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"Revisiting Dollarisation Hysteresis: Evidence From Bolivia, Turkey and Indonesia"

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Evidence from Bolivia, Turkey and Indonesia

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In this paper, we pick up three countries with different inflation experiences and dollarisation levels and we investigate whether dollarisation exhibits different reversibility patterns, as suggested by the literature. The sample includes a country that experienced hyperinflation (Bolivia), a high inflation country (Turkey) and a country that experienced moderate to low inflation (Indonesia). By providing evidence of dollarisation hysteresis in these three countries, this paper challenges the view according to which this phenomenon is confined to highly dollarised economies or to economies that experienced high inflation rates for long periods of time.

JEL Classification: E41, F41.

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Revisiting Dollarisation Hysteresis: Evidence from Bolivia, Turkey and Indonesia

In this paper, we pick up three countries with different inflation experiences and dollarisation levels and we investigate whether dollarisation exhibits different reversibility patterns, as suggested by the literature. The sample includes a country that experienced hyperinflation (Bolivia), a high inflation country (Turkey) and a country that experienced moderate to low inflation (Indonesia). By providing evidence of dollarisation hysteresis in these three countries, this paper challenges the view according to which this phenomenon is confined to highly dollarised economies or to economies that experienced high inflation rates for long periods of time.

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1. Introduction

During periods of macroeconomic and political uncertainty many developing countries experienced a partial replacement of the domestic currency by a foreign currency either as store of value, unit of account or medium of exchange. This phenomenon is usually known as *currency substitution* (CS) or simply *dollarisation*[1].

A question that has been raised is whether dollarisation exhibits a non-reversible behaviour. In fact, there are some reported cases in which the implementation of successful anti-inflation programmes was not enough to lessen sharply the demand for US dollars. This phenomenon has been identified in economies where high inflation rates persisted for long periods of time (evidence of dollarisation hysteresis was presented by Guidotti and Rodriguez, 1992, for Bolivia, Mexico, Peru and Uruguay, by Kamin and Ericsson, 1993, for Argentina, by Clements and Schwartz, 1993, for Bolivia, by Mueller, 1994, for Lebanon, by Mongardini and Mueller, 1999, for the Kyrgyz Republic, by Reding and Morales, 1999, for Bolivia). Yet in countries with moderate inflation rates or in countries that exhibited short-lived high inflation rates, dollarisation has been considered easier to reverse (reversibility patterns were identified in Egypt, Yemen and Chile, by Mueller, 1994, and in some transition economies, by Sahay and Végh, 1996).

To explain the different patterns of dollarisation identified in the empirical literature, some authors developed explanations based on the existence of dollarisation costs. Guidotti and Rodriguez (1992), following Dornbusch, Sturzenegger and Wolf (1990) and Dornbusch and

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[1] In this paper, we follow Savastano (1996), not distinguishing the terms "currency substitution" and "dollarisation". These two terms have been used in the literature with different and often contradictory meanings. For example, Calvo and Végh (1992) reserve the term "currency substitution" to the case in which the foreign currency drives out the domestic currency as means of payment and "dollarisation" to the case in which substitution affects only the store of value role of money. For other uses of the two terms in the literature, see Giovannini and Turtelboom (1994).

Reynoso (1989), argued that there are switching costs associated to the use of foreign currency, such as learning costs, that, once incurred by economic agents, imply persistence. Since learning requires time, in light of these explanations short-lived inflation experiences should not have permanent effects on dollarisation. More recently, Uribe (1998) and Reding and Morales (1999) stressed the role of network externalities in explaining non-reversibility. According to this interpretation, dollarisation costs decline with the aggregate level of dollarisation, giving rise to multiple equilibria and history dependence in the demand for foreign currency. Thus, a temporary high level of inflation can start a dollarisation process that is not necessarily reversed when inflation comes down. Still, the same conclusion holds, that dollarisation hysteresis is more likely to occur in economies that faced high inflation rates for long periods of time.

In this paper, we pick up three countries with different inflation experiences and dollarisation levels and we investigate whether dollarisation exhibits different reversibility patterns, as suggested by the literature. The sample includes a country that experienced hyperinflation (Bolivia), a high inflation country (Turkey) and a country that experienced moderate to low inflation (Indonesia). As in most empirical studies, the extent of CS is measured by the so-called "dollarisation ratio", defined as the ratio of foreign currency deposits (henceforth FCD) to domestic money. Using this proxy, we keep the results comparable with those in the previous literature. Contrary to the previous literature, no restriction reflecting a particular theory of money demand is imposed a priori. Adopting an agnostic functional form, we focus our attention on the significance of the term that captures the ratchet effect. By finding that in a country that always experienced moderate inflation (Indonesia), the dollarisation ratio displays as much hysteresis as in high inflation countries, the paper provides evidence against the view that this phenomenon is confined to highly dollarised economies or to economies that experienced high inflation rates for long periods of time.

The paper is organised as follows. In Section 2, we discuss the implications of using FCD as proxy for foreign money holdings. In Section 3 we discuss alternative specifications to test the CS hypothesis. In Section 4, we present the model to be estimated. In Section 5, we discuss the empirical results. Section 6 concludes.

2. Dollarisation ratios

A major limitation in the empirical analysis of dollarisation is that of measurability. Ideally, the measure of CS should include the total amount of foreign money balances held by the public, regardless the motive was means of payment or store of value. Unfortunately, such a measure is not easy to obtain, due to absence of data on foreign bank notes held by the public[2]. For this reason most empirical studies on CS have relied on imperfect measures, based on the residents' bank deposits denominated in foreign currency. The most popular proxy is the ratio of foreign currency deposits (FCD) to domestic money (the so-called *Dollarisation Ratio*).

In Figure 1, two different dollarisation ratios are depicted for each country. The narrow dollarisation ratio (XH) is defined as the ratio of foreign currency deposits held in the domestic banking system (FCDH) to M2. The broad dollarisation ratio (X) is defined as the sum of XH with the ratio of foreign currency deposits held in banks abroad (FCDA) to M2[3]. As shown in the figure, Bolivia exhibits the highest dollarisation ratios, with the total amount of FCD exceeding more than twice the stock of domestic money. This country had a hyperinflation in the mid-eighties, with the inflation rate reaching 11,750% in yearly terms (see Table 1). The subsequent stabilisation brought down the inflation rate back to one digit in 1993 and 1994, but did not succeeded in reversing dollarisation. The Turkish economy experienced moderate-to-high inflation rates along the sample period. In 1994, the inflation rate reached 100% and the broad dollarisation ratio (X) was already greater than one. In Indonesia, the inflation rate has remained stable and mostly with one digit. In

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[2] Some methods to estimate the amount of dollar bills in circulation were proposed by Melvin and De La Parra (1989) and Kamin and Ericsson (2003). The usefulness of these estimates is however doubtful for time series analysis, given the restrictive assumptions involved.

[3] All data in this paper are from IFS. We use *Cross-border bank deposits of non-banks by residence of depositor* (line 7xrd) for FCDA and *Foreign Currency Deposits* (line 25b) for deposits held in the domestic banking system. Sample sizes differ between countries because of data availability.

this country, both dollarisation ratios are significantly lower than in Turkey and Bolivia, but they exhibit an upward trend. Taking together, the figures suggest that dollarisation (i) is positively related to the past inflation experience and (ii) exhibits limited reversibility.

It is important to qualify the information content of the dollarisation ratios. Indeed, a rise in a dollarisation ratio reflects a shift towards foreign money holdings only if it corresponds to a decline in holdings of domestic money. If this rise is built through the reduction in the amount of other foreign money holdings, such a measure will wrongly suggest a rise in CS. This problem affects particularly the narrow measure of dollarisation (XH) when the institutional framework of FCDH is modified or expected to be modified.

The case of Bolivia illustrates this: between November 1982 and August 1985, FCD were banished from the domestic banking system and converted into domestic currency at a below market exchange rate. As we see in Figure 1, the narrow measure of dollarisation suggests absence of dollarisation in this period, which would be very unlikely in a country facing a hyperinflation[4]. In the same figure, we see that the broad measure of dollarisation is rising until 1985. A symmetric path is observed after the legalisation of FCDH, in August 1985: the narrow dollarisation ratio is likely to be overestimating the dollarisation change, because part of the adjustment was made through the repatriation of FCD held abroad.

Clearly, the broad dollarisation ratio (X) is less sensitive to modifications in the institutional framework of FCDH or to changes in the public expectations regarding its "confiscation risk" than the narrow dollarisation ration (XH). For this reason, only the broad measure of dollarisation will be

[4] In Mexico, Peru and Argentina, FCDH also declined in periods of rising inflation, either because of changes in their legal status or due to the erosion of the public confidence in their legal status (Rogers, 1992, Kamin and Ericsson, 1993, Savastano, 1996).

used in the estimation exercise below (other authors focusing on broad dollarisation ratios include Savastano, 1996, Mueller, 1994, Rogers, 1992, Ratti and Jeong, 1992, Miles, 1978).

It should be noted that the inclusion of FCDA in the dollarisation proxy is not free of criticism. If this component is not liquid enough to be held for transaction purposes, the estimation results will capture both "store of value substitution" and "means of payment substitution". However, such limitation is also hated by the narrow measure of dollarisation. Either using FCDH or FCDA, we are bound to a broad definition of currency substitution, which does not distinguish the different roles of money. This definition accords to Cuddington (1989) - to whom CS is a shift from domestic money balances to foreign money balances, regardless the location - but disagrees with Kamin and Ericsson (1993) - to whom CS refers exclusively to the shift towards foreign money balances held at home.

In general, the role of FCD depends on the development of the domestic capital market. If bonds denominated in foreign currencies are not easily available, interest bearing FCD may play the role of the "missing bond" (Sahay and Végh,1996). If, instead, bonds denominated in foreign currency are available, FCD will be a dominated but more liquid application than bonds, having a role comparable to that of domestic quasi money. As shown below, the estimation results with this particular sample are favourable to the second case.

3. Empirical models of CS

The CS hypothesis has been tested under a variety of specifications. In this section, we discuss the determinants of the relative money demand in light of the two main models of CS, which are the Portfolio Balance Model (Cuddington, 1983, Branson and Henderson, 1985) and the Liquidity Services Model (Miles, 1978, Thomas, 1985). We also refer to the specification proposed by El-Erian (1988) which has been increasingly used in the empirical analysis of currency substitution. In the following discussion, we assume that individual wealth is composed by money and bonds denominated in two different currencies and that foreign residents do not hold domestic money.

The Liquidity Services Model (LSM) motivates the use of money by its means of payment role. According to this theory, money is held despite being dominated by interest bearing bonds, because it facilitates transactions[5]. The individual holdings of domestic and foreign money are optimally chosen, given their "moneyness" and interest foregone. The first order conditions of the individual optimisation problem lead to the equality between the marginal rate of substitution in the production of liquidity services and the ratio of user costs. When the production function of liquidity services is a CES, the relative money demand takes the form[6]:

$$\frac{eF}{M} = \left[\frac{\delta_2}{\delta_1} \left(\frac{i}{i - \hat{e}_{+1}} \right) \right]^{\sigma}, \tag{1}$$

[5] Liquidity services have been specified either as reducing transaction costs or as an argument of the consumer's utility function. Since most utility functions that have been specified for empirical purposes are weakly separable, these two approaches are in general functionally equivalent. For a stochastic model in which money reduces transactions costs, see Thomas (1985). For a stochastic model in which money enters in the utility function, see Smith (1995).

[6] The functional form obtained by Miles (1978) is slightly different from (1) (as noted by Smith, 1995, overestimating the elasticity of CS) because the author assumed that the principal is part of the user cost of money. This is not, however, a common assumption in the money-demand literature (see, for example, Barnett et al., 1992).

where M and F denote the domestic and foreign nominal money holdings, e is the nominal exchange rate, i is the nominal interest rates on domestic bonds, $\hat{e}_{+1}^{E}$ is the expected exchange rate depreciation, δ_1 and δ_2 are the weights in the CES production function and σ is the elasticity of substitution between the two currencies. Thus, the ratio of holdings of foreign to domestic money depends negatively on the expected exchange rate depreciation[7]. In many applications, this variable has been replaced by the (observable) interest rate differential, using the uncovered interest rate parity condition.

A question that has been raised is whether the LSM provides a reasonable description of the money demand in a context in which asset markets are not complete. In fact, the rationale for separability between portfolio and CS decisions is the assumption that domestic and foreign bonds can be used to hedge the currency risk (Thomas, 1985). If bonds denominated in foreign currency are not easily available, as is likely to be the case in most developing countries, other variables apart from the relative user cost may be influencing the money demand. For this reason, Cuddington (1989) defended that in developing countries, where asset market are illiquid, one should follow instead the Portfolio Balance Model.

In light of the Portfolio Balance Model (PMB), money is a simple asset, without any particular feature that makes it distinguishable from other assets. Postulating gross substitutability between all assets, builders of this approach have employed money demand functions that depend negatively on the yield of each alternative asset. Using a semi-log specification and assuming homogeneity of degree zero in all returns, this leads to resident's demands for domestic and foreign currency of the form (see Cuddington, 1983):

[7] Note that the domestic interest rate appears in (1) because the domestic bond is taken as benchmark.

$$\log\left(\frac{M}{P}\right) = f_0 + f_1 i + f_2 (i^* + \hat{e}^E) + f_3 \hat{e}^E + f_4 w \qquad (2)$$

and

$$\log\left(\frac{eF}{P}\right) = g_0 + g_1 i + g_2 (i^* + \hat{e}^E) + g_3 \hat{e}^E + g_4 w, \qquad (3)$$

where P is the domestic price level, i^* is the nominal interest rates on foreign bonds, w is the individual real wealth and all the remaining variables are defined as before. In (2) and (3), all coefficients but g_3, g_4 and f_4 are expected to be negative. According to this theory, the coefficient f_2 captures capital mobility and f_3 measures the degree of CS.

Most empirical applications of the PBM have estimated individual money demand functions like (2) and (3)[8]. When however the dependent variable is a dollarisation ratio, the equation to be estimated becomes:

$$\log\left(\frac{eF}{M}\right) = h_0 + h_1 i + h_2 i^* + h_3 \hat{e}^E_{+1} + h_4 w, \qquad (4)$$

with $h_k = g_k - f_k$, for *k=0,1,2,4* and $h_3 = g_2 + g_3 - f_2 - f_3$. Now, the signs of some coefficients are ambiguous. For example, if bonds are closer substitutes to money balances of the same currency than to money balances of the alternative currency, then $h_1 > 0$ and $h_2 < 0$. But the opposite case is also possible.

In the more restricted case in which bonds are not available at all, the relative money demand becomes a function of the expected exchange rate depreciation alone (see Rojas-Suarez, 1992).

Dollarisation ratios have also been used in estimation under the "agnostic" model introduced by El-Erian (1994) and followed, among others, by Clements and Schwarts (1993), Mueller (1994) and Mongardin and Mueller (1999)[9]:

$$\ln\left(\frac{eF}{M+eF}\right)=\gamma_0+\gamma_1\hat{e}_{+1}+\gamma_2\left(i^F-i^D\right)+\gamma_3\phi\,. \qquad (5)$$

In (5), i^D and i^F denote the domestic and foreign money own yields. The term ϕ is a dummy variable in El-Erian (1988), a time trend in Clements and Schwartz (1993) and a ratchet variable in Mueller (1994) and Mongardin and Mueller (1999). Abstracting from this term and from the differences in the dependent variable, this model differs from (1) and (4) by specifying money own yields. However, by restricting the analysis to the case in which domestic and foreign bonds play no role, this model obviously fails to be agnostic.

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[8] Fore example, Kamin and Ericsson (1993) focused on the demand for domestic money. This approach has the advantage of not facing data restrictions, but it has the inconvenient of not being able to disentangle whether money demand hysteresis is due to dollarisation or financial innovation.

[9] Actually, the dependent variable in this literature is not always the same. FCD are measured as a proportion of broad money in El-Erian (1988) and Clements and Schwarts (1993), and as proportion of total deposits in Mueller (1994) and Mongardin and Mueller (1999).

4. Model and data used

The aim of this paper is not to test among different theories of CS, but rather to investigate the significance of ratchet effects in dollarisation. In order not to restrict the analysis to a particular model, the following structural relation is assumed[10]:

$$x_t = \beta_0 + \beta_1 i + \beta_2 i_t^* + \beta_3 \hat{e}_{t+1}^E + \beta_4 i_t^D + \beta_{56} \phi_t^V + v_t \,. \qquad (6)$$

In (6), x is the log of the broad dollarisation ratio and ϕ^V is a ratchet variable, given by the past-peak value of the pre-determined variable *V*. We test alternative ratchet variables, namely $\phi_t^V = \underset{i}{Max}\{V_i\}$ with $i = 1,2,...,t$ and $V = \{\hat{e}, i, x\}$. The remaining variables are defined as before.

It should be noted that, by omitting the FCD own rate from (6), the risk exist of its influence to be captured by the US dollar money market rate, eventually biasing its sign into a positive value. However, this may not be too inconvenient. As mentioned above, El-Erian (1988), Mueller (1994) and Mongardini and Mueller (1999) interpreted the US dollar money market rate as the FCD own rate, rather than as an opportunity cost. However, none of these authors allowed the estimation results to confirm this, because they forced the dollar money market rate to enter an interest rate differential, without checking previously the sign of its individual coefficient (see equation 5). Since in (6) we consider all interest rates as separate regressors, our results will shed some light as to the exact role of the US dollar money market rate when CS is measured by FCD.

All data in this paper are from the IFS. The sample period is determined by data availability. We use the domestic money market rate for *i*, the US dollar money market rate (3-months, London

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[10] Adding a scale variable (industrial production) to specification (6) did not in general improve the fit. Estimation results including this variable are not displayed so as to save space, but are available upon request.

offer) for $i^{\$}$ and the time deposits rate for i^D. In Indonesia we interpolated two missing observations of the domestic interest rates. Since data on i is not available for Bolivia, we replaced this variable by the current inflation rate. In case of Bolivia we also included a dummy for the period 1982:4-1985:3, during which FCD were banished from the domestic banking system (D1)[11]. The expected exchange rate depreciation is instrumented with current and past values of a pool of variables, according to the rational expectation hypothesis. The instrument set includes all interest rates, the inflation rate, the rate of money creation, the exchange rate depreciation, the real exchange rate and industrial production.

The order of integration of each variable was investigated running both unit roots and stationarity tests. In Table 2 we depict the conclusions regarding the likely order of integration (see Appendix 1 for test details). In general, all variables are stationary in differences. In levels, however, there are different cases: the USD money market rate, the dollarisation proxy in Turkey, the time deposits rate in Indonesia are likely to be non-stationary. The other money market interest rates, the inflation rate in Bolivia, the depreciation rates and the time deposits rate in Turkey are likely to be stationary. The dollarisation proxies in Indonesia and Bolivia and the time deposits rate in Bolivia do not have a clear order of integration.

As for the dynamic specification, we adopt a simple partial adjustment model. The equation to be estimated is:

$$x_t = A' Z_t + (1 - \lambda) x_{t-1} + u_t , \tag{7}$$

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[11] A second dummy was also tested for Bolivia, allowing for a structural break after the introduction of dollar checking accounts in 1987, but no significance was found. This result contrasts to that in Clements and Schwartz (1993), but these authors relied on the narrow measure of dollarisation which, according to the discussion in Section 2, is more sensitive to changes in the institutional framework. In Turkey, a capital account liberalisation in 1987 also constitutes a major change, but it cannot be controlled for in our sample.

where $\alpha_i = \lambda\beta_i \quad \forall\alpha_i \in A$, $u_t = \lambda v_t$ and is Z the vector of explanatory variables in (6). The advantage of this specification is that it can be estimated by conventional methods without requiring stationarity of the dependent variable. When the dependent variable is integrated of order one, co-integration is tested directly by the sign and significance of the adjustment coefficient, λ (Kremers et al, 1992). Since in that case λ has a non-standard distribution, a *t*-value of at least 3 is required to reject the null of no co-integration (Pesaran, Shin and Smith, 1996).

5. Estimation results

For each country, the following equations were estimated:

(i) The full unrestricted model with the exchange rate ratchet (ϕ^e).

(ii) Same as (i), but imposing null restrictions in selected non-significant coefficients.

(iii) Same as (ii) but using the interest rate ratchet (ϕ^i).

(iv) Same as (ii) but using the past-peak value of the dependent variable (ϕ^x).

(v) Same as (ii) but without any ratchet variable.

Whenever the expected exchange rate depreciation is among the explanatory variables the two-stage-least-squares procedure (2SLS) is used. The remaining equations are estimated by OLS.

The estimation results are depicted in Table 3. In general, the signs of the significant coefficients are as expected and the correlation coefficients are high, suggesting overall significance. Sargan validity tests confirm independence between the structural error term and the instruments whenever 2SLS is used. Two equations for Turkey, however, reveal major econometric problems. These are equation (iv), in which the Breusch-Godfrey LM test suggests residuals auto correlation, and equation (v), in which no co-integration is obtained. The correspondent estimates are only displayed for comparative purposes.

As shown in Table 3, in the case of Turkey co-integration is obtained only when the ratchet effect is accounted for. In Bolivia and Indonesia, where the order of integration of the dependent variable is not clear, high t-tests for λ are always obtained, even when no ratchet variable is specified. In general, ratchet variables are significant, confirming hysteresis in dollarisation. The only one with low significance is the interest rate ratchet in Indonesia (equation iii), probably because in this country the highest interest rate peak occurred too early in the sample (1984:3).

The evidence concerning the other explanatory variables is not uniform across countries. In Turkey, the money market interest rate is not significant. In Indonesia, this variable gets a positive and significant sign. In Bolivia we used the current inflation rate instead of the money market interest rate, obtaining a positive and significant coefficient in all equations[12].

The expected depreciation term is always significant and gets a positive sign in Turkey, confirming previous results (see, for example, Akçay et al., 1997). In the other two countries, this variable is not significant[13].

The interest rate on time deposits may affect the dollarisation ratio both through a direct effect on M2 and through a substitutability effect with FCD. Hence, β_4 is expected to be negative. In Table 3, we see that the time deposits rate is non-significant in Bolivia and in Indonesia. In Turkey it gets a negative and significant coefficient[14].

The US money market rate gets a negative sign across equations, and has low significance only in Bolivia. Taking into account that the FCD own rates are omitted from estimation, this is a rather strong result supporting the role of this variable as opportunity cost. The evidence in Figure 2, that deviations from uncovered interest rate parity are stationary in the three countries, also points to

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[12] Probably, this variable captures a Cagan effect in the demand for money, most likely with hyperinflation. Supporting this possibility is the fact that in the other two countries the domestic inflation rate does not improve the fit (results available upon request).

[13] The significance levels of the expected depreciation terms are not related to the performance of the forecasting models used (results for the exchange rate forecasting model are available upon request).

[14] Eventually, the different performance of this variable across countries is related to the different compositions of M2 observed in the data. Indeed, while in Turkey the ratio M1/M2 is small and relatively stable, in Bolivia this ratio is very high (it floats at around 70%) and in Indonesia it declines continuously along the sample period (from 75% to less than 30%).

the idea that the respective money markets are far from isolated from the rest of the world[15]. This evidence has two implications. On one hand, it suggests that FCD have a role comparable to that of domestic quasi money. On the other hand, it challenges the approach of El-Erian (1988) and others, of treating the US dollar money market interest rate as the FCD own rate inside an interest rate differential, without checking previously the signs of the individual coefficients (see equation 5)[16].

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[15] As shown in Figure 2, in Bolivia, the deviation becomes systematically negative after 1986. Although this can be attributed to the fact that we are using for this country the time deposits rate instead of the money market interest rate, this path is also consistent with the emergence of a positive risk premium in peso securities following the hyperinflation episode.

[16] The experiment depicted in Appendix 2, using data for Bolivia illustrates this. Two regressions are computed to explain the narrow measure of dollarisation, one unrestricted and the other imposing an interest rate differential as explanatory. The results show that, although negative coefficients on individual interest rates are obtained in the first regression, a positive and significant coefficient on the interest rate differential is obtained in the second. A possible explanation for such contradiction is that, in the presence of high inflation rates, the time deposits rate is so noisy that the foreign interest rate becomes negligible inside the interest rate differential.

6. Conclusions

This paper investigates whether dollarisation exhibits different reversibility patterns in countries with different degrees of dollarisation and inflation history. To this aim, we test the significance of ratchet effects in an equation exploring the determinants of dollarisation in Bolivia, Turkey and Indonesia. Contrary to other studies, no restriction reflecting a particular theory of money demand is imposed *a priori*. In general, the estimation results support the approach followed. The results do not reveal a common pattern across the three economies, as far as the significance of the various opportunity costs of money is concerned. Ratchet effects, however, are significant in the three countries, regardless their inflation experience and degree of dollarisation. This evidence contradicts the view according to which this phenomenon is confined to highly dollarised economies or to economies that experimented high inflation rates for long periods of time.

Although the evidence provided in this paper might be seen as contradicting the theoretical literature that emphasises the role of dollarisation costs (Guidotti and Rodriguez, 1992, Uribe, 1998, Reding and Morales, 1999), such interpretation should not be overstated. True, we find that dollarisation hysteresis occurs in the three countries at hand, regardless their dollarisation level and past inflation experience. Because of data limitations, however, the exercise is bound to a broad concept of dollarisation that may not be suitable to infer about means of payment substitution - arguably the concept that is more consistent with the theories mentioned above. Still the paper challenges this literature, by contradicting the supporting evidence. Indeed, although the models that have been developed to explain dollarisation hysteresis emphasise the means of payment role money, they all refer to stylised facts which nonetheless are based on the observation of FCD. By showing that FCD exhibit hysteresis in the three countries, regardless their inflation experience and dollarisation level, the results contradict the stylised facts which the literature tries to explain.

Explanations for the phenomenon of dollarisation hysteresis not depending on dollarisation costs or on the past inflation experience include Freitas (2002) and Sahay and Végh (1996). Freitas (2002) argues that hysteresis and non-linearities in the relationship between money and inflation are a natural implication of having two currencies coexisting in the same commodity domain. According to this theory, if the domestic and foreign currencies are perfect substitutes as means of payment, the demand for foreign currency will not necessarily decline after stabilisation. If, however, reversibility takes place (partial or total), the demand for foreign currency will decline smoothly, contrasting with the once-and-for-all change in the inflation rate. In a different reasoning, Sahay and Végh (1996) pointed that in developing countries, where bonds denominated in foreign currency are not easily available, interest-bearing FCD may be an attractive application for savings, playing mostly a store of value role. Hence, they will not necessarily decline with low inflation and nominal interest rates.

By pointing to some substitutability between FCD and foreign bonds, the evidence obtained in this paper does not give support to the idea that dollarisation hysteresis affects only countries where bonds denominated in foreign currency are not available. Still, since FCD are the fraction of foreign money holdings more likely to be held for store of value reasons, the argument that they do not necessarily decline after stabilisation because of portfolio reasons cannot be ruled out. In order to get a clearer picture in respect to the theory that better explains dollarisation hysteresis, one should be able to verify whether this phenomenon affects in the same manner foreign currency deposits and dollar bank notes held by the public. Unfortunately, data limitations make such exercise very difficult to implement.

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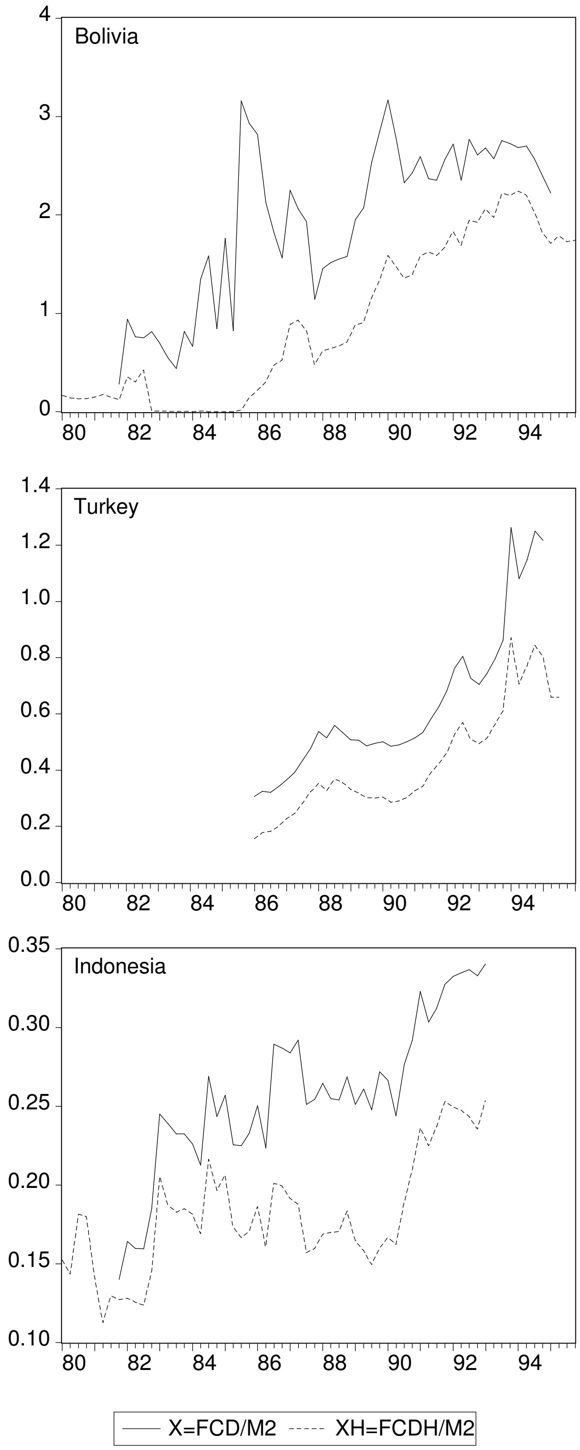
Figure 1: Dollarisation Ratios in Bolivia, Turkey and Indonesia



Source: Own calculations, using IFS data.

Table 1: Annual Inflation Rates in Bolivia, Turkey and Indonesia

	Bolivia	**Turkey**	**Indonesia**
1981	32.1	36.6	12.2
1982	123.5	30.8	9.5
1983	275.6	31.4	11.8
1984	1281.4	48.4	10.5
1985	11749.6	45.0	4.7
1986	276.3	34.6	5.8
1987	14.6	38.8	9.3
1988	16.0	73.7	8.0
1989	15.2	63.3	6.4
1990	17.1	60.3	7.8
1991	21.4	66.0	9.4
1992	12.1	70.1	7.5
1993	8.5	66.1	9.7
1994	7.9	106.3	8.5
1995	10.2	93.6	9.4

Source: IMF, International Financial Statistics.

Table 2 - Summary of the unit root tests

	Bolivia	Turkey	Indonesia
x	*?* [3/]	*I(1)*	*?* [2/]
i [1/]	*I(0)*	*I(0)*	*I(0)*
$\hat{e}$	*I(0)*	*I(0)*	*I(0)*
i^D	*?* [3/]	*I(0)*	*I(1)*
	USD money market rate		
i^*	*I(1)*		

Notes: Test details are in Appendix.
1/ In Bolivia this regressor is the current inflation rate.
2/ ADF and PP reject unit roots, but KPSS rejects stationarity.
3/ ADF not rejected and KPSS rejects stationarity, but PP rejects unit roots.

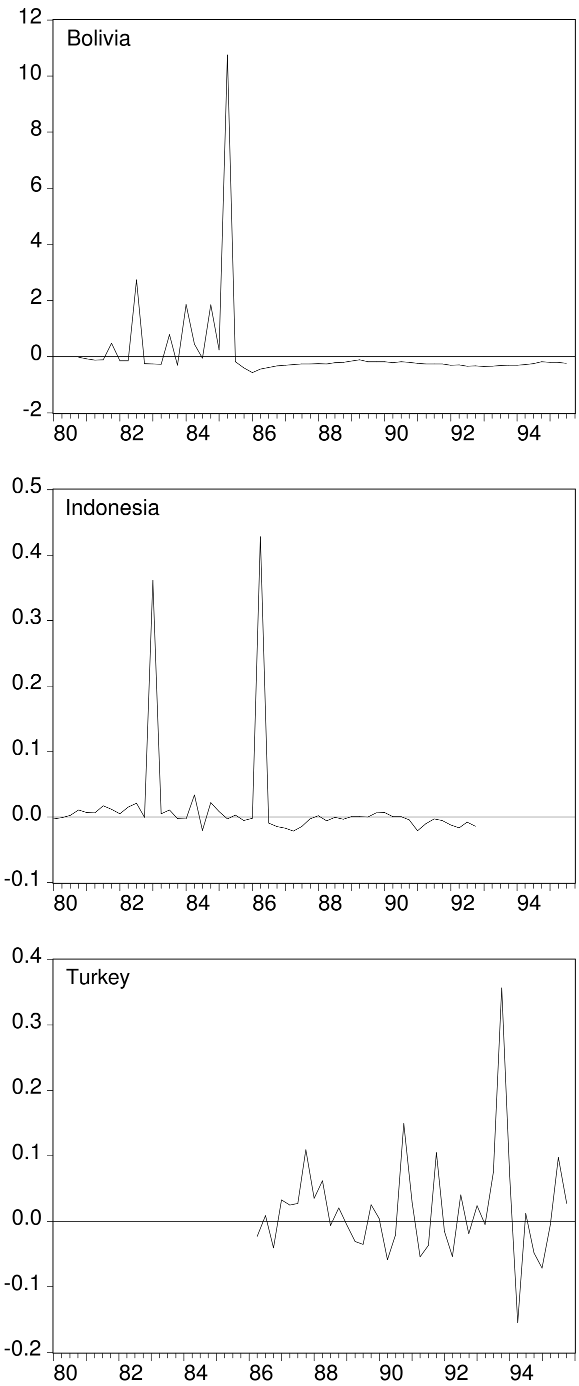
Table 3- Estimates of the determinants of the broad dollarisation in Bolivia, Turkey and Indonesia

	Indonesia					**Turkey**					**Bolivia**			
Sample	1982:1-1992:4					1986:4-1995:1					1982:1-1995:1			
Equation	(i)	(ii)	(iii)	(iv)	(v)	(i)	(ii)	(iii)	(iv)	(v)	(i)	(ii)	(iii)	(iv)
Method	2SLS	OLS	OLS	OLS	OLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	OLS	OLS	OLS
Constant	-0.97	-0.93	-0.74	-0.37	-0.56	-0.16	-	-	0.17	0.26	0.21	-	-	0.81
	-4.2	*-4.5*	*-4.3*	*-4.7*	*-6.1*	*-0.7*			*1.3*	*1.5*	*0.6*			*3.1*
λ - Adjustment coefficient	0.72	0.70	0.54	0.94	0.50	0.27	0.20	0.22	0.84	-0.02	0.65	0.63	0.72	0.55
	5.0	*5.1*	*5.4*	*8.5*	*5.3*	*2.3*	*3.9*	*4.2*	*4.6*	*-0.3*	*5.8*	*7.1*	*7.2*	*6.0*
i - Money market interest rate (1)	5.98	6.12	5.73	4.75	5.78	-0.09	-	0.19	0.31	0.89	0.18	0.18	0.13	0.19
	4.2	*4.5*	*4.1*	*4.4*	*4.1*	*-0.2*		*0.5*	*0.9*	*2.1*	*4.2*	*4.2*	*3.0*	*4.3*
i^* - USD money market rate	-16.58	-16.66	-14.40	-10.17	-15.00	-5.49	-6.67	-8.38	-5.94	-5.03	-0.38	-	-	-0.61
	-4.8	*-5.2*	*-4.5*	*-3.9*	*-4.7*	*-1.8*	*-2.5*	*-3.0*	*-2.3*	*-1.4*	*-1.1*			*-1.9*
iD -Time deposits rate	0.79	-	-	-	-	-1.22	-1.84	-2.11	-1.32	-2.39	0.06	-	-	-
	0.6					*-1.1*	*-4.2*	*-4.3*	*-1.6*	*-2.2*	*0.7*			
$\hat{e}$ - Expected depreciation	-0.03	-	-	-	-	0.49	0.40	0.33	0.03	0.28	-0.02	0.04	-	-
	-0.1					*2.1*	*2.9*	*2.2*	*0.2*	*1.2*	*-0.8*	*6.2*		
ϕ - ratchet effect	0.26	0.29	1.72	0.67	-	0.91	0.80	1.41	0.84	-	0.03	-	0.50	-
	1.5	*2.0*	*1.2*	*5.3*		*3.2*	*4.6*	*4.2*	*5.0*		*2.1*		*6.3*	
D1	-	-	-	-	-	-	-	-	-	-	-0.40	-0.44	-0.36	-0.63
											-2.9	*-5.0*	*-4.2*	*-5.6*
Adjusted R-squared	0.87	0.88	0.87	0.92	0.87	0.97	0.97	0.97	0.98	0.96	0.82	0.81	0.81	0.78
S.E. of regression	0.066	0.064	0.066	0.051	0.066	0.063	0.059	0.058	0.051	0.070	0.22	0.23	0.23	0.24
Sargan (d.f)	0.02(1)	-	-	-	-	0.1(2)	1.8(4)	2.1(3)	0.7(2)	0.9(1)	2.1(3)	-	-	-
LM(4)	0.3	0.2	0.6	1.8	1.3	0.4	1.4	1.5	20.4	4.2	2.0	3.9	6.5	0.8
Number of observations	44	44	44	44	44	34	34	34	34	34	53	53	53	53

Notes: t-ratios in italic. LM(4) denotes the Breusch-Godfrey test for serial correlation of order 4.

(1) In Bolivia this regressor is the current inflation rate.

Figure 2: Deviations from Uncovered Interest Rate Parity



Notes: Deviations (δ) are computed using the 'one quarter ahead' depreciation rate and quarterly adjusted money market rates, according to $\delta = (1+\hat{e}_{+1})(1+i^{*})/(1+i) - 1$. In Bolivia the time deposits rate was used instead of the domestic money market rate. A positive deviation means that the forward rate is underestimating the future exchange rate.
Source: Same as Figure 1.

Appendix 1

Table A1 – Unit root tests

Turkey	Level	ADF	PP	KPSS
x – Dollarisation ratio	1	ADF(t,0)=6.3***	z_μ =6.3***	η_μ =0.092
	0	ADF(c,0)=0.4	z_t =-2.3	η_t =0.141*
i - Money Market rate	1	ADF(t,4)=3.8**	z_μ =11.8***	η_μ =0.046
	0	ADF(t,3)=3.9**	z_t =3.4*	η_t =0.045
$\hat{e}$ *- Depreciation rate*	1	ADF(t,0)=8.5***	z_μ =13.0***	η_μ =0.042
	0	ADF(t,1)=5.2***	z_t =5.0***	η_μ =0.092
i^D *- Deposit rate*	1	ADF(t,1)=7.7***	z_μ =10.1***	η_μ =0.084
	0	ADF(t,0)=3.7**	z_t =3.8**	η_t =0.097
Indonesia				
x – Dollarisation ratio	1	ADF(t,0)=8.6***	z_μ =9.4***	η_μ =0.193
	0	ADF(t,0)=3.9**	z_t =3.8**	η_t =0.127*
i - Money Market rate	1	ADF(t,0)=5.9***	z_μ =11.6***	η_μ =0.031
	0	ADF(t,0)=4.8***	z_μ =4.7***	η_μ =0.141
$\hat{e}$ *- Depreciation rate*	1	ADF(t,3)=6.4***	-	η_μ =0.044
	0	ADF(t,0)=8.1***	z_μ =8.1***	η_μ =0.194
i^D *- Deposit rate*	1	ADF(t,2)=3.5*	z_μ =4.6**	η_t =0.061
	0	ADF(c,1)=1.7	z_μ =1.5	η_t =0.146*
Bolivia				
x – Dollarisation ratio	1	ADF(t,3)=5.5***	z_μ =15.9***	η_μ =0.226
	0	ADF(c,2)=1.8	z_t =5.1***	η_t =0.202***
i - Money Market rate	1	ADF(t,0)=14.0***	z_μ =13.5***	η_μ =0.055
	0	ADF(t,3)=3.2*	z_μ =4.7***	η_t =0.099
$\hat{e}$ *- Depreciation rate*	1	ADF(t,3)=6.2***	z_μ =36.6***	η_μ =0.051
	0	ADF(t,0)=7.6***	z_μ =7.7***	η_μ =0.265
i^D *- Deposit rate*	1	ADF(t,3)=7.3***	z_μ =8.4***	η_μ =0.138
	0	ADF(c,4)=1.1	z_t =3.3*	η_t =0.152*
US dollar 3-months money market rate				
	1	ADF(t,0)=7.4***	z_μ =7.5***	η_μ =0.114
	0	ADF(c,0)=2.0	z_t =2.8	η_t =0.126*

Notes: (*), (**) and (***):significance at 90%, 95% and 99%, respectively. ADF(t,k), ADF(c,k): ADF with *k* lags, with and without time trend, respectively. z_t , z_μ: Phillips-Perron with and without time trend. η_t , η_μ :KPSS test with and without time trend. The number of lags is set according to the Schwert criterion in the KPSS test and the Newey-West criterion in the Phillips-Perron test.

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Appendix 2

Table A2 – Testing an interest rate differential: Bolivia

Country	**Bolivia**	
Sample	1980:4-1995:4	
Dependent	xh	xh
Method	OLS	OLS
Constant	1.34	-0.78
	2.9	*-5.1*
λ - Adjustment coefficient	0.67	0.53
	12.1	*9.4*
i - Inflation rate	0.72	0.64
	3.7	*2.8*
(i*-iD) - Interest rate differential		0.51
		2.6
i* - USD money market rate	-1.40	
	-3.3	
iD -Time deposits rate	-0.80	
	-4.5	
ϕ - ratchet effect	0.05	0.09
	3.4	*5.6*
D1	-3.47	-2.86
	-14.7	*-12.0*
Adjusted R-squared	0.98	0.97
S.E. of regression	0.29	0.35
LM(4)	4.3	13.4
Number of observations	58	58

Notes: t-ratios in italic.