average of the exchange rate prevailing in the exchange markets of commercial banks and exchange houses during the previous week.

	Honduras	January and February	3 (1)	March 13. A new exchange rate system came into operation. With the exception of debt-service payments by the Central Government, and trade with other Central American countries, all foreign exchange transactions will be conducted in an interbank market in which the exchange rate will be allowed to float within margins of 2.5 percent around a reference exchange rate. The latter will be determined by the Central Bank of Honduras and will be adjusted periodically, taking into account market conditions as reflected in the actual exchange rates prevailing in the transactions of the commercial banks and other macroeconomic factors.
	Jamaica	January to September	4 (3)	September 17. An interbank foreign exchange market (Interbank Foreign Exchange Trading System) was established, in which the exchange rate of the Jamaica dollar is market determined. Commercial banks are authorized dealers, and they are allowed to negotiate the buying and selling rates for Canadian dollars, pounds sterling, and U.S. dollars.
	Peru	January to July	4 (3)	August 9. Regulations were enacted to unify and float the exchange rate. Exporters would receive all of their surrendered proceeds in freely disposable foreign exchange certificates (CLDs), which could be used to pay for all imports, nonfactor services, and the servicing of private unguaranteed debt.
1991	Argentina	January to March	1 (4)	March 19. The unified floating exchange rate system was replaced by a peg to the U.S. dollar, and the lower end of the intervention band was raised to narrow the band to A 9,700-A 10,000 per US$1.
	Dom. Rep.	January	4 (1)	January 24. A new exchange system, consisting of an official rate that would apply to all exports and essential imports to be transacted through the Central Bank and an interbank market in which the exchange rate would be freely determined and to be applied to all other transactions was introduced.
	Guyana	January to September	4 (1)	September 30. The official exchange rate was abolished. The Bank of Guyana would apply the average telegraphic transfer cambio rate to transactions channeled through it (that is, export receipts from sugar and bauxite and payments for imports of fuel and official debt-service payments). The Bank of Guyana would quote weekly rates for the CARICOM currencies under the bilateral arrangements with the central banks of the CARICOM.

1992	Costa Rica	January and February	4 (3)	March 1. The system of periodic devaluations of the exchange rate was discontinued, and a new exchange rate arrangement was introduced, under which commercial banks and other financial institutions under the supervision of the Auditor General of Financial Institutions would be allowed to buy and sell foreign exchange at freely negotiated rates immediately for current transactions (including debt-service payments) and as of June 1, 1992 for other capital transactions. Ninety percent of export receipts would have to be sold to the authorized entities at market rates and 25 percent of all foreign exchange purchased by those entities would have to be sold to the Central Bank, which would meet its additional requirements from the market and would intervene, as needed, to smooth out large swings in the exchange rate to meet the foreign reserves target. The Government, its institutions, and public enterprises would purchase and sell foreign exchange through the Central Bank, at the reference buying or selling exchange rate. The reference rates would be calculated by the Central Bank as the weighted average of the respective exchange rates for the foreign exchange transactions performed by the authorized institutions during the previous working day.
	Honduras	January	4 (3)	February 13. The National Assembly approved legislation that legalized the establishment of foreign exchange houses and authorized them to buy and sell foreign exchange at freely negotiated rates for all transactions, except for purchases of export proceeds. The Central Bank would purchase export proceeds in the interbank market at the average exchange rate of the preceding five days in this market. Commercial banks would continue to buy and sell foreign exchange at the interbank market rate and would be required to transfer 40 percent of their purchases to the Central Bank. The Central Bank would meet its remaining foreign exchange requirements with purchases in the market and would intervene, if necessary, to smooth out fluctuations in the market exchange rate.
1993	Nicaragua	None	3 (1)	January 10. The official rate of the cordoba was devalued from C$5 to C$6 per US$1 and would be depreciated daily at an annual rate of 5 percent.
	Uruguay	January and February	3 (4)	The Uruguayan peso replaced the new Uruguayan peso at the ratio of 1,000 to 1 on March 1,1993. The Central Bank of Uruguay intervenes to ensure that the exchange rate will remain within the higher and lower limits of the exchange rate band (currently 7 percent) and allows market rates to float freely within these limits. The Central Bank periodically announces its intervention buying and selling rates.

1994	Brazil	January to June	3 (4)	July 1. A new currency, the real (R$), was introduced to replace the cruzeiro real and the URV at the conversion rate of 2,750 cruzeiro real to R$l. All contracts denominated in URVs were converted to reais at a conversion rate of 1:1. The Central Bank would set a floor of R$l per US$1 in the commercial market and be committed to use its international reserves to maintain the floor for an indefinite period. Otherwise, the exchange rate would be determined by market forces.
	Dom. Rep.	January to August	3 (4)	September 7. The Monetary Board announced that the official exchange rate would be set weekly on the basis of the average of the previous week's exchange rates in the interbank market; previously, the official exchange rate was set on a daily basis to reflect the previous day's exchange rates in the interbank market.
	Honduras	January to July	3 (4)	July 17. The interbank foreign exchange system was suspended temporarily, and a foreign exchange auction system was introduced. Under the auction system, banks and exchange houses are required to sell all their daily foreign exchange purchases to the Central Bank, which auctions at least 60 percent of the foreign exchange it has purchased at a "base price" that it has set. The base price is changed each time the reference exchange rate (the weighted average of successful bids) differs in the same direction from the base prices for 15 consecutive auctions. The bid (buying) price in the auctions cannot be different from the base price by more than 1 percent, and the maximum bid in an auction cannot exceed $200,000.
	Mexico	January to December	4 (3)	December 22. The peso was allowed to float, and the exchange regime based on the crawling peg was abandoned.
	Venezuela	January to June	1 (4)	July 11. The exchange market reopened under a pegged exchange regime, and the exchange rate was set at Bs 170 per US$1.
1995	Costa Rica	January to May	3 (4)	May 30. The Central Bank announced that the spread between the buying and selling rates of authorized banks and financial institutions cannot exceed 0.56 percent.
1996	Venezuela	January to June	3 (1)	On July 8, 1996, an exchange rate band of ±7.5% around the central rate was introduced with an initial rate of Bs 470 per $1. During the remainder of 1996, the central rate was adjusted in line with the inflation target for the first quarter of 1996, i.e., 1.5% a month.
1998	Bolivia	January to December	3 (4)	December 31. The exchange rate arrangement of Bolivia was reclassified to "managed floating" from "independently floating."
1999	Chile	September to December	3 (4)	On September 2, 1999, the central bank suspended the crawling band and allowed the peso to float. Thus, the exchange rate arrangement was reclassified to the category independently floating from the category crawling band. *Authors' note: All 1999 is classified as managed float (3) by the IMF. Based on our review, September to December correspond to float (4).*

Brazil	January	4 (3)	Before February 1, 1999, there were two official exchange markets…. The Central Bank of Brazil (CBB) established an adjustable band for the external value of the national currency…. On February 1, 1999, the exchange rate was unified.
Colombia	January to September	4 (3)	On September 25, 1999, the BR abandoned the crawling band and allowed the peso to float.
Ecuador	January	4 (3)	The dual exchange rate system was unified on February 12, 1999, when the sucre was floated.

Appendix B. Construction of marginal effects based on estimators from Table 1

This Appendix first discusses the estimation strategy behind Table 1 in Section IV. It then shows the marginal effects of the ordered and multinomial logit models based on the estimators displayed in Table 1. Regarding the estimation strategy, in order to identify accurately what happen with the probabilities per category for the ordered logit models shown in Table 1, Columns 1 and 2, the marginal effects should be calculated around particular level for the covariates, which is usually the mean of the covariates. We first identify a score function evaluated at the mean of the covariates outside the electoral window: $S_0 = \bar{\mathbf{x}}'\hat{\boldsymbol{\beta}}$**.** Given the estimated score, we define the probabilities of the regimes as follows:

$$P(regime = 1[\text{fixed}] \mid \bar{\mathbf{x}}) = (1 + \exp(S_0 - k_1))^{-1},$$

$$P(regime = 2[\text{crawl}] \mid \bar{\mathbf{x}}) = (1 + \exp(S_0 - k_2))^{-1} - (1 + \exp(S_0 - k_1))^{-1},$$

$$P(regime = 3[\text{manage}] \mid \bar{\mathbf{x}}) = (1 + \exp(S_0 - k_3))^{-1} - (1 + \exp(S_0 - k_2))^{-1},$$

$$P(regime = 4[\text{float}] \mid \bar{\mathbf{x}}) = 1 - (1 + \exp(S_0 - k_3))^{-1}, \quad \text{(A1)}$$

where $k_3 > k_2 > k_1$ are the cutoffs estimated in the model.[27] If, for example, $\hat{\beta}_1 > 0$, then $\frac{\partial P(regime=1[\text{fixed}] \mid \bar{\mathbf{x}})}{\partial x_1} < 0$ and $\frac{\partial P(regime=4[\text{float}] \mid \bar{\mathbf{x}})}{\partial x_1} > 0$, but it is not possible to identify the signs of $\frac{\partial P(regime=2[\text{crawl}] \mid \bar{\mathbf{x}})}{\partial x_1}$ and $\frac{\partial P(regime=3[\text{manage}] \mid \bar{\mathbf{x}})}{\partial x_1}$ when $\hat{\beta}_1 > 0$. In Table B1, Panel A, below we present the full set of marginal effects for each of the categories for the ordered logit model.

The interpretation of the estimators in the multinomial logit model is different from the ordered logit model. By selecting inconsistent de jure fixed regimes as the base category ($y = 3$), the estimated probabilities are as follows:

$$P(regime = 1[\text{consistent fixed}] \mid \bar{\mathbf{x}}) = \frac{\exp(\bar{\mathbf{x}}'\hat{\boldsymbol{\beta}}_1)}{1+\exp(\bar{\mathbf{x}}'\hat{\boldsymbol{\beta}}_1)+\exp(\bar{\mathbf{x}}'\hat{\boldsymbol{\beta}}_2)+\exp(\bar{\mathbf{x}}'\hat{\boldsymbol{\beta}}_4)},$$

[27] For N categories there are $N - 1$ cutoffs.

$$P(regime = 2[\text{inconsistent flex}] \mid \bar{\mathbf{x}}) = \frac{\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_2)}{1+\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_1)+\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_2)+\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_4)},$$

$$P(regime = 3[\text{inconsistent fixed}] \mid \bar{\mathbf{x}}) = \frac{1}{1+\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_1)+\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_2)+\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_4)},$$

$$P(regime = 4[\text{consistent flex}] \mid \bar{\mathbf{x}}) = \frac{\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_4)}{1+\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_1)+\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_2)+\exp(\bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_4)}. \tag{A2}$$

Note that $\ln\left(\frac{P(regime=j \mid \bar{\mathbf{x}})}{P(regime=3 \mid \bar{\mathbf{x}})}\right) = \bar{\mathbf{x}}'\widehat{\boldsymbol{\beta}}_j$ for $j = 1,2,4$. Now for a vector of marginal increase, $\Delta\mathbf{x}$, we get $\ln\left(\frac{P(regime=j \mid \bar{\mathbf{x}}+\Delta\mathbf{x})}{P(regime=3 \mid \bar{\mathbf{x}}+\Delta\mathbf{x})}\right) - \ln\left(\frac{P(regime=j \mid \bar{\mathbf{x}})}{P(regime=3 \mid \bar{\mathbf{x}})}\right) = \Delta\mathbf{x}'\widehat{\boldsymbol{\beta}}_j$**.** If we analyze only an increase in the covariate 1 equal to $\Delta x_1 = 1$, keeping all the other covariates constant at the mean values, then we get $\ln\left(\frac{P(regime=j \mid \bar{\mathbf{x}}+\mathbf{1})}{P(regime=3 \mid \bar{\mathbf{x}}+\mathbf{1})}\right) - \ln\left(\frac{P(regime=j \mid \bar{\mathbf{x}})}{P(regime=3 \mid \bar{\mathbf{x}})}\right) = \hat{\beta}_{1j}$, which is precisely what is called the "relative risk ratio". In other words, each coefficient is understood as the increase in the probability of category $j = 1,2,4$ in relation to category 3 (base) if $\hat{\beta}_{1j} > 0$, given a marginal increase of x_{1j}. Note that for two non-based categories $j \neq k$,

$$\ln\left(\frac{P(regime=j \mid \bar{\mathbf{x}}+\mathbf{1})}{P(regime=k \mid \bar{\mathbf{x}}+\mathbf{1})}\right) - \ln\left(\frac{P(regime=j \mid \bar{\mathbf{x}})}{P(regime=k \mid \bar{\mathbf{x}})}\right) = \hat{\beta}_{1j} - \hat{\beta}_{1k},$$

which indicates that we can also compare categories beyond the base category by comparing the estimators of categories j and k for the same covariate 1. We also provide the full set of marginal effects in Table B1, Panel B.

The marginal effects of the ordered logit and multinomial logit models estimated and displayed in Table 1 are shown in Table B1. The table shows the change in the probability of all the categories for a marginal increase in a particular variable, evaluated at the mean values of all covariates and at 0 for all the dummy variables (*Default* and **govch**s). Take for instance the coefficient of terms of trade *ToT* (-2.564***) in the RR de facto regime classification of Table 1, Column 2: Panel A shows that a 1% increase of *ToT* leads the probability of fix and crawl regimes to increase by 0.385 p.p. and 0.242 p.p., while manage and float decrease by 0.36 p.p. and 0.267 p.p. For dummy variables, the marginal increases are calculated by estimating the difference $\mathrm{P}(regime\ Y_{it} = y|\ \bar{\mathbf{X}}, D = 1) - \mathrm{P}(regime\ Y_{it} = y|\ \bar{\mathbf{X}}, D = 0)$. Standard errors are calculated by applying the Delta Method. We could not calculate the marginal effect under crawl with the IMF de jure regimes because the underlying probability of this category is zero, given that no preannounced crawl was registered in our sample.

Table B1. Marginal effects of ordered and multinomial logit models in Table 1

Panel A. Marginal effect from ordered logit model				
IMF de Jure regime:	Fix	Band/crawl	Manage	Float
$ln\ Portfolio_{t-12}$	-0.008*	-	0.000	0.008*
$ln\ Foreign.Liab.pc_{t-12}$	-0.002	-	0.000	0.002
$ln\ Size_{t-12}$	0.001	-	0.000	-0.001
$ln\ ToT_{t-12}$	0.114	-	-0.003	-0.111
$ln\ U.S.Interest_{t-1}$	0.404*	-	-0.011	-0.393*
$ln\ Openness_{t-1}$	-0.080*	-	0.002	0.078*
$Default_{t-12}$	-0.085*	-	-0.030*	0.115*
$govch(-3)^q$	-0.008	-	0.000	0.008
$govch(-2)^q$	-0.021	-	-0.001	0.022
$govch(-1)^q$	-0.027	-	-0.002	0.029
$govch(0)^q$	-0.033	-	-0.003	0.036
$govch(+1)^q$	-0.086*	-	-0.031	0.117*
$govch(+2)^q$	-0.073*	-	-0.020	0.093*
$govch(+3)^q$	-0.054*	-	-0.010	0.064
$govch(+4)^q$	-0.061*	-	-0.013	0.074*
RR de facto regime:	Fix	Band/crawl	Manage	Float
$ln\ Portfolio_{t-12}$	0.007*	0.004*	-0.006*	-0.005*
$ln\ Foreign.Liab.pc_{t-12}$	0.007*	0.005*	-0.007*	-0.005*
$ln\ Size_{t-12}$	0.006	0.004	-0.006	-0.004
$ln\ ToT_{t-12}$	0.385*	0.242*	-0.360*	-0.267*
$ln\ U.S.Interest_{t-1}$	0.003	0.002	-0.002	-0.002
$ln\ Openness_{t-1}$	0.045*	0.028*	-0.042*	-0.031*
$Default_{t-12}$	-0.139*	-0.216*	0.081*	0.274*
$govch(-3)^q$	-0.040*	-0.033	0.038*	0.035
$govch(-2)^q$	-0.057*	-0.051*	0.055*	0.053*
$govch(-1)^q$	-0.058*	-0.054*	0.056*	0.056*
$govch(0)^q$	-0.060*	-0.056*	0.058*	0.058*
$govch(+1)^q$	-0.077*	-0.079*	0.073*	0.083*
$govch(+2)^q$	-0.063*	-0.060*	0.061*	0.062*
$govch(+3)^q$	-0.044*	-0.036	0.042*	0.038
$govch(+4)^q$	-0.035	-0.028	0.034	0.029

Panel B. Marginal effect from multinomial logit model				
Regimes:	Fixed-Consistent	Flexible-Inconsistent	Fixed-Inconsistent	Flexible-Consistent
$ln\ Portfolio_{t-12}$	0.012	0.005	-0.006*	-0.011*
$ln\ Foreign.Liab.pc_{t-12}$	0.007*	-0.009*	-0.011*	0.013*
$ln\ Size_{t-12}$	0.025*	-0.032*	-0.058*	0.064*
$ln\ ToT_{t-12}$	-0.322*	0.982*	-0.204*	-0.455*
$ln\ U.S.Interest_{t-1}$	0.530*	-0.456*	0.084*	-0.157*
$ln\ Openness_{t-1}$	-0.026	-0.047*	-0.04*	0.113*
$Default_{t-12}$	-0.355*	-0.117*	-0.019*	0.492*
$govch(-3)^q$	-0.083	-0.020	-0.006	0.109*
$govch(-2)^q$	-0.136*	-0.011	0.005	0.141*
$govch(-1)^q$	-0.134*	-0.223	0.009	0.147*
$govch(0)^q$	-0.143*	-0.007	0.009	0.141*
$govch(+1)^q$	-0.129*	-0.04	-0.016	0.185*
$govch(+2)^q$	-0.095	-0.020	-0.016	0.131*
$govch(+3)^q$	-0.002	-0.037	-0.010	0.049
$govch(+4)^q$	0.009	-0.002	-0.004	-0.003

Notes: For ordered logit models from Panel A, marginal effects calculated with set of equations (A1) as $\frac{\partial P(regime = \mathrm{y} \mid \bar{\mathbf{x}})}{\partial x_j}$ for $y =$ {Fix, Crawl, Manage, Float}, with all time-varying covariates at their mean values ($\bar{x}_j$). No observations reported in our dataset for category "crawl" under IMF de jure regime, which prevents us from calculating marginal effects (estimated probability is zero). For dummy variables (D_j), marginal effects calculated as $P(regime = \mathrm{y} \mid \bar{\mathbf{x}},\ D_j = 1) - P(regime = \mathrm{y} \mid \bar{\mathbf{x}},\ D_j = 0)$. Similar approach used to calculate marginal effects of multinomial logit model with set of equations (A2) for categories $y =$ {Fixed-consistent, Flexible-inconsistent, Fixed-inconsistent, Flexible-consistent}. Standard errors calculated by applying the Delta Method. 95% significance or higher denoted with (*).

Appendix C. Real exchange-rate misalignments with Hamilton filter

Hamilton (2018) has shown that the HP filter introduces undesired spurious dynamic relations. As a robustness procedure, we employ here his method to show that our results under the HP filter for the misalignments of the real exchange rate of Section V are not substantially driven by such spurious dynamics. That is, we find here qualitatively similar results under the Hamilton method.

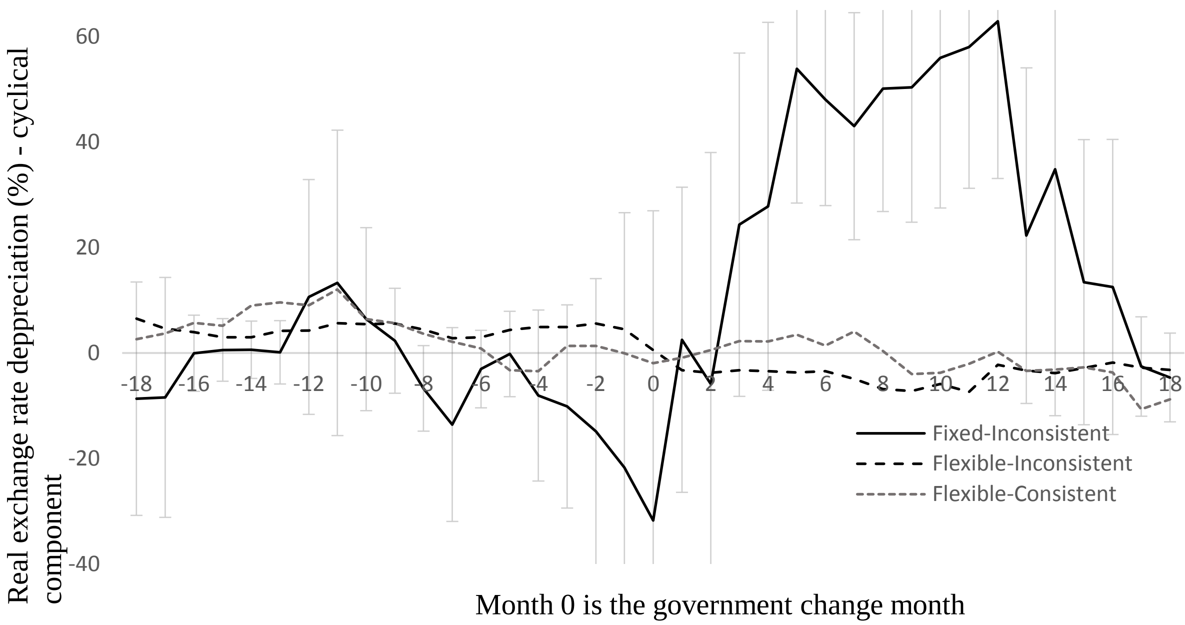
As we mentioned in the main text, we have identified that the first difference of the natural log of Real exchange rate ($\Delta \ln RER$), our main dependent variable, is $I(0)$. We have run Dickey Fuller test on $\ln RER$ and $\Delta \ln RER$ for each country to confirm that they are $I(1)$ and $I(0)$, respectively. We complement the DF test results with a visual inspection of the time series per country of $\Delta \ln RER$, where we have corroborated visually that $\Delta \ln RER$ is $I(0)$.[28]

Given that $\Delta^d \ln RER$ is $I(0)$ for $d = 1$, the Hamilton filter predicts that the cyclical component of $\ln RER$ is asymptotically equivalent to

$$\ln RER_{t+h}^{cycle} = \ln RER_{t+h} - \ln RER_t.$$

We decided to use $h = 12$ in monthly data in order to avoid losing so many observations due to the differentiation process. Below we see the estimation results of Equation (5) in Figure C1 for the Hamilton Filter, as we have done in Figure 5 for the HP filter.

Figure C1. Real exchange rate misalignments around government changes with Hamilton filter



Notes: Graphic representation of the estimators of Equation (5) for cyclical component of RER, estimated by detrending the RER series with Hamilton filter (with differential parameter h=12). Results are relative to fixed-consistent episodes. Vertical bars represent 95% confidence intervals of estimators based on robust standard errors for fixed-inconsistent estimators around the government change date.

[28] Results of DF test and time series graphs are available on request.

We observe that the patterns in Figure 5, for the HP filter, and Figure C1, for the Hamilton filter, are similar. There are still some discrepancies. While the 5% history threshold is hit at month 10 before the government change for the HP filter, here it is hit at month -4. For the HP filter the history/peak stage lasts 10 months with an appreciation peak of 37% at the government change month, while in here it only lasts 4 months with an appreciation peak of 32%, also at government change month. The peak/end stage lasts 3 months with the HP filter in Figure 5, while here it lasts between 1 and 2 months. We also observe an overshooting here, with two undervaluation peaks, one at month +5, and the other one at month +11. The second peak reaches a 63% undervaluation, which is corrected in 6 months. Something similar happened with the HP filter of Figure 5, where the peak at month +6 was corrected smoothly in a 12 month period, doubling the correction found with the Hamilton filter.

In Table C1, we also show the six-month collapsed dummy variables to capture average medium-term behavior, as we have done for the HP filter in Table 3. For direct comparison purposes, in Column 1 we paste results for the HP filter, which are also available at Table 3 in the main text, while Column 2 displays the result under the Hamilton filter. As we explained in the main text, we observe that for the fixed-inconsistent regimes the six-month average overvaluation hits a six-month average of 25.5% during the last semester before government change for the HP filter, while for the Hamilton filter such overvaluation hits a six-month average of 14% for the last semester before government change.

The overvaluation is reverted during the first semester after government change and we observe a six-month average of 16% undervaluation during the second semester after government change for HP filter. This overvaluation is exacerbated under the Hamilton filter, to 25% and 54% in the first and second semester after government change, respectively.

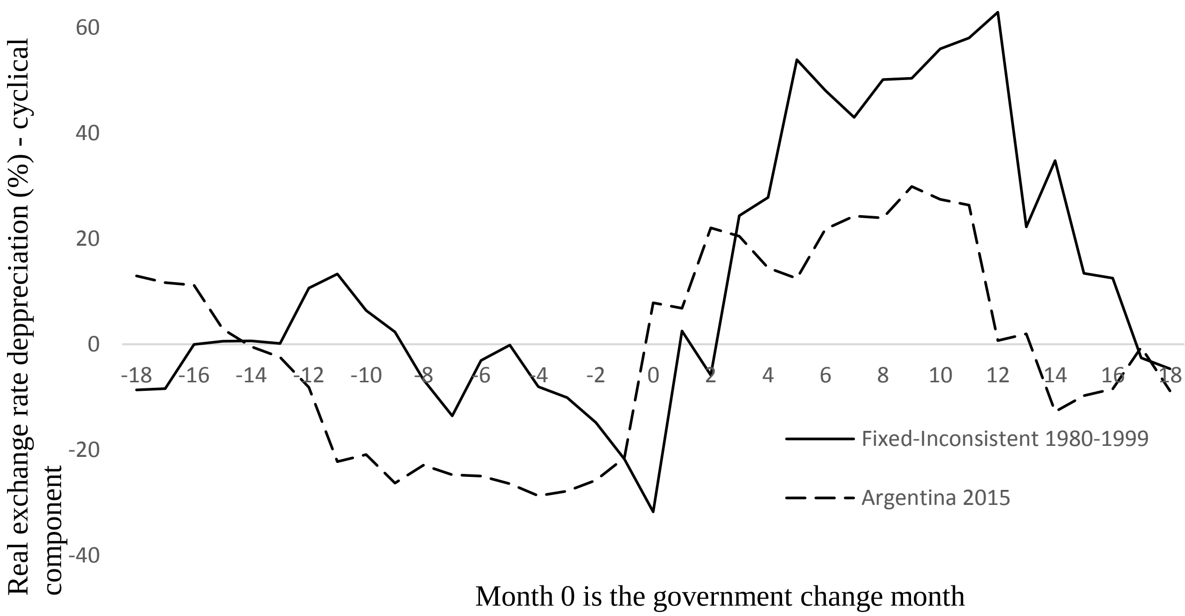
Table C1. Real exchange rate misalignment: six-month dummy variables, HP and Hamilton filters

Dependent variable: $ln(RER_{it})_{cycle}$	(1) HP-filter		(2) Hamilton Filter	
$govch(-2)^s$	-1.5133	[1.008]	-2.5633**	[1.186]
$govch(-1)^s$	-2.6879***	[0.962]	-1.2911	[1.208]
$govch(0)^s$	-3.0408***	[0.934]	0.7652	[1.366]
$govch(+1)^s$	0.9747	[1.070]	5.0749**	[2.234]
$govch(+2)^s$	1.0024	[1.035]	5.4999***	[1.966]
$govch(+3)^s$	-0.4071	[0.926]	3.1308***	[1.051]
$govchFI(-2)^s$	-4.5510**	[2.191]	0.5835	[3.075]
$govchFI(-1)^s$	-12.3445***	[3.038]	-0.1517	[3.747]
$govchFI(0)^s$	-25.4982***	[5.556]	-14.0769**	[7.143]
$govchFI(+1)^s$	2.9813	[6.359]	25.1382***	[7.530]
$govchFI(+2)^s$	16.2924***	[2.226]	54.1979***	[5.572]
$govchFI(+3)^s$	9.7655***	[1.305]	12.1803**	[6.062]
$govchFEI(-2)^s$	1.7333	[1.090]	3.8436**	[1.654]
$govchFEI(-1)^s$	3.7338***	[1.004]	4.5296***	[1.594]
$govchFEI(0)^s$	5.0865***	[1.012]	4.2582**	[1.659]
$govchFEI(+1)^s$	-0.9429	[1.046]	-3.6386	[2.251]
$govchFEI(+2)^s$	-2.2551**	[1.128]	-5.7436***	[2.034]
$govchFEI(+3)^s$	-1.2326	[1.021]	-2.9709**	[1.247]
$govchFEC(-2)^s$	0.4289	[1.283]	7.1340***	[1.803]
$govchFEC(-1)^s$	1.6642	[1.582]	5.1829*	[2.732]
$govchFEC(0)^s$	2.8312	[2.096]	-0.9248	[2.955]
$govchFEC(+1)^s$	2.1107	[1.710]	1.5014	[2.885]
$govchFEC(+2)^s$	1.3806	[1.474]	-0.8325	[2.812]
$govchFEC(+3)^s$	0.4493	[1.295]	-5.3511**	[2.410]
Observations	2,127		2,127	
R-squared	0.119		0.165	

Notes: Estimation of Equation (5) for cyclical component of RER using six-month dummy variables, detrending the RER series with Hodrick-Prescott filter (smoothing parameter of 129,600) in Column 1, and with Hamilton filter (for $h = 12$) in Column 2. Reduced sample used. Results relative to fixed-consistent episodes. Non-democratic episodes excluded based on Polity IV Project. Dollarization episodes also excluded. OLS used for estimations. Robust standard errors reported in brackets at right of each estimator. (*) [**] {***} stands for significance at (10%) [5%] {1%}.

On the other hand, in our comparison between 1980-1999 versus 2000-2016 periods of Section VI, we observe that the patterns in Figure 6, for the HP filter, and Figure C2, for the Hamilton filter, are similar.

Figure C2. Real exchange rate misalignments in 1980-1999 period and Argentina 2015 with Hamilton filter



Notes: Graphic representation of estimators of Equation (5) for cyclical component of RER for 1980-1999 period, detrending RER series with Hamilton (2018) filter ($h = 12$). Results shown for fixed-inconsistent regime (solid line) and Argentina 2015 general election (stripped line).

Appendix D. Allowing for conditional exchange rate volatility

D1. Robustness analysis of exchange-rate variation around government changes in Table 2

Exchange rate volatility may be particularly high around general elections and government changes. Alesina and Wagner (2006) also note that macroeconomic instability could be highest during fear of pegging regimes. Despite the inclusion of time-varying controls listed and included under Equation (2) – which should partially address conditional volatility–, we might not identify such instability accurately. We are not trying to study the volatility (or uncertainty) of present or future exchange rates here. Our question is whether an alternative variance structure does not substantially bias the results of our mean equations, where a traditional robust variance structure is employed. In other words, do our results in the main text survive the inclusion of bias correction for conditional volatility? To estimate macroeconomic volatility, we allow the variance to be conditional on government change and exchange-rate regimes, where uncertainty due to electoral process (especially in our fixed-inconsistent category) may be higher than in non-electoral years. This is a straightforward way to model macroeconomic instability through the volatility of the exchange rate. The results, shown in Table D1, Columns (1) and (2), allow examining the robustness of the estimates in Columns (2) and (3) of Table 2, respectively.

Table D1. Exchange rate variation around government changes with conditional variance

Dependent variable $\Delta ln\, RER$	(1)		(2)	
Type of sample	Reduced		Extended	
Type of dummy variable	$x = q$	$x = a$	$x = q$	$x = a$
Equation	Mean	Variance	Mean	Variance
$govch(-3)^x$	0.101		-0.316	
$govch(-2)^x$	0.390**		-0.383	
$govch(-1)^x$	-0.125		-0.572	
$govch(0)^x$	-0.652	0.813	-0.644*	-0.885
$govch(+1)^x$	0.383	-0.503	-0.656**	-1.304**
$govch(+2)^x$	0.660***		-1.038***	
$govch(+3)^x$	-0.163		-0.610**	
$govch(+4)^x$	0.082		-0.333	
$govchFI(-3)^x$	-0.187		2.654**	
$govchFI(-2)^x$	-0.461		-0.505**	
$govchFI(-1)^x$	1.443*		0.474	
$govchFI(0)^x$	4.100***	-0.759	3.126***	0.347
$govchFI(+1)^x$	0.211	0.873	7.137***	2.128
$govchFI(+2)^x$	11.278***		8.454***	
$govchFI(+3)^x$	-0.543*		-1.007*	
$govchFI(+4)^x$	1.026		4.112	
$govchFEI(-3)^x$	0.144		-0.244	
$govchFEI(-2)^x$	0.146		-1.028***	
$govchFEI(-1)^x$	0.398*		0.087	
$govchFEI(0)^x$	0.387	-2.381***	-0.249	-2.179***
$govchFEI(+1)^x$	-0.421	-1.434***	0.400**	-1.540***
$govchFEI(+2)^x$	-0.817***		0.231	
$govchFEI(+3)^x$	-0.161		-0.298	
$govchFEI(+4)^x$	0.285		-0.327	
$govchFEC(-3)^x$	0.270		-0.059	
$govchFEC(-2)^x$	1.734		1.158	
$govchFEC(-1)^x$	2.117**		1.486*	
$govchFEC(0)^x$	2.986	0.989	1.142*	0.676

$govchFEC(+1)^x$	0.477	1.846**	0.778**	0.672
$govchFEC(+2)^x$	0.623		0.689*	
$govchFEC(+3)^x$	0.303		0.569	
$govchFEC(+4)^x$	-0.148		-0.399	
$L1.arch$		2.695***		3.798***
Observations	2,236		4,008	
Linear Combination 1	-3.888***		4.011***	
Linear Combination 2	2.974*		5.995***	
Linear Combination 3	1.955		3.830***	
Linear Combination 4	1.770*		3.237***	

Notes: Estimation of Equation (2). Reduced sample includes Argentina, Bolivia, Brazil, Colombia, Costa Rica, Dominican Republic, Ecuador, Guatemala, Honduras, Mexico, Nicaragua, Panama, Peru, Uruguay, and Venezuela over 1980-1999 period. Extended sample also includes Barbados, Chile, El Salvador, Guyana, Jamaica, and Paraguay. *FI*, *FEI*, and *FEC* stand for fixed-inconsistent, flexible-inconsistent, and flexible-consistent regimes; flexible-consistent is omitted category. Non-democratic episodes excluded based on Polity IV Project. Dollarization episodes also excluded. MLE-ARCH used for estimations. Linear Combination k = 1,2,3,4: $govchFI(+1)^q - govchFI(0)^q$, $\frac{1}{2}(govchFI(+2)^q + govchFI(+1)^q - govchFI(0)^q - govchFI(-1)^q)$, $\frac{1}{3}(govchFI(+3)^q + govchFI(+2)^q + govchFI(+1)^q - govchFI(0)^q - govchFI(-1)^q - govchFI(-2)^q)$, $\frac{1}{4}(govchFI(+4)^q + govchFI(+3)^q + govchFI(+2)^q + govchFI(+1)^q - govchFI(0)^q - govchFI(-1)^q - govchFI(-2)^q - govchFI(-3)^q)$. Robust standard errors not reported due to space limitation. (*) [**] {***} stands for significance at (10%) [5%] {1%}.

The dummy variables of the mean equations are quarterly based, while the ones from the variance equation are annually based, because the algorithms used to reach convergence in an ARCH setting fail when the amount of variables in both equations, mean and variance, are high. For the reduced sample in Column 1, the coefficients of the linear combinations at the bottom of Table D1 fall in relation to those of Table 2, Column 2. However, Linear Combinations (LC) 2 and 4 are still significantly positive, although the magnitudes are considerably smaller (e.g., the value of LC2 is 2.974^{*} here and 15.38^{**} in Table 2, Column 2). When we employ the extended sample instead, the results of the mean equation under the conditional variance structure are also a bit smaller than those in Table 2, Column 3, but the statistical significance is higher. For example, LC1 is 4.011^{***} here and 12.47^{**} in Table 2, Column 3. Moreover, for medium run differences, such as LC3 or LC4, the coefficients are very similar in both Table 2 and Table D1.

We briefly discuss now the findings of the variance equation. Since results from both samples, reduced in Column 1 and extended in Column 2 of Table D1, show very similar patterns, we only describe the results of the variance equation of Column 2. For the fixed-inconsistent category, $govchFI(0)^a$ and $govchFI(+1)^a$, though volatility tends to be higher after the government change, it is not significant. For the flexible-inconsistent category, volatility is significantly lower not only before government change but also afterwards. This suggests that the Calvo and Reinhart (2002) finding of fear of floating, in which the exchange rate is stabilized though reserve usages and interest rates, is particularly acute around elections. Finally, $govchFEC(0)^a$ and $govchFEC(+1)^a$ are not significant, although positive. This is totally congruent with consistent de jure flexible regimes, where the monetary authority lets the exchange rate float.

D2. Robustness analysis of real exchange rate misalignments in Table 3

As in our robustness analysis of the exchange rate dynamics under a conditional variance structure of Appendix D1, we produce the same alternative conditional variance approach in Appendix D2 for the exchange rate misalignment analysis of Section V. As in Section V, the dummy variables of the mean equations are semestral based, while we decided to use annually based dummy variables for the variance equation given that algorithms used to reach convergence in an ARCH setting fail when the amount of variables in both equations, mean and variance, are high.[29] We apply it for both filtering techniques, the Hodrick-Prescott and Hamilton (2018) filters. Results are shown below in Table D2. The estimation of Column 1 corresponds to the conditional variance robustness of the estimation in Column 1 of Table C1, where the HP filter is used for the misalignment with the traditional robust variance approach, while the estimation of Column 2 of Table D2 corresponds to Column 2 of Table C1 for the Hamilton technique.

[29] Standard errors shown in Figure 5 start to become higher within the ± 12-month window. Therefore, we decided to use 12-month based dummy variables for the variance equation instead of 18-month based dummy variables.

Table D2. Real exchange rate misalignments with six-month dummy variables: conditional variance robustness

Dependent variable $ln(RER_{it})_{cycle}$	(1)		(2)	
Filter type used	H-P		Hamilton 2018	
Type of dummy variable	$x = s$	$x = a$	$x = s$	$x = a$
Equations	Mean	Variance	Mean	Variance
$govch(-2)^x$	1.366***		1.6953	
$govch(-1)^x$	1.394**		1.5079	
$govch(0)^x$	1.607***	-1.777**	2.358*	-2.069***
$govch(+1)^x$	1.414**	-1.450***	2.079*	-2.301***
$govch(+2)^x$	1.918***		3.925***	
$govch(+3)^x$	0.341		3.287***	
$govchFI(-2)^x$	-9.903***		3.178	
$govchFI(-1)^x$	-12.573***		-5.735***	
$govchFI(0)^x$	-21.458***	3.737***	-5.129	2.846***
$govchFI(+1)^x$	13.568***	1.095	47.774***	0.062
$govchFI(+2)^x$	12.507***		45.079***	
$govchFI(+3)^x$	9.501***		-6.779***	
$govchFEI(-2)^x$	-0.397		-0.845	
$govchFEI(-1)^x$	-0.362		-0.484	
$govchFEI(0)^x$	0.427	0.382	-1.319***	-0.530
$govchFEI(+1)^x$	0.187	-1.078*	-1.928***	0.239
$govchFEI(+2)^x$	-1.152***		-4.644***	
$govchFEI(+3)^x$	-0.271		-3.446***	
$govchFEC(-2)^x$	-2.940***		-1.781**	
$govchFEC(-1)^x$	-1.789***		-3.088*	
$govchFEC(0)^x$	1.886	2.048***	-2.015	2.784**
$govchFEC(+1)^x$	3.927**	2.962***	-1.819**	1.902**
$govchFEC(+2)^x$	0.382		-3.656***	
$govchFEC(+3)^x$	0.779		-2.415**	
L1. arch		1.257***		1.234***
Observations	2,127		2,127	

Notes: Estimation of Equation (5) for cyclical component of RER using six-month dummy variables, detrending RER series with Hodrick-Prescott filter (smoothing parameter of 129,600) in Column (1), and with Hamilton (2018) filter ($h = 12$) in Column (2). Reduced sample used. MLE-ARCH used for estimations. Results relative to fixed-consistent episodes. Non-democratic episodes excluded based on Polity IV Project. Dollarization episodes also excluded. Robust standard errors not reported due to space limitation. (*) [**] {***} stands for significance at (10%) [5%] {1%}.

Using the HP filter, the misalignment of the mean equation follows a similar pattern to the one found with the traditional robust variance estimation. For example, the last semester before government change we observe an overvaluation of the real exchange rate of about 21% (i.e., $govchFI(0)^s = -21.458^{***}$), while we found one of 25% in Table 3 under the traditional robust variance estimation (i.e., $govchFI(0)^{s} = -25.4982^{***}$). However, the first semester after government change we observe a significant undervaluation of about 13% (i.e., $govchFI(+1)^{s} = 13.568^{***}$), which indicates that when controlling for the conditional variance we are able to identify earlier the sharper nominal devaluation on the mean equation that ends up in a temporary undervaluation. Under the traditional robust variance approach, this only shows up in the second semester after government change (i.e., in Table 3, $govchFI(+1)^{s} = 2.9813$ indicates a non-significant undervaluation of 3%, while during the second semester there is a significant undervaluation of $govchFI(+1)^{s} = 16.292^{***}$). Despite this, the patterns look pretty similar and consistent with our explanation in Section V, i.e., an overvalued exchange rate before government changes, followed by a quick correction done through a strong nominal devaluation, leading to overshooting that brings about an undervalued exchange rate, which is smoothly corrected later through inflation differentials.

We focus now on analyzing the variance equation. The volatility of the real exchange rate misalignment is higher for the fixed-inconsistent category the year before government changes ($govchFI(0)^{a} = 3.737^{***}$, while $govchFI(+1)^{a} = 1.095$). Although this might look like a contradiction with what we found in Appendix D1, where the volatility of the exchange rate dynamics is higher (although insignificantly so) after government changes, it is not. In the case of misalignment, volatility tends to be higher before government changes precisely because sustaining a steady nominal exchange rate before elections while levels of inflation differ across countries exacerbates the misalignment (overvaluation in this case) volatility within and between counties under the fixed-inconsistent category. After government changes, the nominal exchange rate adjustment is tailored to the level of overvaluation in each country (that is, there tends to be more volatility in the exchange rate dynamics after government change due to higher volatility in the overvaluation before government change). Very similar conclusions can be drawn for the Hamilton Filter.