South African Reserve Bank Working Paper Series
WP/14/09

Variance Bounds as Thresholds for 'Excessive' Currency Volatility: Inflation Targeting Emerging Economies

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Authorised for distribution by Chris Loewald

December 2014

South African Reserve Bank

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Variance Bounds as Thresholds for 'Excessive' Currency Volatility: Inflation Targeting Emerging Economies *

Shaista Amod † Shakill Hassan ‡

November 2014

Abstract

At what level does a currency's volatility become 'excessive', in a concrete sense? Any claim that an exchange rate is excessively volatile needs a benchmark for 'normal' variability. We compute variance bounds implied by exchange rate models as the norm, for a set of particularly volatile emerging market currencies; and find that *long-run* exchange rate volatility does not breach the upper bound implied by the present value of underlying fundamentals – for each currency in our sample, except the Brazilian real. However, nominal exchange rate variances get closer to implied upper bounds under inflation targeting. We also find a reduction in real exchange rate misalignment under inflation targeting.

Keywords: currency volatility; variance bounds; monetary exchange rate models; inflation targeting; emerging markets.

JEL Classifications: F31; E52

*This paper previously circulated under the title 'Are Emerging Market Currencies Excessively Volatile Under Inflation Targeting?'. With thanks, for helpful comments and/or discussions, to Chris Loewald, Ugo Panizza, Stan du Plessis, John Taylor, Cédric Tille, Nicola Viegi; to seminar participants at the Graduate Institute of International and Development Studies, Geneva; the South African Reserve Bank; University of Pretoria; and University of Stellenbosch; and to an anonymous external referee. We are entirely responsible for any errors, of course. The views expressed herein are those of the authors, and not necessarily those of the South African Reserve Bank.

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1 Introduction

1.1 Motivation

Consider an initial state where the volatility of a given currency is at a level with which most economic agents are comfortable with. Now suppose the volatility of the exchange rate starts to rise. At what level does a currency's variability become 'excessive', in a concrete sense?

The case for 'sand in the wheels' of currency markets rests largely (though not entirely) on exchange rates being 'too volatile' relative to a defensible benchmark, e.g., some function of the currency's underlying determinants (Eichengreen, Tobin and Wyplosz (1995)). If not, stabilizing the path of the currency's long-term macroeconomic determinants, or/and reducing uncertainty about the evolution of these determinants, is the natural *first step* to reduce (long-term) exchange rate volatility.

The recent volatility of capital flows has brought this discussion back to the centre of policy fora (e.g., IMF (2012), Rey (2014)). The discussion often turns to the role of inflation targeting as the framework for monetary policy. Its adoption in emerging economies either followed or coincided with more flexible exchange rate regimes. It is therefore not surprising that the nominal exchange rates of some inflation targeting countries (but not all), would become more volatile under the new monetary policy regime. Evidence for this shift is mixed, as reported by De Gregorio, Tokman and Valdés (2005), and Berganza and Broto (2012); with evidence to the contrary from Edwards (2007) and Rose (2007).[1]

[1] De Gregorio, Tokman and Valdés (2005), report an increase in short run volatility in Chile. Berganza and Broto (2012) find an increase in short run volatility over a large cross-section of countries (but also that foreign exchange interventions are more effective under inflation targeting). Edwards (2007) reports, for some countries, a reduction in conditional exchange rate volatility once the change in exchange regime is controlled for. He also shows that the introduction of inflation targeting had no significant effect on exchange rate volatility in Australia and Canada, which had floating exchange regimes for prolonged periods of time prior to the adoption of inflation targeting. Rose (2007) finds that once the effects of other factors are controlled for, both nominal and real exchange rate volatility are generally lower for inflation targeters.

Irrespective of whether volatility increased or decreased with inflation targeting, it need not follow that currencies became 'excessively' volatile, relative to the volatility implied by changes in the present value of expected fundamentals; nor that they became more unstable when we look at measures of real exchange rate misalignment. If long-run exchange rate volatility does not breach the maximum level implied by long-term macro fundamentals, stabilizing the latter should precede, or at the very least be concurrent with, interventions aimed at stabilizing exchange rates.

1.2 This paper's contribution

We calculate upper bounds for the long-term variance of a set of emerging market currencies, implied by standard monetary approach specifications where, like other asset prices, the exchange rate is determined by the present value of current and the expected future (relative) path of fundamentals. High long-run volatility is potentially detrimental to economic growth, and it is our focus in the present paper – the conclusions in the present paper do *not* necessarily apply to higher frequency, short-term volatility.

It is shown that: 1) exchange rate volatility is within rational bounds in the sub-samples prior to the adoption of inflation targeting for all countries in the sample, and for all currencies except the Brazilian real, over the period under inflation targeting; but 2) nominal exchange rates moved closer to being 'too volatile' after the adoption of inflation targeting – there is an overall increase in the ratio of exchange rate volatility, to a function of variance in fundamentals.[2] Misalignment, measured by the average deviation of the real effective exchange rate from its equilibrium value, reduced (in five of the six economies) after the adoption of inflation targeting.

The use of variance bounds implied by the standard monetary approach requires justification, given the widely documented failure of monetary models to outperform a random walk in predicting nominal exchange rate levels out-of-sample, at short run horizons (between one month and one year). Our defence is twofold. First, recent evidence supports improvements in forecasting performance in the long run (Mark and Sul (2001), Groen (2005)).[3] Our calculations are based on long term horizons, between 15 and 61 quar-

[2] Our results are mainly presented in terms of variance ratios – the ratio of the variance of a nominal exchange rate to its theoretic upper bound. We emphasize how this ratio changed with the adoption of inflation targeting, rather than statistical tests of the variance inequalities.

[3] See also Balg and Metcalf (2010) for evidence of a long run relationship between exchange rate volatility and the volatility of monetary fundamentals.

ters. Second, monetary models imply that exchange rates are determined by the discounted sum of current and expected future fundamentals, like other asset prices. Engel and West (2005) show that if the discount factor in the present value equation is close to one, equilibrium nominal exchange rates will behave approximately as random walks. Hence the evidence of poor out-of-sample predictability is not inconsistent with the monetary approach. The practical importance of this observation depends on how close discount factors are to unity. Recent empirical evidence, using survey data on exchange rate expectations, points to discount factors around 0.98 (Sarno and Sojli (2009)), comfortably above the 0.90 level required for the Engel and West (2005) result to hold.

We make the connection between the size of the discount factor in the general class of present value equations studied by Engel and West (2005), and the variance bounds implied by monetary models explicit: high discount factors, which imply greater proximity to random walk behavior, also imply larger exchange rate variance bounds. Still, our use of macro variance bounds is driven by the role of macroeconomics as closer to a type of engineering (in the sense of finding solutions to practical questions that may require theoretically imperfect tools) than science, as discussed in Mankiw (2006).

1.3 Related literature

Variance bounds were introduced by Shiller (1981) as tests of stock market efficiency, and adapted to the study of exchange rate behavior by Huang (1981), Vander Kraats and Booth (1983), Honohan and Peruga (1986), Diba (1987), Wadhwani (1987), Bartolini and Bodnar (1996), Bartolini and Giorgianni (2001). Their sample sets consist of advanced economy currencies, primarily over the early post-Bretton Woods period. The findings of the early studies, specifically Huang (1981), Vander Kraats and Booth (1983), and Wadhwani (1987), substantiated the (enduring) view that exchange rates over this period had been excessively volatile for consistency with the behavior of fundamentals. However, Diba (1987) shows that correction for calibration errors in Huang (1981) and Vander Kraats and Booth (1983), results in no evidence of excess volatility. Honohan and Peruga (1986) show that modifying the variance bounds for consistency with exchange rate 'overshooting' also results in no evidence of excessive exchange rate volatility.

We add to this literature by extending the analysis of variance bounds to emerging market currencies, where exchange rate volatility is arguably a more central policy concern; by showing the effect of inflation targeting on the difference between observed volatility and the maximum level justified by fundamentals; and by showing that proximity to, or violation of nominal exchange rate variance bounds, need not imply increased real exchange rate misalignment.

The remainder of the paper is organized as follows. Section 2 is a brief summary of the monetary approach to exchange rates. In section 3 we show the variance bounds implied by flexible and sticky-price exchange rate models, discuss the sensitivity of the bounds to parameter choice, and relate the variance bounds to the discount factor in present value models for the exchange rate. Section 4 contains the empirical findings, including a subsection on real exchange rate misalignment. Section 5 is our brief conclusion.

2 Monetary models of the exchange rate

2.1 Common assumptions

Most exchange rate models used in empirical work are based on the standard monetary approach. The common assumption in flexible and sticky-price variants of this approach is money market equilibrium, given by (with a log-linear functional form for money demand):

$$m_t = p_t + \beta y_t - \alpha i_t, \tag{1}$$

where m_t is the log of domestic money supply, p_t is the log of the domestic price level, y_t is the log of output, and i_t is the level of the domestic interest rate (at time t, with maturity at time $t+1$). A similar relationship holds in the foreign economy. The analogous variables are m_t^*, p_t^*, y_t^* , and i_t^*, with identical parameters α and β in the demand for money equation. α is the interest semi-elasticity of demand for money; β is the income elasticity of demand for money. Equilibrium in the domestic and foreign money markets implies (subtracting the foreign from the domestic equilibrium condition)

$$\begin{aligned} p_t - p_t^* &= x_t + \alpha(i_t - i_t^*), & (2) \\ x_t &= (m_t - m_t^*) - \beta\left(y_t - y_t^*\right), & (3) \end{aligned}$$

where the fundamentals given by x_t reflect ordinary supply and demand effects in the currency market, other than interest-rate effects.

The nominal interest rate differential is tied to the exchange rate through uncovered interest parity:

$$(i_t - i_t^*) = E_t s_{t+1} - s_t \tag{4}$$

where s_t is the log of the domestic currency price of one unit of foreign currency, and $E_t \equiv E(\cdot \mid \Omega_t)$, is the expectations operator conditional on Ω_t, the information set available at time t.

2.2 Standard specification with flexible prices

The flexible price version of the monetary approach (Frenkel (1976), Mussa (1976)) assumes that purchasing power parity always holds [4]

$$s_t = p_t - p_t^*, \tag{5}$$

implying, from equations 2 and 4,

$$s_t = x_t + \alpha \left[E_t s_{t+1} - s_t \right].$$

Combining equations 2, 4, 5, using the law of iterated expectations, and solving, gives the nominal exchange rate as an asset price, with the no-bubbles forward solution given by:

$$s_t = \frac{1}{1+\alpha} E_t \left\{ \sum_{i=0}^{\infty} \left(\frac{\alpha}{1+\alpha} \right)^i x_{t+i} \right\}. \tag{6}$$

Like other asset prices, the exchange rate is determined by the present value of current and the expected future (relative) path of fundamentals. The equation summarizes the most important and robust insight from the monetary approach: exchange rates change in response to changes in *expectations* about variables that affect supply and demand in the (domestic and foreign) money markets. If the path of fundamentals is highly uncertain, and expectations subject to frequent and large revisions, the exchange rate will be more volatile.

2.3 Overshooting model (sticky prices)

If purchasing power parity is not assumed (Dornbusch (1976), Frankel (1979)), combining equations 2 and 4 gives:

$$s_t = \widetilde{x}_t + \widetilde{\alpha} \left[E_t s_{t+1} - s_t \right], \tag{7}$$

with

$$\widetilde{x}_t = (1+\delta)^{-1} \left[(m_t - m_t^*) - \beta (y_t - y_t^*) + \delta (p_t - p_t^*) \right], \tag{8}$$

[4] Strictly speaking, as observed by Gros (1989), it is only required that purchasing power parity always holds in the expected sense, which will be the case if the real exchange rate follows a random walk.

where $\widetilde{\alpha} = \alpha/(1+\delta)$ and $\delta = 1 - \theta - \theta^*$. In this setting, the price level is decomposed into domestic prices and foreign prices expressed in terms of domestic currency. θ is the weight of the domestic component in the price index, given by $\{\theta p_t + (1-\theta)(s_t + p_t^*)\}$.

3 Implied variance bounds

3.1 Variance bound under flexible prices

Equation 6 can be re-written as,

$$(s_t - x_t) = E_t A_t, \tag{9}$$

where

$$A_t = \sum_{i=1}^{\infty} \left(\frac{\alpha}{1+\alpha} \right)^i Dx_{t+i}, \tag{10}$$

and D is the difference operator, defined by $Dx_t \equiv x_t - x_{t-1}$. A_t is the value of $(s_t - x_t)$ under perfect foresight of the subsequent path of Dx_t. Define $u_t \equiv A_t - E_t A_t$, the realized forecast error, linearly independent of Ω_t. Hence, $s_t - x_t = A_t - u_t$, and we have

$$Var(A_t) = Var(s_t - x_t) + Var(u_t) \geq Var(s_t - x_t). \tag{11}$$

Huang (1981), building on the approach introduced by Shiller (1981) and now standard, shows that the following inequality (I1) follows: [5]

$$Var(s_t - x_t) \leq \alpha^2 Var(Dx_t). \tag{12}$$

Note that the inequality is only a necessary but not sufficient condition for acceptance of the monetary model. Its observation does not preclude under-performance of the flexible-price monetary model, relative to a random walk, in out-of-sample exchange rate forecasting. [6]

[5] Huang (1981) tests two other variance bounds, but these are redundant if the first is satisfied.

[6] See also Vander Kraats and Booth (1983), Wadhwani (1987), Bartolini and Bodnar (1996), Bartolini and Giorgianni (2001). Engel and West (2005) develop an argument on the usefulness of monetary exchange rate models even if outperformed in out-of-sample forecasting.

3.2 Variance bound under overshooting (sticky prices)

Honohan and Peruga (1986) reformulate the setting in Huang (1981) to obtain the analogous bound when purchasing power parity is not assumed. The associated inequality (I2), implied by a sticky-price or 'overshooting' version of the monetary approach is the following:

$$Var(s_t - \widetilde{x}_t) \leq \widetilde{\alpha}^2 Var(D\widetilde{x}_t). \tag{13}$$

The same caveat regarding forecasting performance applies.

3.3 Discussion

3.3.1 Plausible values for the interest semi-elasticity of demand for money

Early studies of variance bounds concluded that the exchange rates of the major advanced economies were generally 'too volatile' relative to the variability of their fundamentals in the post Bretton Woods period – or that the exchange rate models were rejected. (See Huang (1981), Vander Kraats and Booth (1983).) These findings were due to errors in the choice of values for α, the interest semi-elasticity of demand for money, to which the calculation of variance bounds is highly sensitive – see inequalities I1 and I2 and note that the left-hand side is independent of α while the right-hand side increases monotonically with α. Diba (1987) shows that once correctly calibrated, the variance bounds are satisfied in the same samples – i.e. the currencies were not too volatile for reasonable values of α (provided the models are valid).[7]

Previous estimates from studies of money demand and exchange rates suggest $\alpha = 10$ as a conservatively low estimate for quarterly data, with estimates up to $\alpha = 60$ (Frankel (1979), Engel and West (2005), Engel, Mark and West (2007)). Consider a range for nominal interest rates between 0.05 and 0.20, or quarterly rates between 0.0125 and 0.05. Using the Baumol-Tobin interest elasticity of money demand coefficient of 0.5 (see Diba (1987)), implies the following range for α :

$$\frac{0.5}{0.05} = 10 \leq \alpha \leq 40 = \frac{0.5}{0.0125}.$$

[7] Huang (1981) set α equal to 3, justified by an interest elasticity coefficient between 0.12 and 0.15, and a range between 0.05 and 0.15 for the nominal interest rate (the semi-elasticity is the ratio of the former to the latter). These values might be appropriate on annual observations, but generate unjustifiably low values for α when using monthly observations, as Huang (1981) does. Vander Kraats and Booth (1983) follow the same calibration.

Alternatively, Obstfeld and Rogoff (2003) show that $\alpha \approx 1/\epsilon\bar{i}$, where i is the steady-state nominal interest rate and ϵ is taken to be marginally above one. At the lower bound of ϵ, it would take a steady state nominal interest rate of 40 per cent per annum or above, for quarterly $\alpha \leq 10$.

We set $\alpha = 10$ in our calculations but, as Diba (1987), present for each exchange rate (and hypothesized value of β), the lowest value of α such that the variance bound is not breached. This is given by

$$\alpha^* = \left[\frac{Var(s_t - x_t)}{Var(Dx_t)}\right]^{\frac{1}{2}} \tag{14}$$

for I1, and by

$$\alpha^* = (1+\delta)\left[\frac{Var(s_t - \widetilde{x}_t)}{Var(D\widetilde{x}_t)}\right]^{\frac{1}{2}} \tag{15}$$

for I2.

This approach allows the policy maker to determine whether the reported value of α^* is too large, given real time interest rates and estimates of the interest elasticity of money demand, when forming a view on whether exchange rate volatility is excessive relative to the volatility of fundamentals.

3.3.2 Relation to discount factor and intuition

For intuition on the sensitivity of the variance bounds to α, and relation to the more recent literature, consider the following general form for present-value models of asset prices (Engel and West (2005)):

$$s_t = (1-b)\sum_{i=0}^{\infty} b^i E_t(\mathbf{a}_1'\mathbf{x}_{t+i}) + b\sum_{i=0}^{\infty} b^i E_t(\mathbf{a}_2'\mathbf{x}_{t+i}), \tag{16}$$

where b is a discount factor, $\mathbf{a}_1$ and $\mathbf{a}_2$ are $n \times 1$ vectors, and $\mathbf{x}$ is the vector of fundamentals.

The monetary model equation 6 is a special case where $b = \alpha/(1+\alpha)$, $\mathbf{a}_1'\mathbf{x}_{t+i} = \left(m_{t+i} - m_{t+i}^*\right) - \beta\left(y_{t+i} - y_{t+i}^*\right)$, and $\mathbf{a}_2'\mathbf{x}_{t+i} = 0$. The simple sticky-price equation 7 corresponds to the special case where $b = \alpha/(1+\alpha)$, $\mathbf{a}_1'\mathbf{x}_{t+i} = (1+\delta)^{-1}\left[(m_t - m_t^*) - \beta\left(y_t - y_t^*\right) + \delta\left(p_t - p_t^*\right)\right]$, and $\mathbf{a}_2'\mathbf{x}_{t+i} = 0$.

In this setting, Engel and West (2005) show that if x_t has a unit root and its first differences follow a first-order autoregressive process, then as b approaches unity: 1) the exchange rate approaches a random walk; and 2) the variance of the exchange rate increases. The connection is that, since α determines the discount factor in the monetary approach, larger values of α imply higher 'justified' volatility.

For intuition, consider a decomposition of $\mathbf{a}_1'\mathbf{x}_{t+i}$ into the sum of a random walk permanent component, and a transitory stationary component (see Engel and West (2005), Engel, Mark and West (2007)). The variance of the random walk component increases in proportion to i, while the variance of the transitory component approaches a constant as i becomes large. So as we increase the time horizon into the future, an increasing share of expected variability in $\mathbf{a}_1'\mathbf{x}_{t+i}$ is driven by the permanent random walk component. As $b \to 1$, the model puts relatively more weight on the effect of changes in fundamentals far in the future (large i) on the current exchange rate. Hence, as the discount factor approaches one (or as α increases), the variance of the present value of fundamentals becomes increasingly dominated by the variance of the random walk component.

4 Empirical results

4.1 Data and basic statistics

We calculated exchange rate variance bounds I1 and I2, from inequalities 12 and 13, for Brazil, Chile, Mexico, South Africa, South Korea and Turkey, using data from the first quarter of 1994 to the second quarter of 2013. These are floating emerging market currencies for which Gagnon and Hinterschweiger (2011) document the highest long-term volatilities.[8] Variations owing to limited data availability are noted below. The variables include domestic exchange rates per US dollar, money supply, seasonally adjusted real gross domestic product (GDP) and consumer price indices (CPIs). All variables were converted by taking logs.

We obtained nominal exchange rates for all countries from Bloomberg, as well as data on money supply, CPIs and GDP for Brazil (1994Q3 to 2013Q2), Mexico, South Korea and Chile. GDP for Chile is from the International Monetary Fund's (IMF) International Financial Statistics (IFS), from 1996Q1 to 2013Q2. Turkish money supply was obtained from the Organization for Economic Cooperation and Development (OECD) and the domestic statistical agency (Turkstat) provided CPI data, while Bloomberg was the

[8] More precisely, these are the six most volatile currencies from the set of currencies under a flexible exchange rate regime, which are also heavily traded. Data from the Bank for International Settlements show that, as of 2013, the Mexican, Turkish, South Korean (no longer categorized as an emerging market), South African, and Brazilian currencies are, respectively, the 8th, 16th, 17th, 18th, and 19th most heavily traded by average daily turnover. Interestingly, quite a few of the most volatile currencies in Gagnon and Hinterschweiger (2011) are those of countries with fixed exchange rate regimes.

Table 1: Basic Statistics

	Brazil		Chile		Mexico		South Africa		South Korea		Turkey	
	Mean	ρ1	*Mean*	ρ1	*Mean*	ρ1	*Mean*	ρ1	*Mean*	ρ1	*Mean*	ρ1
s	0.5861	0.94	6.2747	0.91	2.2792	0.88	1.8772	0.92	6.9777	0.86	0.1454	0.90
	[0.3733]		[0.1520]		[0.2960]		[0.2997]		[0.1775]		[0.5357]	
*m-m**	0.9964	0.95	0.8654	0.95	1.2705	0.95	0.8337	0.96	0.6436	0.93	5.3174	0.93
	[0.4430]		[0.2585]		[0.5121]		[0.3819]		[0.2180]		[0.9497]	
*y-y**	-0.0722	0.96	0.0486	0.96	-0.0391	0.85	0.0266	0.98	0.2018	0.96	-0.1191	0.96
	[0.0753]		[0.1095]		[0.0296]		[0.0725]		[0.1197]		[0.1175]	
*p-p**	2.4110	0.94	-0.8264	0.92	-1.0748	0.94	-1.1508	0.95	-0.8369	0.93	3.6970	0.93
	[0.2295]		[0.0491]		[0.3403]		[0.1918]		[-0.9469]		[0.7033]	
x	1.0325	0.94	0.8411	0.94	1.2901	0.95	0.8204	0.96	0.5427	0.93	5.3769	0.93
	[0.4195]		[0.2083]		[0.5094]		[0.3560]		[0.1732]		[0.9037]	
x_hon	-1.0353	0.94	3.3423	0.94	4.8374	0.96	3.7835	0.96	2.6121	0.92	7.8969	0.93
	[0.7283]		[0.4563]		[0.7833]		[0.6261]		[0.3682]		[1.2436]	
s-x	-0.4464	0.96	5.4336	0.95	0.9891	0.95	1.0513	0.94	6.4350	0.77	-5.2315	0.95
	[0.4330]		[0.2603]		[0.2608]		[0.2443]		[0.1381]		[0.5070]	
Dx	0.0219	0.06	0.0108	-0.02	0.0239	0.40	0.0146	0.19	0.0085	0.64	0.0556	0.49
	[0.0417]		[0.0283]		[0.0276]		[0.0249]		[0.0204]		[0.0605]	

Table reports mean, standard deviation (in parenthesis) and coefficient of autocorrelation (ρ1) for each variable per country over the entire sample (1994Q1 to 2013Q2, exceptions noted in text).

source for GDP (1998Q1 to 2013Q2). US money supply, CPI and GDP is from Federal Reserve Economic Data (FRED). The source for South African money supply and GDP was the South African Reserve Bank (SARB), while CPI data was provided by Statistics South Africa and the SARB. Table 1 shows some basic statistics for the input series.

Each country adopted inflation targeting at a different date, and the sub-sample periods differ accordingly. The shortest sample is however fifteen quarters long, so the reported variances represent long run volatility. Table 2 shows the time periods for the pre- and post-inflation targeting samples for each country.

Table 2: Pre- and Post-IT Sample Periods

	Pre-IT	Post-IT	Float
Brazil	1994 Q3 -- 1999 Q2	1999 Q3 -- 2013Q2	1999 Q1
Chile	1996 Q1 -- 1999 Q3	1999 Q4 -- 2013 Q2	1999 Q3
Mexico	1994 Q1 -- 2000 Q4	2001 Q1 -- 2013 Q2	1994 Q4
South Africa	1994 Q1 -- 1999 Q4	2000 Q1 -- 2013 Q2	1996 Q4
South Korea	1994 Q1 -- 1998 Q1	1998 Q2 -- 2013 Q2	1997 Q4
Turkey	1998 Q1 -- 2005 Q4	2006 Q1 -- 2013 Q2	2001 Q1

Unit root tests using the Augmented Dickey-Fuller and Phillips-Perron methods revealed a unit root in the exchange rate (s_t) and in the fundamentals ($x_t, \widetilde{x}_t$) for all countries except Mexico and Turkey. Although the first-order autocorrelation coefficient ρ_1 is much lower for the right-hand side of the inequality (Dx_t), it remains high for the left-hand side ($s_t - x_t$). However, note that we require only the variance and do not use these variables in a linear regression test.

We test the inequalities over a standard range of possible β values (β=0.5, β=1 and β=1.5) as is common in the literature, rather than imposing one such value. For I2, we set δ=-0.6, calculated using the ratio of imports to gross domestic expenditure in South Africa of $(1 - \theta)$. The value of α is set to 10, and we present the minimal values, as explained in the preceding discussion.

For estimating real exchange rate misalignment, we needed additional series. GDP data for the USA, Japan and the euro area is from Bloomberg, as is the budget balance data for Brazil (1996Q4 to 2013 Q2), Mexico, Turkey and South Korea.[9] GDP, terms of trade and the dependent variable (real

[9] The proxy for the euro area before 1999 is Germany. While studies completed soon

effective exchange rate) are from the World Bank high-frequency dataset for all other countries. The IMF's World Economic Outlook (WEO) database is the source for all population figures.

4.2 Variance bounds before inflation targeting

Consider first the period preceding the adoption of inflation targeting for each country in the sample. Table 3 shows, for each country and for alternative values of the income elasticity parameter β, the sample estimates of the left and right hand sides of inequality I1, given by equation 12, with α set to 10. It also shows, again for each country and alternative values of β, the smallest value of α for which the sample variance does not exceed the theoretic upper bound, given by $\alpha^* = \sqrt{Var(s_t - x_t)/Var(Dx_t)}$.[10]

The inequalities are comfortably satisfied for each country. The values of α^* are very low, given our use of quarterly data, with the exception of South Korea if $\beta = 0.5$. There is therefore no evidence of excessive long-term volatility, for any of the currencies, over the sample periods prior to the adoption of inflation targeting.

Sticky price ('overshooting') models are, by design, better at accounting for exchange rate volatility, and this implies wider bounds. Hence, the absence of any breach under I1 implies no evidence of excessive volatility under I2, which is confirmed by the results in the bottom section of Table 3, where $\alpha^* = (1+\delta)\sqrt{Var(s_t - \widetilde{x}_t)/Var(D\widetilde{x}_t)}$.

4.3 Variance bounds during inflation targeting

The values of α^* (defined in equation 14) increase for all countries, except again for South Korea where it remains relatively unchanged (decreasing slightly), for the sample period beginning with the formal adoption of inflation targeting. This is shown in the upper half of Table 4.

The inequality is violated (i.e. the variance bound is breached) for Brazil, under all values for β; and the left hand side is approximately equal to the right hand side for Chile and South Africa, if $\beta = 0.5$.[11] The finding is illus-

after the euro's introduction used Germany as a proxy for the euro area (see MacDonald and Ricci (2004)), MacDonald and Dias (2007) state that Germany diverges from the euro area in later years and thus is not a good proxy for all times.

[10] Recall that higher values of parameter α correspond to higher variance bounds, or reduced likelihood that a given level of volatility breaches the upper bound.

[11] The inequality is also violated for South Africa if we include the period from the fourth quarter of 2001 to the third quarter of 2002. This was a period of abnormal behaviour in the rand exchange rate, which led to investigations for suspected market manipulation.

Table 3: Sample Values for Inequalites I1 and I2 Before Inflation Targeting

(I1) Pre-IT		β=0.5				β=1.0				β=1.5		
	LHS		RHS	α*	LHS		RHS	α*	LHS		RHS	α*
Brazil	0.0162	<	0.2583	2.51	0.0157	<	0.2924	2.32	0.0160	<	0.3598	2.11
Chile	0.0029	<	0.0451	2.54	0.0023	<	0.0468	2.22	0.0021	<	0.0635	1.81
Mexico	0.0372	<	0.0769	6.95	0.0344	<	0.1035	5.77	0.0321	<	0.1478	4.66
South Africa	0.0043	<	0.0448	3.08	0.0033	<	0.0424	2.79	0.0028	<	0.0419	2.58
South Korea	0.0253	<	0.0304	9.11	0.0223	<	0.0791	5.31	0.0204	<	0.1570	3.60
Turkey	0.0779	<	0.4582	4.12	0.0676	<	0.5595	3.48	0.0588	<	0.6875	2.92

Table reports sample estimates of left hand side (LHS) and right hand side (RHS) of inequality I1, with α
and the smallest values of α for which sample variances do not violate the same inequality.

(I2) Pre-IT		β=0.5				β=1.0				β=1.5		
	LHS		RHS	α*	LHS		RHS	α*	LHS		RHS	α*
Brazil	0.0437	<	8.3143	0.72	0.0558	<	9.9883	0.75	0.0734	<	12.9670	0.75
Chile	0.0191	<	1.6321	1.08	0.0187	<	1.6845	1.05	0.0208	<	2.3223	0.95
Mexico	0.0795	<	2.9046	1.65	0.0711	<	3.1570	1.50	0.0654	<	4.0968	1.26
South Africa	0.0274	<	1.8172	1.23	0.0342	<	1.6928	1.42	0.0439	<	1.6432	1.64
South Korea	0.0703	<	0.8777	2.83	0.0492	<	2.2833	1.47	0.0341	<	4.8290	0.84
Turkey	0.1951	<	14.4332	1.16	0.1635	<	16.8691	0.98	0.1414	<	20.3496	0.83

Table reports sample estimates of left hand side (LHS) and right hand side (RHS) of inequality I2, with α
and the smallest values of α for which sample variances do not violate the same inequality.

Table 4: Sample Values for Inequalities I1 and I2 During Inflation Targeting

(I1) Post-IT	β=0.5				β=1.0				β=1.5			
	LHS		RHS	α*	LHS		RHS	α*	LHS		RHS	α*
Brazil	0.2497	>	0.1411	13.30	0.2121	>	0.1412	12.26	0.1782	>	0.1468	11.02
Chile	0.0848	<	0.0857	9.95	0.0593	<	0.0890	8.16	0.0398	<	0.0972	6.40
Mexico	0.0110	<	0.0330	5.76	0.0091	<	0.0450	4.50	0.0077	<	0.0643	3.47
South Africa	0.0516	<	0.0570	9.52	0.0413	<	0.0594	8.34	0.0341	<	0.0639	7.30
South Korea	0.0158	<	0.0206	8.75	0.0113	<	0.0262	6.56	0.0128	<	0.0381	5.79
Turkey	0.0158	<	0.0475	5.78	0.0128	<	0.0584	4.68	0.0117	<	0.0922	3.57

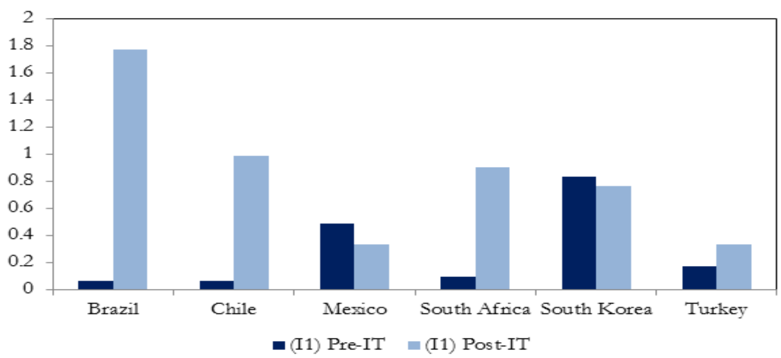
Table reports sample estimates of left hand side (LHS) and right hand side (RHS) of inequality I1, with α
and the smallest values of α for which sample variances do not violate the same inequality.

(I2) Post-IT	β=0.5				β=1.0				β=1.5			
	LHS		RHS	α*	LHS		RHS	α*	LHS		RHS	α*
Brazil	0.6623	<	5.5792	3.45	0.5090	<	5.5407	3.03	0.3780	<	5.7175	2.57
Chile	0.2647	<	3.2066	2.87	0.1569	<	3.3268	2.17	0.0868	<	3.6326	1.55
Mexico	0.0714	<	1.1500	2.49	0.0584	<	1.6261	1.89	0.0481	<	2.3906	1.42
South Africa	0.1655	<	2.4654	2.59	0.1236	<	2.5981	2.18	0.1005	<	2.8134	1.89
South Korea	0.0277	<	0.9183	1.74	0.0233	<	1.1397	1.43	0.0566	<	1.6080	1.88
Turkey	0.0458	<	1.5569	1.72	0.0301	<	1.7886	1.30	0.0270	<	2.9123	0.96

Table reports sample estimates of left hand side (LHS) and right hand side (RHS) of inequality I2, with α
and the smallest values of α for which sample variances do not violate the same inequality.

trated in Figure 1, which shows, for each country, $Var(s_t - x_t)/\alpha^2 Var(Dx_t)$, i.e., the ratio of the long-term variance of the exchange rate (deviations from fundamentals) to the maximum variance justified by the respective country's fundamentals – before and after the adoption of inflation targeting. This ratio increases substantially, after the adoption of inflation targeting, in Brazil, Chile, South Africa and Turkey; and stays relatively unchanged, decreasing slightly, in Mexico and South Korea.

Figure 1: Variance Ratios, Pre- and Post-IT, I1



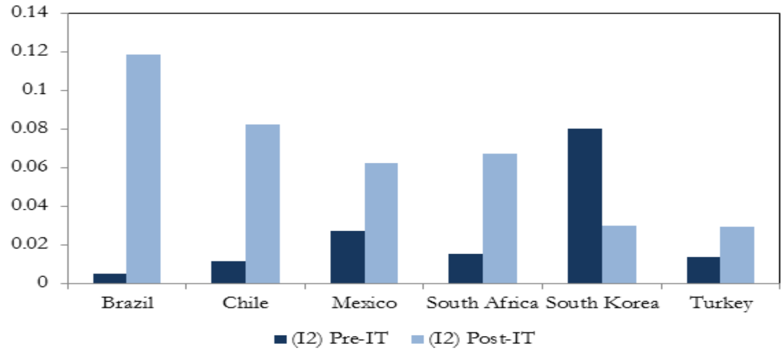
The bounds implied by the overshooting model can account for the excess volatility of the Brazilian real (see the lower half of Table 4), but we observe the same pattern for inequality I2: compared to the sample results for the period(s) prior to inflation targeting, the variances are all closer to the respective upper bounds. This implies an increase in long run exchange rate volatility relative to volatility in fundamentals during inflation targeting.

Figure 2 illustrates the same ratio as Figure 1, using the bounds implied by the sticky-price model. It shows the same pattern, with larger differences between the two sub-samples, and an increase also in Mexico.

In sum, averaging across countries, the ratio of long-term exchange rate variance to its upper bound, increases substantially in the inflation targeting period. In the case of Brazil, the ratio exceeds one for the sample period after the adoption of inflation targeting, indicating exchange rate volatility beyond what can be rationalized by the standard monetary exchange rate model. The ratio is approximately equal to one for Chile, indicating that long-term variance is as high as can be explained by long-term fundamentals. Only South Korea shows a marked reduction in this ratio after the adoption

South Africa adopted inflation targeting in February 2000, with 2002 as the first target year (see Farrell, Hassan and Viegi (2012)).

Figure 2: Variance Ratios, Pre- and Post-IT, I2



of inflation targeting.

Examination of the right hand sides of tables 3 and 4 shows an increase in the variance ratios, for all currencies except South Korea's.

4.4 Misalignment and real exchange rate instability

In terms of implications for economic growth and welfare, the variance of nominal exchange rate changes (in our case the variance of deviation from fundamentals) is not necessarily the best measure of exchange rate instability, as it may not reflect persistent misalignment nor increases in the probability of extreme valuations, as observed by De Gregorio, Tokman and Valdés (2005). [12] Theoretic (and empiric) research shows potential output costs of real exchange rate instability, but not necessarily a clear relationship between economic growth and excessive nominal exchange rate volatility – e.g., Eichengreen (2008), Aghion, Bacchetta, Ranciere, and Rogoff (2009).[13]

We follow De Gregorio, Tokman and Valdés (2005) and Hausmann, Panizza, and Rigobon (2006) in using regression residuals to compute the mean of squared deviations of the real exchange rate around a proxy for its equilibrium value, as an indication of real exchange rate instability.

[12] The possible exception are economies characterized by high levels of foreign currency debt, leading to un-hedged currency mismatch in the aggregate balance sheet. Exchange rate depreciations then raise the (domestic currency) value of aggregate liabilities, without a matching increase in aggregate revenues, leading to output contraction, and constraining monetary policy. See for example Eichengreen, Hausmann and Panizza (2007).

[13] Hausmann, Panizza, and Rigobon (2006) show that the long run volatility of real exchange rates in developing countries is on average two times larger than that of advanced economies.

Our estimates of equilibrium real exchange rates are uncomplicated, and based on De Gregorio, Giovannini, and Wolf (1994), and De Gregorio and Wolf (1994), who introduced demand factors to the Balassa-Samuelson model, where the real exchange rate depends entirely on supply factors. The explanatory variables are relative labour productivity, the log of terms of trade, and the fiscal balance. We rely on a proxy for relative productivity constructed as the ratio of log domestic output per capita, to the trade-weighted log output per capita in the foreign country (as MacDonald and Ricci (2004)). Each 'foreign country' was determined as a weighted average of the USA, Japan, and the euro area, which in sum comprised more than 50 per cent of trade for each of the countries in our sample. The weights were taken from Bayoumi, Lee and Jayanthi (2006). The dependent variable is the real effective exchange rate.

We measure misalignment as the difference between the actual and fitted values from the regression equation. Table 5 reports the mean of squared deviations of the real exchange rate around its estimated equilibrium (rather than its own mean), for the periods before and after the adoption of inflation targeting.

Table 5: Mean Squared Misalignment

	Brazil	Mexico	South Africa	South Korea	Turkey
Pre-IT	0.0368	0.0240	0.0082	0.0640	0.6531
Post-IT	0.0135	0.0141	0.0074	0.0452	1.2724

The table shows a reduction in the measure of real exchange rate instability after the adoption of inflation targeting, for all countries except Turkey. Chile is excluded due to unavailability of quarterly data – De Gregorio, Tokman and Valdés (2005) show a reduction in misalignment volatility after floating and the adoption of inflation targeting in Chile, using the same approach. Hence the group as a whole experienced a reduction in extreme deviations from equilibrium under inflation targeting. To the extent that inflation is lower under inflation targeting regimes, this finding is consistent with research showing that real effective exhange rate volatility increases with inflation (e.g., Bleaney and Francisco (2010)).

5 Conclusion and a caveat

We find little support for claims of 'excessive' exchange rate volatility, when volatility is measured over the long run, and the highest level of volatility consistent with a standard asset pricing equation for the exchange rate, establishes the upper bound for 'normal' volatility. However, we also find an overall increase in exchange rate variance as a ratio of (a function of) volatility in fundamentals, with the adoption of inflation targeting.

A caveat is in order. Our findings are based on second moments calculated over long sample periods before and after adoption of inflation targeting. These do not preclude shorter periods of excessive volatility. For shorter horizons, variance bounds based on monetary models would however be inadequate.

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