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Abstract
This paper examines the effects of inflation and currency substitution volatility on the average rates of inflation and currency substitution for twelve emerging market economies. Using a bivariate GARCH-in-Mean model, which accommodates for asymmetric and spillover effects of inflation and currency substitution innovations on their volatilities, we find that for the majority of the countries in the sample the variability of inflation exerts a positive influence on both the average rates of inflation and currency substitution. Similarly, higher uncertainty in currency substitution displays enhancing effects on inflation and currency substitution. These results indicate an alternative avenue that stresses the importance of currency substitution for the conduct of monetary policy in terms of price stability, and provide an additional explanation to the phenomenon of dollarization hysteresis.

JEL Classification: C32; E31; E52; F31

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1. Introduction

Currency substitution (CS), the replacement of domestic with foreign currencies, has been a salient feature of several developing countries in recent years. Such a phenomenon has been particularly pervasive in countries that have experienced severe inflationary periods and/or uncertainty about domestic macroeconomic policies.[1] The widespread use of foreign currencies, especially since the early 1990s in many emerging economies, has triggered a lot of research in pursuit of the determinants of CS (Agénor and Khan 1996, De Nicoló et al. 2003), of the economic elements that are affected by its development (Barnett and Ho 1996, Akçay et al. 1997, Berg and Borenzstein 2000, Lange and Sauer 2005), and of explanations of its inertia (Clements and Schwartz 1993, Kamin and Ericsson 1993, Tandon and Wang 1999, 2003, Oomes 2003, de Freitas 2004).[2] In addition, a recently developed line of work examined the extent to which CS has been important for the conduct of effective monetary policy (see Kamin et al. 1998, Reinhart et al. 2003, Mishkin 2004).

A common characteristic of most of these studies is their focus on the relationship between the rate of inflation and the degree of CS. On the one hand, they emphasize the direct effect the average rate of inflation has on CS as private agents substitute out of the domestic currency to hedge against the erosion of its value, which intensifies CS. On the other hand, they stress the feedback effect of the degree of CS on the inflation rate as the base of the inflation tax shrinks and a financially constrained government monetizes its budget deficit, which, in turn, leads to higher inflation (Chang 1994, Sturzenegger 1997, Bahmani and Domac 2003, Levy-Yeyati 2004). The interplay between inflation and CS is also present in studies that examine the inertia of CS. Theoretical explanations of this phenomenon have mainly focused on the transactions costs of inflation involved in switching between two currencies (Guidotti and Rodriguez 1992, Uribe 1997, and de Freitas 2004), while empirically CS hysteresis has been accounted for by means of

[1] For definitions of the different types of CS and a survey of the literature, see Giovannini and Turtelboom (1994), Calvo and Végh (1992), and more recently Reinhart et al. (2003).

[2] Also coined *dollarization hysteresis*, the inertia of CS describes the persistence of demand for foreign currencies even long after the elapse of inflationary episodes or reductions in the exchange rate depreciation of the national currency.

"ratchet effects".[3] In these studies, the variable that has most commonly been used to capture the ratchet effects that lead to CS persistence is the rate of inflation (see, for example, Kamin and Ericsson 1993).

On the policy making side, CS has attracted a lot of attention with regard to its influence in the design and transmission of monetary policy. Although the earlier literature expressed the concern that CS, by raising the volatility of money demand, may impede the ability of the central bank to conduct monetary policy and reduce inflation, recent work by Reinhart et al. (2003), De Nicoló et al. (2003), and Havrylyshyn and Beddies (2003) find no empirical evidence in favour of this hypothesis. Accepting the notion that the primary goal of monetary policy is to attain and maintain a low rate of inflation, these studies provide evidence which supports the coexistence between high dollarization and low inflation. In this spirit, Masson et al. (2003) imply that in such low-inflation environment inflation targeting could possibly serve as a potential monetary policy framework.[4]

The above considerations, although illustrative of the interaction between CS and inflation, limit the analysis in the bi-directional causality of their first moments ignoring the possible impact of the second moments of inflation and CS on their average rates. This is an important consideration that could provide further insights for the determinants of inflation and CS and their potential spillover effects, especially in the process of explaining CS hysteresis and designing monetary policies suitable for high-inflation and emerging market economies.[5] To that extent, our approach utilizes a recently developed econometric technique (bivariate VARMA, GARCH-in-Mean model) by Grier et al. (2004) to study the impact of the uncertainty of inflation and CS on their average rates. The distinctive features of our analysis are as follows. First, we examine whether CS is affected by its own volatility, in addition to the volatility of inflation. Second, we depart

[3] Ratchet effects describe the asymmetric response of a dependent variable to changes in an explanatory variable. In this way, positive shocks in an explanatory variable out way the effects on the dependent variable from negative shocks of equal magnitude, and imply persistence. Ratchet effects are usually captured by the past peak value of key right-hand-side variables.

[4] Additional prerequisites include a degree of independence of monetary policy and proper policy instruments linked to inflation.

[5] Exceptions are the studies by Piontkovsky (2003) who shows that CS depends on inflation volatility, and Ize and Levy-Yeyati (2003) and Peiers and Wrase (1997) who illustrate that CS hysteresis is partially explained by the volatility of inflation.

from earlier studies of CS hysteresis and provide an additional explanation through the observed hysteresis in the volatilities of CS and inflation. Third, we study the effect of the volatility of CS on inflation and its subsequent implications for the effectiveness of monetary policy. Fourth, the generality of our methodology allows for asymmetric effects of the variability of inflation and CS innovations on their average rates and the spillover of volatility from one series to the other. As a result, our methodology allows us to assess the evidence in support of any given hypothesis using data from twelve emerging market economies.

Our findings suggest that for the majority of the countries in our sample, the variability of inflation has a positive effect both on the average rate of CS and the average rate of inflation. The first result points towards the existence of direct enhancing effects of inflation rate volatility on the rate of CS as implied by the minimum variance portfolio (MVP) allocation model developed by Ize and Levy-Yeyati (2003).[6] Moreover, the positive impact of inflation uncertainty on average inflation is in line with a model developed by Cukierman and Meltzer (1986), who show that the agent's uncertainty over the policymaker's objectives and the monetary growth process (and therefore inflation) gives the policymaker the incentive for surprise inflation to achieve an output stimulus. This reasoning is in contrast to Holland (1995) who claims that monetary authorities can diminish the negative effects of inflation uncertainty on output growth by decreasing inflation as uncertainty rises, therefore establishing a negative effect of inflation variability on inflation.

At the same time, the findings indicate that the uncertainty about the degree of CS exerts a positive influence on both the average levels of inflation and CS. These results show that the volatility of CS is an important determinant of the level of CS, not identified in the traditional literature, and of the inflation rate, highlighting the (second-order) adverse effects of CS on monetary policy. Exploration of the asymmetric effects of shocks indicates that, for most of the countries, positive CS innovations raise uncertainty about CS by more than negative innovations do. The results, however, for the innovations in the rate of inflation are mixed. Therefore, although both inflation and CS are

[6] The MVP selection approach dictates the level of dollarization by the relative difference of the volatilities of inflation and real exchange rate depreciation. For instance, an increase in the variability of inflation, for a given volatility in the rate of depreciation, increases dollarization.

characterized by own-variance asymmetry, the ratchet effects in the volatility of CS imply persistence. This result along with the positive effects of inflation and CS volatilities on CS imply an alternative explanation of hysteresis in CS.

These issues are of particular interest for the transition economies that are new members of the European Union (EU) or at stages of candidacy to join the EU, where CS is widespread, in order to examine the benefits from a subsequent monetary union membership. According to our findings, such benefits would, in part, be attained by giving up their monetary policy to the European Central Bank (ECB).[7] The adoption of the euro as their national currency will benefit them from decreasing rates of inflation, not only directly due to the decline in CS, but also indirectly through the reduction of its own variance and the volatility of inflation.

The rest of the paper is structured as follows. Section 2 describes the data and the econometric methodology we utilize. Section 3 presents the results and discusses their implications. Finally, section 4 concludes.

2. Data and econometric methodology

While there exists a substantial number of empirical studies on CS, a common challenge has been the measurement of foreign currency circulating in an economy. The most important limitations have to do with the lack of data on foreign currency notes in circulation in the domestic economy, foreign currency deposits of domestic residents abroad, and foreign currency deposits of foreign residents domestically.[8] Due to these restrictions, the majority of the studies resort to measure the level of CS with the foreign currency denominated deposits in the domestic banking system, coined in the literature as financial dollarization (see Clements and Schwartz 1993, Akçay et al. 1997, Baliño et al. 1999, Honohan and Shi 2002). One needs to keep in mind, however, that this measure represents the lower bound of CS and it abstains from issues of the time maturity of deposits.

[7] We refer to partial benefits of joining the EMU since a comprehensive analysis also requires an examination of these countries' business-cycle correlation with the Euro zone (see, for instance, Furceri and Karras 2006).

[8] Notable exceptions are studies that have approximated the amount of foreign exchange bills and coins in circulation for selected developing countries (see, for instance, Kamin and Ericsson 1993, and Oomes 2003). As argued by Savastano (1992, 1996), however, the usefulness of these estimations is questionable as they are based on very restrictive assumptions.

In light of these constraints, and to allow for comparison across studies, in what follows we measure CS as the ratio of foreign currency deposits (FCD) of private agents to a broad monetary aggregate (M2). Both monthly series, CS and inflation, are collected by the *International Financial Statistics* dataset of the IMF, when available, and additionally from the website of each country's Central Bank.[9] We measure inflation as the annualized monthly difference of the logarithm of the consumer price index (by multiplying the log difference by 1200).[10] Similarly, the series for CS is transformed to its annualized monthly difference (by multiplying the first-differenced series, ΔCS, by 1200). Summary statistics and diagnostic tests are presented in Table 1, Panel A and B respectively.

We investigate the stationarity properties of the series by employing a number of Augmented Dickey Fuller (ADF) tests (Dickey and Fuller 1979). The results of these tests indicate that we can treat the *change* in CS (ΔCS) and the inflation rate in each country as stationary processes.[11] Both series for all countries display significant amounts of kurtosis while inflation is positively skewed. The series for CS suggests mixed results as far as the sign of skewness is concerned while both series fail to satisfy the null hypothesis of the Jarque-Bera (1980) normality test. Finally, the Ljung-Box (1979) test for up to twelve lags serial correlation indicates the strong presence of serial dependence in the data. Similarly, a Ljung-Box test for serial correlation in the squared data provides evidence of conditional heteroscedasticity in the data (except for Bulgaria in the case of inflation). These tests suggest that second moment (nonlinear) dependencies are significant, supporting the use of a specification that captures the instability of the variances of inflation and CS.

Equation (1) presents the approach we use to model both the change in CS (s_t) and the rate of inflation (π_t) simultaneously. For this purpose, a VARMA-GARCH-in-Mean model is adopted by Grier et al. (2004) (also see Bredin and Fountas, 2005). This

[9] The countries in the sample are: Armenia, Bulgaria, the Czech Republic, Estonia, Georgia, Kyrgyz Republic, Latvia, Poland, Romania, the Russian Federation, Turkey, and Ukraine.
[10] This practice is standard in the time series literature (see, for example, Grier et al. 2004).
[11] For all countries, the original CS series failed the stationarity tests in levels. Since stationarity is a necessary condition of our econometric methodology we use the first difference of the series. In addition, the ADF tests for the ΔCS in Turkey and for the inflation rate in Bulgaria, the Czech Republic, Latvia and Turkey indicate that trend stationary processes are more appropriate. As shown later, to accommodate for these cases we include an additional term in our model that captures the trend of ΔCS and inflation.

specification simultaneously estimates equations for both change in CS and inflation, where it takes into account the uncertainty of each variable by including in their mean equations the conditional standard deviations of both series as explanatory variables. The standard Schwartz Bayesian information criterion is used to determine the appropriate lag of the VARMA process, taking also into account the GARCH-in-mean effects.

$$
\begin{aligned}
&Y_t = c + zt + \sum_{i=1}^{p} \Gamma_i Y_{t-i} + \Psi \sqrt{h_t} + \sum_{j=1}^{q} \Theta_j \varepsilon_{t-j} + \varepsilon_t \\
&\varepsilon_t \mid \Omega_t \sim (0, H_t) \\
&H_t = \begin{bmatrix} h_{s,t} & h_{s\pi,t} \\ h_{\pi s,t} & h_{\pi,t} \end{bmatrix}
\end{aligned}
\qquad (1)
$$

where

$$
Y_t = \begin{bmatrix} s_t \\ \pi_t \end{bmatrix};\ \varepsilon_t = \begin{bmatrix} \varepsilon_{\mathrm{s,t}} \\ \varepsilon_{\pi,\mathrm{t}} \end{bmatrix};\ \sqrt{\mathrm{h_t}} = \begin{bmatrix} \sqrt{h_{\mathrm{s,t}}} \\ \sqrt{h_{\pi,\mathrm{t}}} \end{bmatrix};\ \mathrm{c} = \begin{bmatrix} \mathrm{c_s} \\ \mathrm{c}_\pi \end{bmatrix};\ \mathrm{z} = \begin{bmatrix} z_s \\ z_\pi \end{bmatrix};\Gamma_i = \begin{bmatrix} \gamma_{11}^{(i)} & \gamma_{12}^{(i)} \\ \gamma_{21}^{(i)} & \gamma_{22}^{(i)} \end{bmatrix};
$$

$$
\Psi_{\mathrm{i}} = \begin{bmatrix} \psi_{11} & \psi_{12} \\ \psi_{21} & \psi_{22} \end{bmatrix};\ \Theta_j = \begin{bmatrix} \theta_{11}^{(j)} & \theta_{12}^{(j)} \\ \theta_{21}^{(j)} & \theta_{22}^{(j)} \end{bmatrix}
$$

Matrix *z* captures the trend of ΔCS and the rate of inflation wherever applicable (see footnote 11). For the cases where the trend term is unnecessary the corresponding coefficients of the *z* matrix are set equal to zero. The assumption $\varepsilon_t \mid \Omega_t \sim (0, H_t)$, where Ω_t represents the information set available at time *t*, is made only to motivate the use of quasi-maximum likelihood estimation of the parameters; non-normality is allowed for by using robust "sandwich" standard errors for the parameter estimates. In addition, the estimation of parameters is subject to the condition that H_t is positive definite for all values of ε_t in the sample. To model the conditional variance-covariance matrix, H_t, we adopt the asymmetric version of the BEKK model of Engle and Kroner (1995). By using this model we are able to capture any possible asymmetric effects of innovations of inflation and change in CS on the conditional variances and covariance of the two series. The asymmetry in the conditional variance process can be incorporated by defining good and bad news respectively. More specifically, if both inflation and change in CS are higher than expected (positive innovations), we consider that to be bad news. Therefore, with bad news the residuals of the two series will be positive, and we define $\xi_{s,t}$ and $\xi_{\pi,t}$ as $\max(\varepsilon_{k,t},0)$, where k = s, π.[12] The asymmetric conditional covariance can be expressed using:

$$H_t = K_0^{*'}K_0^* + A_{11}^{*'}\varepsilon_{t-1}\varepsilon_{t-1}^{'}A_{11}^* + D_{11}^{*'}\xi_{t-1}\xi_{t-1}^{'}D_{11}^* + B_{11}^{*'}H_{t-1}B_{11}^* \tag{2}$$

where

$$K_0^* = \begin{bmatrix} \kappa_{11}^* & \kappa_{12}^* \\ 0 & \kappa_{22}^* \end{bmatrix};\ \mathrm{A}_{11}^* = \begin{bmatrix} \alpha_{11}^* & \alpha_{12}^* \\ \alpha_{21}^* & \alpha_{22}^* \end{bmatrix};\ \mathrm{D}_{11}^* = \begin{bmatrix} \delta_{11}^* & \delta_{12}^* \\ \delta_{21}^* & \delta_{22}^* \end{bmatrix};\ \mathrm{B}_{11}^* = \begin{bmatrix} b_{11}^* & b_{12}^* \\ b_{21}^* & b_{22}^* \end{bmatrix};\ \xi_{\mathrm{t}} = \begin{bmatrix} \xi_{\mathrm{s,t}} \\ \xi_{\pi,\mathrm{t}} \end{bmatrix}$$

An important difference between this specification and other models (see, for instance, Grier and Perry 2000), by construction, is that it allows testing rather than imposing diagonality and symmetry restrictions. In this manner, we avoid potential specification errors and remove the unnecessary terms when appropriate. For example, if there were no asymmetry present, then the coefficient matrix D_{11}^* would not be statistically significant and equation (2) would be the symmetric BEKK model (Engle and Kroner, 1995).

3. Empirical Results

3.1. Specification Tests and Description of Results

The model described by equations (1) and (2) is jointly estimated with the Bollerslev and Wooldridge (1992) quasi-maximum likelihood method, which allows for

[12] We examine the validity of our definition of positive innovations in the two series by using the Engle and Ng (1993) test for asymmetry in volatility. Both series are found to exhibit positive sign and size bias supporting our definition of bad news and implying that their absence from the model may lead to a specification error. Exceptions are Georgia for which there is no significant size bias in neither of the two series, and Poland and Russia where only inflation and CS respectively are found to have a size bias.

robust standard errors.[13] A series of diagnostic tests were carried out for each country individually to establish the adequacy of the model specification. These tests examine the presence of conditional heteroscedasticity in the data and the existence of a non-diagonal covariance process. Both tests consider the significance of the elements of the matrices A^{*}_{11}, B^{*}_{11}, and D^{*}_{11}, where conditional heteroscedasticity requires the joint significance of all their elements while non-diagonal covariance specification calls for the joint significance of the off-diagonal elements. The results of Table A2 (in the Appendix) suggest the strong conditional heteroscedasticity and the significance of the off-diagonal elements in the covariance structure at the 1% level.

Table 2 presents the outcome of two additional tests with respect to coefficients ψ_{ij} and δ^{*}_{ij}, where the remainder of the analysis will focus upon. A test regarding the joint significance of the effect of uncertainty about inflation and change in CS on their average rates (Panel A), and a test about the presence of asymmetry in the covariance process (Panel B). The upper panel of Table 2 illustrates the joint significance of the elements of matrix Ψ implying that the exclusion of inflation and CS uncertainty as explanatory variables of their average rates would lead to a misspecified model for all twelve countries in our sample. Similarly, the lower panel indicates that the matrix of the asymmetric terms is jointly significant for most of the countries at the 1% level. Exceptions are Georgia, Poland, and the Russian Federation.[14] For the countries with significant asymmetry coefficients we can conclude that the covariance process is asymmetric while for the rest of them we can exclude the asymmetry terms, and hence set $D^{*}_{11} = 0$, to estimate a more parsimonious model.

Finally, we perform standardized residual diagnostic tests to evaluate the fitness of the model to the data. These appear in Table 3, where we observe the mean and variances of the standardized residuals to be consistent with values of zero and one respectively, for all the economies at any level of significance (apart from Russia as it concerns the mean values of both series which are different than zero at the 10% level). In addition, the Ljung-Box statistics indicate the absence of serial correlation up to

[13] The lag structure of the VARMA process differs across our sample according to the Schwartz Bayesian information criterion. The resulting choices of p and q appear in the Appendix, Table A1.

[14] These results are in line with the outcome of the preliminary Engle and Ng (1993) test reported in footnote 12.

twelfth order in the standardized and squared standardized residuals in both inflation and the change in currency substitution equations, suggesting that the model is well specified.

Having established the appropriateness of our specification, next we concentrate in the description of the coefficients of our concern, ψ_{ij} and δ^*_{ij}.[15] The coefficients ψ_{11} and ψ_{12} examine the impact of a change in CS uncertainty and inflation uncertainty on the change of CS respectively.[16] Similarly, ψ_{21} and ψ_{22} test for the impact of a change in CS uncertainty and inflation uncertainty on the rate of inflation. The (a)symmetry-related coefficients δ^*_{ij} express the own- and cross-variance asymmetry of CS and inflation. In particular, δ^*_{11} (δ^*_{22}) indicates whether CS (inflation) displays positive, negative or zero own-variance asymmetry. In general, the individual significance and sign of each coefficient provide information about the effects of uncertainty on CS and inflation and its potential asymmetric and spillover effects. These considerations, in turn, provide further insights for the determinants of inflation and CS, especially in the process of explaining CS hysteresis and designing monetary policies suitable for high-inflation and emerging market economies, as in our sample.

Table 4 presents the related findings of our estimation methodology. A first important result, not considered in the traditional literature of the determinants of CS, becomes apparent from the estimates of ψ_{11}. In half of the countries (Bulgaria, the Czech Rep., Estonia, the Kyrgyz Rep., Romania, and Ukraine) the volatility of CS positively influences its own average rate in a significant way. Therefore, although in the remaining countries such a relationship is found to be insignificant at conventional levels, uncertainty about the future level of CS deserves attention as an additional explanatory factor of CS. As a result, countries that wish to discourage the use of foreign currency would benefit by employing policies that aim to reduce the uncertainty associated with the future levels of CS.

Furthermore, the values of ψ_{12} depict the significance of inflation uncertainty for holdings of foreign currency. For Bulgaria, Georgia, the Kyrgyz Rep., Latvia, Poland, and Russia uncertainty about future inflation leads to higher CS as a precautionary

[15] Details on *all* the estimated coefficients of equations (1) and (2) are not provided due to space considerations. The results, however, are available upon request from the authors.

[16] This implies that the uncertainties related with the change in CS and inflation also affect the *level* of CS, since, for instance, positive values of ψ_{11} and ψ_{12} increase the change in CS and subsequently its current level.

motive against the erosion of the domestic currency.[17] These results support Ize and Levy-Yeyati's (2003) theory of the minimum variance portfolio allocation, where CS is partially explained by the second moments of inflation as agents change the currency composition of their portfolios trying to hedge against changes in the distribution of expected returns. An important implication of this theory is that shifts in inflation do not play any role in the decision to hold foreign currency denominated assets; what matters is the volatility of inflation. The authors present empirical support for their approach for a sample of five Latin American countries. Similar results about the effect of inflation volatility on CS are attained by Piontkovsky (2003) for nine transition economies, De Nicoló et al. (2003) for 100 countries, and more recently by Rennhack and Nozaki (2006) for 62 developing countries. Therefore, we view our results as complementing these authors' findings, who, in contrast to our methodology however, utilize a panel data estimation approach.

The potential impact of the uncertainty related with CS on the rate of inflation is captured by ψ_{21}. The results for the eight countries that appear in bold illustrate the importance of CS uncertainty for inflation (Armenia, Bulgaria, Georgia, Kyrgyz Republic, Latvia, Romania, Turkey, and Ukraine), where for seven of them the effect is significantly positive (except for Latvia). From this, it follows that studies that support the notion that a high degree of CS does not seriously impede the effective conduct of monetary policy (see Reinhart et al. 2003 and Havrylyshyn and Beddies 2003), and, therefore, high CS and monetary control can coexist, do not seem to fully account for the effects of CS. In particular, by restricting themselves in the study of the direct effects of CS on inflation, they ignore the second-order adverse effects of CS on inflation through its uncertainty.

Finally, ψ_{22}, the most studied element of matrix Ψ, describes the effect of inflation uncertainty on inflation. Theoretical arguments of this relationship were developed by Cukierman and Meltzer (1986), on the one hand, and Holland (1995), on the other. As described in the introduction, the first predict a positive relationship while the second a negative. Our results strongly support the Cukierman-Meltzer hypothesis for the majority

[17] Surprisingly enough the inverse relationship holds for Ukraine and marginally for the Czech Rep., where foreign currency holdings are reduced because of high inflation uncertainty.

of the countries (Bulgaria, the Czech Rep., Estonia, Georgia, Kyrgyz Rep., Latvia, and Poland), while evidence for the Holland hypothesis is found for Armenia and Ukraine. In addition, the findings for Romania, Russia, and Turkey contradict both hypotheses since inflation uncertainty does not affect inflation. These results accord well with recent evidence on this issue in studies that utilize the GARCH approach in a bivariate model of the joint determination of output growth and inflation. For example, Baillie et al. (1996) find evidence supporting the Cukierman-Meltzer hypothesis in the UK, Argentina, Brazil, and Israel. Similar results are reported by Grier and Perry (1998) for Japan and France, and Bredin and Fountas (2005) for Canada, Germany, and Italy. Grier et al. (2004) and Bredin and Fountas (2005), on the other hand, differ from these studies since they reject the Cukierman-Meltzer hypothesis for the US. In particular, the first support Holland (1995), while the second reject both of the hypotheses. The current study, however, is the first to our knowledge that examines such issues for economies in transition.

In addition to the uncertainty effects just described, the lower panel of Table 4 depicts the importance of the asymmetries and spillovers in the variance-covariance structure for inflation and CS. The results highlight this consideration since most of the countries have at least one element of matrix D^*_{11} that appears statistically significant (exceptions are Poland and Russia). In particular, we emphasize the significance of δ^*_{11} for eight countries of our sample indicating that CS displays own-variance asymmetry. This means that in Armenia, Bulgaria, Kyrgyz Rep., Romania, Turkey, and Ukraine positive innovations in CS lead to higher uncertainty about CS than negative innovations of equal size do. This, in turn, implies the presence of CS ratchet effects in the volatility of CS.[18] That is, the variability of CS displays persistence due to the asymmetric substitution between domestic and foreign currencies. The own-variance asymmetry of inflation, on the other hand, is not that extensive in the sample as illustrated by δ^*_{22}. There are, however, countries (Bulgaria, Latvia and Romania) whose results are in line with the US experience (Grier et al. 2004, Bredin and Fountas 2005) and that of the UK, Egypt, Morocco, India, Colombia, and Peru (Daal et al. 2005), as they exhibit positive asymmetry to inflationary shocks.

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[18] A study that also used the degree of CS as a ratchet variable is Mongardini and Mueller (2000). They, however, found ratchet effects on the level of CS.

3.2. Discussion and Policy Implications

The results presented above clearly illustrate the statistical significance of inflation and CS uncertainty for explaining the behaviour of average inflation and CS. These results, in effect, provide an explanation for the observed hysteresis in CS. They also carry important implications for the effectiveness of monetary policy and the potential benefits these emerging economies can gain by committing themselves to inflation targeting or during their integration process to the EU, as some of them are already (candidates to become) members.

Our evidence that uncertainty related with both CS and inflation has an enhancing effect on the level of CS, as captured by the significance of ψ_{11} and ψ_{12} in most countries, and also the finding of inertia in the variability of CS, as illustrated by δ^{*}_{11}, suggest that the persistence of CS is a rational response of agents to macroeconomic instability. This argument is in line with Peiers and Wrase (1997), Piontkovsky (2003), and Rennhack and Nozaki (2006), who recognize that even if the macroeconomic fundamentals improve substantially within a short time horizon agents will not switch back to the domestic currency before they become convinced of the stability of the current macroeconomic situation. Although these studies have mostly focused on the rate of inflation as an indicator of the macroeconomic environment and the necessity to maintain it in low and stable levels to encourage the use of the domestic currency, we also found the stability of CS to be an additional determinant of this process. Therefore, according to our results, confidence in the domestic currency can be promoted by sustainable policies that aim to reduce the uncertainty associated with the future levels of both CS and inflation.

An issue relative to the effective use of policies in emerging market economies with high levels of CS refers to the ability of monetary policy to succeed its goals of a low and stable rate of inflation and minimum volatility of aggregate output (Fischer 1994). Concentrating on the first of these two goals, recent studies by Reinhart et al. (2003), De Nicoló et al. (2003), and Havrylyshyn and Beddies (2003) find that a high degree of CS does not inhibit the design and effective conduct of monetary policy, in particular with respect to the achievement of inflation control. By showing that the degree of CS has no distinctive effect on the duration of disinflation and simultaneously that successful disinflations generally have not been accompanied by large declines in the level of CS, they infer that high CS can co-exist with low inflation. Therefore, they do not find any first-order adverse effects of CS on monetary policy in the form of inflationary pressures. Our evidence, however, as represented by the strong significance of ψ_{21} for the majority of the countries in our sample, suggests that CS has important augmenting second-order effects on the rate of inflation. Hence, it appears that an effective monetary policy would need to consider, among other factors, the influence of CS variability in retaining a stable inflation rate.

With the above issues in mind, a policy that could potentially reverse the process of CS and at the same time reduce its related uncertainty by promoting macroeconomic stability is inflation targeting (IT). This notion is best described by Ize and Levy-Yeyati (2003) and De Nicoló et al. (2003) who argue that IT joined with nominal exchange rate flexibility should discourage the use of foreign currencies and gradually reduce CS. Further empirical evidence, centred on emerging economies, reinforces this argument since countries that have opted for the IT regime experienced greater stability in the form of lower levels of inflation, and inflation and growth volatility (Gonçalves and Salles 2005). However, as pointed out by Mishkin (2000), a serious problem these countries may face as they adopt a floating exchange rate is abrupt depreciations in the domestic currency that would raise the burden of the foreign currency-denominated debt and, thus, increase the risks of a financial crisis. Therefore, a viable IT policy for economies with high CS would require strong fiscal and financial institutions to ensure that the economy can successfully combat exchange rate shocks, and strong institutional commitment to price stability and instrument independence of the monetary authorities (Mishkin and Savastano 2001, Mishkin 2004).

The above issues and our results of the relationship between uncertainty in CS and inflation and average rates of CS and inflation acquire exceptional importance for the countries of our sample that are new members of the EU as of May 1, 2004 (Czech Rep., Estonia, Latvia, and Poland) and for the countries that are candidates for membership in a future date (Bulgaria, Romania, and Turkey). With reference to these seven countries, our evidence suggests that the decline in both the variability of CS and inflation due to the eventual adoption of the single currency, once the countries meet the convergence criteria, will prove beneficial for their welfare. As illustrated by the estimates of ψ_{21} and

ψ_{22}, the *net* expected gains from joining the EMU and surrendering the independent monetary policy to the ECB are significant. In particular, notice the evidence in favour of the Cukierman-Meltzer hypothesis in all the EU-involved countries, except for Romania. Therefore, the countries of which the monetary authorities display high inflationary bias in the pre-EMU period will benefit the most from the European integration. Finally, a movement towards unilateral euroization for countries currently ineligible to join the EU seems to be supported by the evidence for Georgia, the Kyrgyz Republic, and Ukraine.

4. Conclusion

The purpose of this paper has been to examine the impact of inflation and CS uncertainty on the average rates of inflation and CS. We have addressed this issue with the use of a bivariate GARCH-in-Mean model that allows us to simultaneously estimate the effects of the dynamic volatilities of monthly inflation and CS on their conditional means. By using an approach that considers for the possibilities of asymmetries and spillovers in the variance-covariance structure we avoid a potential specification error and at the same time nest other simpler symmetric and diagonal models.

Our results for a sample of twelve emerging market economies depict that the volatility of inflation and CS exhibit statistically significant effects on the average levels of inflation and CS for the majority of the countries. More specifically, the estimated conditional variance coefficients provide support to the minimum variance portfolio allocation model of Ize and Levy-Yeyati (2003) with the disclose of positive effects of inflation volatility on the level of CS. Similarly strong effects are also obtained in favour of the Cukierman-Meltzer hypothesis where in eight countries from our sample a change in the degree of uncertainty surrounding the rate of inflation has been found to create inflationary pressures. Finally, a new element in our analysis has been the consideration of the variability of CS, which has emerged as an important determinant of both the degree of CS and the rate of inflation since in most cases it created enhancing effects.

These findings emphasize the importance of macroeconomic stability, both in terms of more stable rates of inflation and CS, as a necessary step for encouraging the use of the domestic currency in countries with high levels of CS, and for promoting the reversal of the CS process. They also highlight, contrary to recent studies on the effectiveness of monetary policy in such countries, that although CS may not have direct impeding effects on the control of the inflation rate, it does have second-order effects through its variability. Therefore, the implementation of inflation targeting has been proposed as an appropriate policy for these countries as long as they take the necessary steps to strengthen their policy institutions. In general, our methodological approach and results characterize the significance of uncertainty on the conduct of monetary policy and the proper choice of exchange rate regimes, and highlights the need for further research on this area.

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Table 1. Summary Statistics and Diagnostic Tests

Panel A: Summary Statistics

Country	Sample Period	Obs.	Mean		Std. Dev.		Skewness		Kurtosis	
			ΔCS	Infl.	ΔCS	Infl.	ΔCS	Infl.	ΔCS	Infl.
Armenia	Jun1995-Aug2005	127	0.09	6.29	25.49	28.60	-0.67	0.23	4.35	3.64
Bulgaria	Jul1995-Mar2005	116	0.84	39.35	32.41	150.8	1.78	7.88	15.05	73.25
Czech Rep.	Jan1993-Mar2004	134	-0.19	5.74	5.19	7.95	3.74	2.08	26.54	10.18
Estonia	Jan1993-Aug2005	151	-0.01	11.20	12.84	15.20	-1.35	2.51	16.38	12.14
Georgia	Jan1995-Aug2005	127	0.81	9.97	27.51	29.74	-3.30	4.94	24.74	38.37
Kyrgyz Rep.	Jan1996-Sep 2005	116	3.50	12.81	22.92	22.28	-0.63	0.85	6.20	5.29
Latvia	Jan1993-Sep2004	140	-0.55	9.48	15.00	14.75	1.08	2.39	19.41	13.39
Poland	Jan1993-Aug2005	151	-0.72	11.08	6.33	12.69	1.19	1.32	12.89	5.41
Romania	Dec1990-Jun2005	174	2.16	52.07	56.39	57.48	2.31	0.88	9.40	17.21
Russian Fed.	Dec1993-Jun 2005	138	-0.99	34.85	20.76	50.39	0.83	3.66	13.28	21.31
Turkey	Jan1990-May 2005	232	1.18	45.47	16.75	30.33	-0.17	1.81	4.92	13.52
Ukraine	Dec1992-Jul 2005	151	1.15	35.21	24.61	62.61	1.69	4.23	11.90	26.02

Panel B: Diagnostic tests

Country	Jarque-Bera		Q(12)		$Q^2(12)$		ADF tests	
	ΔCS	Infl.	ΔCS	Infl.	ΔCS	Infl.	ΔCS	Infl.
Armenia	19.16 [0.00]	3.32 [0.19]	71.231 [0.00]	80.552 [0.00]	43.916 [0.00]	31.603 [0.00]	-12.613 [0.00]	-7.275 [0.00]
Bulgaria	762.99 [0.00]	25050 [0.00]	24.554 [0.02]	47.382 [0.00]	35.784 [0.00]	1.456 [0.99]	-8.346 [0.00]	-7.265 [0.00]
Czech Rep.	3407.3 [0.00]	384.31 [0.00]	40.935 [0.00]	91.077 [0.00]	33.701 [0.00]	48.865 [0.00]	-10.934 [0.00]	-5.848 [0.00]
Estonia	1173.1 [0.00]	684.66 [0.00]	43.579 [0.00]	367.55 [0.00]	42.659 [0.00]	95.827 [0.00]	-12.702 [0.00]	-5.651 [0.00]
Georgia	2731.2 [0.00]	7135.7 [0.00]	59.334 [0.00]	29.281 [0.00]	36.283 [0.00]	34.666 [0.00]	-11.721 [0.00]	-4.886 [0.00]
Kyrgyz Rep.	56.92 [0.00]	39.31 [0.00]	38.010 [0.00]	61.502 [0.00]	38.923 [0.00]	51.718 [0.00]	-15.831 [0.00]	-5.941 [0.00]
Latvia	1597.5 [0.00]	763.30 [0.00]	42.964 [0.00]	207.35 [0.00]	37.122 [0.00]	59.569 [0.00]	-12.136 [0.00]	-5.618 [0.00]
Poland	651.26 [0.00]	80.22 [0.00]	34.215 [0.00]	317.57 [0.00]	39.106 [0.00]	139.45 [0.00]	-6.222 [0.00]	-4.890 [0.00]
Romania	451.10 [0.00]	1486.4 [0.00]	210.59 [0.00]	157.33 [0.00]	137.73 [0.00]	32.657 [0.00]	-6.622 [0.00]	-4.487 [0.00]
Russian Fed.	624.05 [0.00]	2235.5 [0.00]	41.636 [0.00]	177.88 [0.00]	43.999 [0.00]	29.826 [0.00]	-11.467 [0.00]	-3.354 [0.02]
Turkey	36.89 [0.00]	1195.9 [0.00]	34.327 [0.00]	45.026 [0.00]	90.677 [0.00]	26.848 [0.01]	-13.437 [0.00]	-8.717 [0.00]
Ukraine	570.24 [0.00]	3784.6 [0.00]	27.158 [0.01]	267.79 [0.00]	33.811 [0.00]	68.595 [0.00]	-9.122 [0.00]	-12.001 [0.00]

Notes: *p*-values appear in squared brackets. $Q(p)$ and $Q^2(p)$ are the Ljung-Box test statistics for *p*th order serial correlation in the standardized and squared standardized residuals respectively. The ADF test statistic represents the values obtained when applied with a constant and trend for the ΔCS in Turkey and for inflation in Bulgaria, the Czech Republic, Latvia and Turkey. For the rest of the countries the ADF test statistic represents the values obtained when applied only with a constant.

Table 2. Hypotheses Tests for the Volatility in Mean and Asymmetry Terms

Panel A: $H_0 : \psi_{ij} = 0$, for all *i, j*		
Country	**Likelihood ratio test**	**p-value**
Armenia	10.28	[0.036]
Bulgaria	21.84	[0.000]
Czech Rep.	48.10	[0.000]
Estonia	54.54	[0.000]
Georgia	21.80	[0.000]
Kyrgyz Rep.	15.76	[0.000]
Latvia	56.04	[0.000]
Poland	23.92	[0.000]
Romania	18.30	[0.001]
Russian Fed.	13.04	[0.011]
Turkey	47.48	[0.000]
Ukraine	8.22	[0.084]
Panel B: $H_0 : \delta^*_{ij} = 0$, for all *i, j*		
Country	**Likelihood ratio test**	**p-value**
Armenia	37.56	[0.000]
Bulgaria	37.50	[0.000]
Czech Rep.	8.10	[0.088]
Estonia	37.00	[0.000]
Georgia	1.48	[0.830]
Kyrgyz Rep.	18.34	[0.000]
Latvia	30.54	[0.000]
Poland	2.12	[0.714]
Romania	63.52	[0.000]
Russian Fed.	4.7	[0.320]
Turkey	29.50	[0.000]
Ukraine	32.86	[0.000]

Note: *p*-values appear in squared brackets.

Table 3. Standardized Residual Diagnostics

Country	Mean		Variance		Q(12)		Q^2(12)	
	$\varepsilon_{\Delta CS}$	$\varepsilon_{Infl.}$	$\varepsilon_{\Delta CS}$	$\varepsilon_{Infl.}$	$\varepsilon_{\Delta CS}$	$\varepsilon_{Infl.}$	$\varepsilon_{\Delta CS}$	$\varepsilon_{Infl.}$
Armenia	0.034 [0.649]	0.124 [0.081]	1.001 [0.497]	0.943 [0.325]	16.597 [0.120]	10.307 [0.503]	4.478 [0.954]	9.195 [0.604]
Bulgaria	-0.051 [0.291]	0.051 [0.291]	0.882 [0.184]	0.831 [0.100]	9.829 [0.456]	15.832 [0.198]	7.927 [0.636]	8.904 [0.541]
Czech Rep.	0.016 [0.427]	0.006 [0.472]	0.905 [0.218]	0.920 [0.256]	3.174 [0.988]	17.190 [0.142]	1.242 [0.998]	10.634 [0.474]
Estonia	0.047 [0.282]	0.007 [0.466]	1.017 [0.441]	0.991 [0.469]	10.474 [0.575]	13.553 [0.330]	5.842 [0.924]	6.146 [0.906]
Georgia	-0.089 [0.158]	-0.050 [0.287]	0.961 [0.378]	1.076 [0.272]	7.140 [0.788]	12.818 [0.616]	9.122 [0.611]	3.134 [0.989]
Kyrgyz Rep.	0.020 [0.415]	0.031 [0.369]	1.058 [0.329]	1.089 [0.249]	8.857 [0.546]	8.901 [0.542]	5.752 [0.836]	14.657 [0.145]
Latvia	-0.012 [0.443]	-0.028 [0.370]	1.028 [0.593]	1.023 [0.576]	9.122 [0.611]	7.591 [0.749]	15.732 [0.151]	15.123 [0.177]
Poland	0.016 [0.422]	0.001 [0.495]	1.015 [0.448]	0.991 [0.469]	8.559 [0.740]	12.761 [0.309]	11.853 [0.458]	5.305 [0.870]
Romania	0.016 [0.416]	0.032 [0.337]	0.966 [0.376]	0.987 [0.452]	7.016 [0.798]	15.432 [0.164]	13.757 [0.184]	16.483 [0.124]
Russian Fed.	-0.138 [0.056]	-0.126 [0.074]	1.075 [0.267]	0.985 [0.450]	10.965 [0.446]	13.596 [0.256]	9.556 [0.571]	16.412 [0.127]
Turkey	-0.034 [0.302]	-0.016 [0.404]	0.975 [0.394]	0.885 [0.108]	17.874 [0.120]	11.593 [0.395]	4.391 [0.975]	4.318 [0.977]
Ukraine	-0.028 [0.365]	-0.064 [0.216]	0.928 [0.266]	0.949 [0.329]	15.659 [0.207]	17.919 [0.118]	4.014 [0.983]	9.519 [0.658]

Notes: p-values appear in squared brackets. *Q(p)* and $Q^2(p)$ are the Ljung-Box test statistics for *p*th order serial correlation in the standardized and squared standardized residuals respectively.

Table 4. Asymmetric Multivariate GARCH-M Model (ψ and δ coefficients)

Country	ψ_{11}	ψ_{12}	ψ_{21}	ψ_{22}
Armenia	0.068 [0.715]	-0.013 [0.928]	**0.448** **[0.000]**	**-0.426** **[0.001]**
Bulgaria	**0.141** **[0.001]**	**0.142** **[0.003]**	**0.248** **[0.073]**	**0.492** **[0.013]**
Czech Rep.	**0.951** **[0.041]**	**-0.424** **[0.098]**	-0.795 [0.156]	**0.824** **[0.074]**
Estonia	**0.607** **[0.072]**	-0.654 [0.130]	-0.228 [0.395]	**0.990** **[0.012]**
Georgia	0.139 [0.712]	**0.160** **[0.080]**	**0.098** **[0.054]**	**0.176** **[0.046]**
Kyrgyz Rep.	**0.348** **[0.040]**	**0.240** **[0.041]**	**0.363** **[0.083]**	**0.573** **[0.069]**
Latvia	0.019 [0.953]	**0.449** **[0.093]**	**-0.897** **[0.003]**	**1.953** **[0.000]**
Poland	-0.069 [0.754]	**0.229** **[0.019]**	0.019 [0.957]	**0.843** **[0.003]**
Romania	**0.285** **[0.004]**	0.020 [0.815]	**0.470** **[0.000]**	0.361 [0.122]
Russian Fed.	0.114 [0.973]	**0.387** **[0.004]**	-0.140 [0.964]	1.208 [0.148]
Turkey	0.283 [0.301]	0.105 [0.459]	**0.679** **[0.013]**	0.271 [0.254]
Ukraine	**0.732** **[0.000]**	**-0.475** **[0.010]**	**0.569** **[0.018]**	**-0.445** **[0.038]**

Country	δ^*_{11}	δ^*_{12}	δ^*_{21}	δ^*_{22}
Armenia	**0.555** **[0.011]**	**0.979** **[0.000]**	-0.052 [0.430]	-0.145 [0.403]
Bulgaria	**0.334** **[0.052]**	-0.084 [0.444]	**0.626** **[0.017]**	**0.981** **[0.000]**
Czech Rep.	-0.024 [0.738]	-0.006 [0.946]	**-0.338** **[0.001]**	0.001 [0.986]
Estonia	**-0.246** **[0.011]**	0.224 [0.171]	**-0.244** **[0.052]**	**-0.333** **[0.047]**
Georgia	**-0.067** **[0.000]**	0.105 [0.730]	0.089 [0.478]	-0.085 [0.595]
Kyrgyz Rep.	**0.537** **[0.071]**	0.011 [0.979]	0.396 [0.442]	**-0.822** **[0.000]**
Latvia	-0.152 [0.435]	**1.494** **[0.000]**	0.007 [0.958]	**0.311** **[0.035]**
Poland	0.350 [0.220]	0.100 [0.565]	0.036 [0.798]	-0.037 [0.467]
Romania	**0.821** **[0.000]**	0.248 [0.119]	0.040 [0.745]	**0.428** **[0.043]**
Russian Fed.	0.178 [0.749]	0.072 [0.537]	0.588 [0.551]	2.010 [0.701]
Turkey	**0.601** **[0.001]**	0.051 [0.535]	0.141 [0.400]	0.020 [0.572]
Ukraine	**0.805** **[0.000]**	**-0.394** **[0.000]**	**0.738** **[0.000]**	**-0.524** **[0.000]**

Notes: p-values appear in squared brackets. Significant coefficients appear in bold.

Appendix

Table A1. VARMA order

Country	AR(*p*) order	MA(*q*) order
Armenia	1	0
Bulgaria	1	1
Czech Rep.	1	0
Estonia	1	0
Georgia	1	0
Kyrgyz Rep.	1	2
Latvia	1	0
Poland	1	2
Romania	1	2
Russian Fed.	1	1
Turkey	1	6
Ukraine	3	2

Table A2. Model Specification Tests

Country	Model Test	*p*-value
Armenia	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Bulgaria	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Czech Rep.	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Estonia	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Georgia	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Kyrgyz Rep.	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Latvia	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Poland	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Romania	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Russian Fed.	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Turkey	No GARCH	[0.000]
	Diagonal GARCH	[0.000]
Ukraine	No GARCH	[0.000]
	Diagonal GARCH	[0.000]

Notes: the No GARCH test refers to testing the null hypothesis of $H_0 : \alpha^*_{ij} = b^*_{ij} = \delta^*_{ij} = 0$, while the Diagonal GARCH hypothesis test is $H_0 : \alpha^*_{12} = \alpha^*_{21} = b^*_{12} = b^*_{21} = \delta^*_{12} = \delta^*_{21} = 0$.