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Emerging Floaters: pass-throughs and (some) new commodity currencies

Emanuel Kohlscheen**

Abstract

In spite of early skepticism on the merits of floating exchange rate regimes in emerging markets, 8 of the 25 largest countries in this group have now had a floating exchange rate regime for more than a decade. Using parsimonious VAR specifications covering the period of floating exchange rates, this study computes the dynamics of exchange rate pass-throughs to consumer price indices. We find that pass-throughs have typically been moderate even though emerging floaters have seen considerable nominal and real exchange rate volatilities. Previous studies that set out to estimate exchange rate pass-throughs ignored changes in policy regimes, making them vulnerable to the Lucas critique. We find that, within the group of emerging floaters, estimated pass-throughs are higher for countries with greater nominal exchange rate volatilities and that trade more homogeneous goods. These findings are consistent with the pass-through model of Floden and Wilander (2006) and earlier findings by Campa and Goldberg (2005), respectively. Furthermore, we find that the Indonesian Rupiah, the Thai Baht and possibly the Mexican Peso are commodity currencies, in the sense that their real exchange rates are cointegrated with international commodity prices.

Keywords: commodity currency; exchange rate; emerging market; inflation; pass-through
JEL Classification: E31; E52; E58; F31; F41

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1. Introduction

The succession of exchange rate crises in emerging markets during the 1990s seems to have left at least one lasting mark in the developing world: a greater proportion of countries have chosen to implement a floating rather than a fixed exchange rate regime. Starting with Mexico in late 1994, several governments that were effectively forced off pegs have refrained from setting new pre-announced targets for the level of the exchange rate. In spite of early skepticism on the merits of floating exchange rate regimes in emerging markets (e.g. Carmen Reinhart's "*The Mirage of Floating Exchange Rate Regimes*" (2000), or Guillermo Calvo and Carmen Reinhart's "*Fixing for Your Life*" and "*Fear of Floating*" papers (2000 and 2002, respectively)), 8 of the 25 largest developing countries have now had a floating exchange rate regime for more than a decade.[1]

This paper focuses on the pass-through of exchange rate variations to consumer prices as a key variable which may explain the (perhaps surprising) observed resilience of the exchange rate regimes in this group of countries - which we henceforth call *emerging floaters*. Since our results - as well as those of Campa and Goldberg (2001 and 2005) - show that higher pass-throughs are associated with greater exchange rate volatilities, a high pass-through to final consumer prices typically implies higher inflation variance. Hence, it follows that a country with a high pass-through may want to refrain from letting its currency float if the chief objective of the monetary authority is to stabilize the rate of inflation around a pre-announced target.

In our study we use the nominal and real effective exchange rate indices released by the BIS to estimate the short and long-term pass-throughs for the emerging floaters. For this we resort to parsimonious VAR specifications. We note that, by and large, existing studies have estimated exchange rate pass-throughs ignoring changes in policy regimes. In order to make the estimates less vulnerable to the Lucas critique, however, we depart from these studies in that we restrict the samples to cover a unique

[1] Three of them have a GDP in excess of $750 bn each: Brazil, South Korea, Mexico. The three largest emerging markets that do not have a floating exchange rate regime are India, that has had a *de facto* peg since August 1979, China that has been on a peg since August 1992 and Russia - on a peg since December 1999.

exchange rate regime. Our concern is that the exchange rate regime itself can affect the degree of price stickiness and therefore the exchange rate pass-through.[2]

Contrary to early estimations (e.g. Borensztein and de Gregorio (1999), Calvo and Reinhart (2000), Goldfajn and Werlang (2000)), we find that the pass-throughs in emerging floaters have typically been moderate and that, in some respects, the effects of the exchange rates in these countries resemble those found in developed economies with floating exchange rate regimes. Pass-throughs are very far from being complete even in the long-run and we find no evidence that variations in domestic price levels feed straight back into exchange rate variations.

We then try to explain observed pass-throughs. Overall, we find that the relation between estimated pass-throughs and the level and volatility of inflation rates is not unambiguously a positive one in this group of countries. Pass-throughs however are clearly increasing both with the volatility of the exchange rate and with the presence of homogeneous products in trade flows. The former effect is consistent with the model of Floden and Wilander (2006) that features local currency pricing and price-setters that follow S-s type adjustment rules, whereas the latter is in line with the findings of Campa and Goldberg (2001, 2005) for developed countries. Moreover, the fact that exchange rate volatility is associated with higher - and not lower - pass-throughs clearly contradicts the mechanisms highlighted by Froot and Klemperer (1989), Krugman (1989) and Devereux and Engel (2002).

Finally, since incomplete pass-through renders the PPP assumption invalid, and given that uncovered interest parity has been consistently rejected (see Engel (1996), for instance), we conjecture that the monetary model of the exchange rate is unlikely to become a useful guide in predicting the behaviour of the exchange rate in this group of countries as well. We therefore look at whether commodity prices can explain nominal exchange rate variations of emerging floaters. Based on cointegration and causality tests, we conclude that the only countries in which there are clear indications that variations in exchange rates are linked with variations in international commodity prices are Indonesia, Thailand and Mexico. While we find stable cointegrating relationships for the Indonesian Rupiah and the Thai Baht, we fail to do so for the Mexican Peso.

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[2] The relationship between price stickiness and pass-through in the United States has been analyzed by Gopinath and Itskhoki (2009).

The paper proceeds as follows. Section 2 discusses the country selection criterion of this study and shows that emerging floaters have indeed experienced considerable exchange rate volatilities. Section 3 sets out to estimate short and long-run exchange rate pass-throughs under flexible exchange rate regimes. The section that follows aims to relate estimated pass-throughs to their potential determinants that have been highlighted in the literature. Finally, Section 5 tests whether the currencies of emerging floaters can be described as commodity currencies in the sense that their valuation hinges primarily on international commodity prices. The conclusion provides some directions for further research.

2. Volatile Exchange Rates

A central objective of this paper is to estimate the dynamics of exchange rate pass-throughs and to provide an answer as to whether emerging market currencies are *de facto* commodity currencies. Importantly, the answers to these questions are to be found within a single policy regime, contrasting with previous literature that mixes up different regimes. A crucial step therefore is to distinguish between exchange rate policy regimes. In order to determine whether a country has had a floating exchange rate regime we follow the *de facto* regime classification of Reinhart and Rogoff (2004) and later updates of it by the IMF. All 54 countries for which the Bank of International Settlements regularly publishes exchange rate data are analyzed. This leads to the identification of 8 emerging countries that have had an uninterrupted floating exchange rate regime for at least 10 years. 4 of these are Asian, 2 Latin American, 1 African and 1 Eastern European. Ranked by size those are, respectively, Brazil, South Korea, Mexico, Indonesia, South Africa, Thailand, the Czech Republic and the Philippines. Throughout, we refer to this group of countries as the emerging floaters.

To begin with, we analyze the behaviour of exchange rates since the inception of the floating exchange rate regimes. For this we use the monthly effective nominal and real exchange rates that are regularly published by the Bank of International Settlements. Figure 1 shows that the real effective exchange rate tends to track the nominal effective exchange rate quite closely for all emerging floaters.[3] It is evident

[3] Note that in Figure 1 the nominal effective exchange rate index is inverted, so that a devaluation is associated with an increase in the index.

that, at least in the short run, the nominal exchange rate is the main driver of real exchange rates.

[Figure 1 about here]

[Table 1 about here]

Table 1 shows the calculated (effective) nominal exchange rate volatilities. The table presents the proportion of the sample in which absolute monthly effective exchange rate variations exceeded a given threshold (set to 1, 2 and 5%). It is apparent that the volatility of the nominal effective exchange rates of all emerging floaters is greater than that observed for the United States or the Eurozone. Even though these new floaters practice a managed float, their observed nominal exchange rate volatility is considerable. Moreover, the lower part of the table shows that in general the volatility measures exceed those associated with existing and previous fixed exchange rate regimes in major developing countries. Among the emerging floaters, Brazil, South Africa and Indonesia stand out as having the most volatile exchange rates. [4]

3. Exchange Rate Pass-Throughs: Estimation

The extent of exchange rate pass-through to domestic prices has been a topic of great interest to international macroeconomists in the last few decades. Indeed, the appeal of many theoretical models ultimately hinges on the magnitude of this parameter. Its precise estimation is therefore paramount in an open economy, and in one with a floating exchange rate regime in particular. From an applied perspective, the quantification of pass-through is relevant not only for the conduct of monetary policy geared towards domestic inflation, but also for the assessment of the effectiveness of the

[4] Over time, the exchange rate volatility of the emerging floaters has tended to show a downward trend for most of the time leading up to 2007. Nevertheless there seem to be few changes in relative positions. The exception to this pattern is South Africa that saw a marked increase in exchange rate volatility in the first half of the current decade. The downward trend in volatilities of the emerging market currencies by and large occurs in a period in which the volatility of the G-3 currencies also trended downwards. It is therefore difficult to establish whether the greater stability reflects a maturing process of the new regime or is purely due to a benign global environment up to 2007.

active use of exchange rate policies to correct for eventual international imbalances. A number of analyses have therefore tried to pin down the theoretical determinants of this parameter. Among others, the extent of exchange rate pass-through has been linked with the level and variability of inflation (Taylor (2000), Gagnon and Ihrig (2004), Choudhri and Hakura (2006)), country size (Dornbusch (1987)), openness and trade characteristics (Campa and Goldberg(2001, 2005)), market structure (Krugman (1987), Froot and Klemperer (1989)), and exchange rate variability (Krugman (1989), Froot and Klemperer (1989), Floden and Wilander (2006)). The contribution of this paper is however empirical. In particular, it asks which of these theories turns out to better explain the observed pattern of exchange rate pass-throughs in emerging markets that have allowed their currencies to fluctuate since the mid-1990s.

It should be noted that the paper focuses exclusively on the pass-through to consumer prices as this is likely to be the most relevant indicator when it comes to the selection and sustainability of an exchange rate regime.[5] Obviously, this is by no means the first study to estimate short and long term pass-throughs and our contribution belongs to rapidly growing empirical literature on the subject.[6] However, we differ from other studies that include developing countries (e.g. Goldfajn and Werlang (2000), Frankel, Parsley and Wei (2005), Choudhri and Hakura (2006) and Ca'Zorzi, Hahn and Sanchez (2007)) as they have pooled different monetary and exchange rate regimes together without distinction, casting doubt on the validity of their estimated coefficients. [7] Because of this limitation they are particularly vulnerable to the Lucas critique. Taking advantage of the fact that 8 emerging markets have now had a floating exchange rate regime for more than a decade we proceed to estimate the dynamics of exchange-rate pass-throughs, relying on parsimonious VAR specifications. More specificically, we estimate

[5] Campa and Goldberg (2005) analyze pass-through to import prices, while Frankel, Parsley and Wei (2005) study the pass-through to 8 different product prices.
[6] Previous studies include Borensztein and de Gregorio (1999), Goldfajn and Werlang (2000), Campa and Goldberg (2001, 2005), Frankel, Parsley and Wei (2004), Gagnon and Ihrig (2004), Choudhri and Hakura (2006) and Ca'Zorzi, Hahn and Sanchez (2007).
[7] To our knowledge, the only exception to this pattern is the study of Calvo and Reinhart (2000). Frankel, Parsley and Wei's (2005) interesting dataset spans only the period from 1990 to 2001 - a period with considerable changes in exchange rate regimes in developing countries. The approach of that paper is however considerably different from the one followed here since it effectively imposes complete pass-through in the long-run. In other words, relative PPP is not tested for, but is assumed to hold.

$$X_t = K + \sum_{i=1}^{P} \Phi_i X_{t-i} + \eta_t \ , \qquad (1)$$

where X_t is a vector of endogenous variables, K is a vector of constants and η_t is a vector of white-noise disturbances. Identification is achieved by using a Choleski decomposition of the variance-covariance matrix of the reduced form residuals. To obtain the short and long-term pass-throughs for each country we use nominal effective exchange rates and consumer price data at monthly frequencies. Nominal effective exchange rates were obtained from the Bank of International Settlements, while inflation and output data were taken from the IMF's IFS database. Equation (1) is then estimated using 12 lags. Our sample starts at the inception of the floating exchange rate regime as defined by Reinhart and Rogoff (2004) and extends until October 2008. All data sources are described in detail in the Appendix. As in each case our sample excludes observations prior to the start of the floating exchange rate regime, our results are less subject to the Lucas type critique that implies that pass-through itself may hinge on the policy regime. It is important to note that even though this represents a clear improvement relative to the studies that mix different policy regimes, we cannot claim complete immunity from the Lucas critique because countries may have selected floating exchange rate regimes because they have more or less flexible prices than countries that have chosen fixed regimes.

Our study differs from papers such as Borensztein and de Gregorio (1999) and Goldfajn and Werlang (2000) in at least one other important aspect: we use effective exchange rates, as opposed to bilateral US Dollar exchange rates. This difference is potentially important as pass-through estimates that are based exclusively on the exchange rate against the US Dollar are likely to be biased. This occurs as exchange rates capture not only the variation in value of the domestic currency but also of the foreign. Effective and bilateral exchange rates will therefore typically differ. As effective and bilateral rates are associated with the same observed variation in the price level a potential bias emerges. This bias will lead to an underestimation of the pass-through if the variations in the value of the US currency are more important than those of other international currencies, and to an overestimation if other international currencies are more volatile.[8] The use of effective rates aims at providing a better insulation of the measured pass-throughs from shocks affecting the US economy. Our measures are therefore not affected by the choice of the base country.

[Table 2 about here]

As our first inspection showed that the time series are non-stationary,[9] we ran the VARs in (log) differences using the Cholesky ordering [Δe Δp]. Evidently, one additional advantage of using VARs for the characterization of the dynamics of exchange rate pass-throughs is that they allow for the possibility that price changes that were induced by an initial shock to the exchange rate eventually feedback into new variations of the exchange rate. The accumulated responses to an exchange rate shock are shown in Table 2.

[Table 3 about here]

The first observation is that pass-throughs in emerging floaters show considerable variation between countries. Thailand, the Philippines and Mexico are characterized by very moderate pass-throughs to CPIs, whereas Indonesia and the Czech Republic have the largest coefficients. Secondly, the pass-through is typically fast. Most of the variation in prices occurs within the first six months that follow a shock: the median pass-through for the eight countries is 16.1% after 6 months and 20.6% after a year (averages of 18% and 28.4%, respectively). The Granger-causality tests in Table 3 indicate that causality runs from nominal exchange rate variations to inflation in most of the countries. The results also imply a rejection of the reverse causal link in all cases. Furthermore, the impulse responses show that the reaction of exchange rates to an innovation in the price level is not statistically significant at the usual confidence levels.

[Table 4 about here]

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[8] A second observation that follows is that the measured exchange rate pass-through will typically be more volatile if contract prices are specified in terms of a vehicle currency, as prices will also be affected by changes in the value of the vehicle currency.

[9] Overall, Phillips-Perron and Kwiatwoski *et. al.* unit root tests lead to the rejection of stationarity in 95.4% of the cases.

To check for the robustness of the above results to the particular specification of the VAR, we also run a VAR that includes output data (proxied by the log change in the industrial production index), using the Cholesky ordering [Δy Δe Δp].[10] It is apparent from Table 4 that the estimated pass-throughs are not greatly affected by the change in specification. Relative to the bivariate VAR, the greatest difference is observed for Mexico, with the 6-month pass-through increasing by a modest 2 percentage points, and the 12-month pass-through by 3.2%.[11]

The above estimations tell us that, with the exception of Indonesia and the Czech Republic, less than one third of exchange rate shocks end up being passed into consumer prices after an entire year. Even in Indonesia and the Czech Republic - the countries with the highest estimates - the pass-through is far from complete.[12] All in all, it is fair to say that the CPI pass-throughs in the group of emerging floaters have been moderate. In some cases they are even comparable with the pass-throughs observed in developed countries with well established floating exchange rate regimes.[13]

The results of this Section stand in marked contrast with those of early studies in this literature. For instance, Calvo and Reinhart (2000) found that pass-throughs in emerging markets were on average four times higher than in developed ones. Goldfajn and Werlang (2000) found an average 12-month CPI pass-through of 91.2% in emerging countries, as opposed to 60.5% in developed ones between 1980 and 1998.

4. What Explains Pass-Throughs?

We now turn to the question of relating pass-throughs estimated in the previous section to the variables that have been identified as their potential determinants in the

[10] Our ordering choice here and before follows that of McCarthy (2000). Table A1 in the appendix shows the sensitivity of the variance decomposition to the Choleski ordering in the bivariate case.

[11] Note that this table does not include Indonesia and Thailand due to the lack of data on industrial production indices at monthly frequencies in these countries during the sample period.

[12] In all countries the null of complete pass-through is rejected at the 5% confidence level.

[13] For instance, Choudhri and Hakura (2006) estimate an average accumulated 12-month pass-through of 14% for countries with low average inflation rates. For Germany

literature. Table 5 lists these factors, while Table 6 shows the correlations between these indicators and 6- and 12-month accumulated pass-throughs.

[Tables 5 and 6 about here]

A look at the simple correlations in the two first lines of Table 6 seems to suggest that the countries with higher inflation rates and higher variability of inflation are also the ones with the higher pass-throughs. This link has been emphasized by Taylor (2000), Gagnon and Ihrig (2004) and Choudhri and Hakura (2006), among others. However, a closer inspection reveals that the Spearman rank correlations tells a different story: ranking the countries according to their anti-inflation performance during the period gives a very poor match of the pass-through ranking. To see this, note that these rankings lead to at least two notable exceptions: the Czech Republic, which has had the lowest inflation throughout has very high pass-throughs, whereas Mexico - a country with one of the highest rates of inflation - has a pass-through that resembles that of a developed country.

Second, greater openness to trade, as measured by the average value of trade flows (i.e. exports plus imports) relative to GDP is associated with lower pass-throughs. While this may seem counter intuitive at first, one possibility is that these strongly negative correlations are capturing that more open economies also tend to have lower exchange rate volatilities. This follows from models such as Hau (2000) or Obstfeld and Rogoff (2000) and has been empirically confirmed by Hau (2002). Indeed, in our sample the correlation between openness and exchange rate volatility is -0.763. Third, the last column suggests that the link between pass-throughs and country size is weak and does not allow us to confirm the Dornbusch (1987) prediction that larger countries tend to have lower pass-throughs.

Elsewhere, Campa and Goldberg have argued that the composition of trade is a key determinant of the degree of pass-through. In particular, pass-through is likely to be higher for homogeneous goods that are traded in international markets. To proxy for these characteristics, we consider the share of trade that is made up by food and energy. Table 6 suggests a clear positive association between this measure of trade homogeneity

- a country that had a floating exchange rate regime throughout their sample - they find a pass-through of 13%.

and pass-throughs - a result in line with that obtained by Campa and Goldberg (2005) for developed countries.

Finally, among emerging floaters higher exchange rate volatilities are associated with higher pass-throughs. This is consistent with Floden and Wilander (2006) in which there is local currency pricing and price setters follow S-s type adjustment rules. The positive correlations however contradict theories that associate low pass-throughs with high exchange rate volatilities (e.g. Krugman (1989), Froot and Klemperer (1989) and Devereux and Engel (2002)).[14]

5. Are We Talking of Commodity Currencies?

Incomplete pass-through clearly renders the PPP assumption invalid. If one adds the fact that uncovered interest parity has been consistently rejected, the odds seem to be stacked against the flexible price monetary model of the exchange rate once more. In this section we therefore look at whether commodity prices can explain the variations in the nominal exchange rates of the emerging floaters.

5.1. Cointegration Tests

To explore the possibility that the currencies of the emerging markets analyzed in this paper show the patterns that are typical of commodity currencies, we test whether there is evidence that the real exchange rates of emerging floaters are cointegrated with international commodity prices. This approach has been used earlier in the comprehensive study of Cashin, Cespedes and Sahay (2004). Their sample period however spanned the period from 1980 to March 2002, so that their estimation covered different exchange rate regimes, with a clear predominance of fixed exchange rates in developing countries.[15]

[14] In Krugman (1989) and Devereux and Engel (2001) the rationale is that, as low pass-throughs imply only small substitution effects after a change in the exchange rate, greater exchange rate variations are required for the economy to reach its new equilibrium following a shock. In Froot and Klemperer (1989) lower pass-throughs result when nominal exchange variability is high as exporters try to maintain their market share.

[15] While the authors do not express any concern with the different policy regimes that were practiced during this time interval, they do effectively allow for one single

Again, we face the same issue that the exchange rate regime itself could matter for whether a given currency is identified as a commodity currency or not. Since the real exchange rate is clearly affected by the nominal rate, it may well be that a country's currency over a given time period may not show the characteristics of a commodity currency simply because it is not allowed to fluctuate like a commodity currency in a world with price rigidities.

To test whether the currencies of the emerging floaters are *de facto* commodity currencies or not, we use the non-fuel primary commodity index computed by the IMF - which is based on the prices of about 40 major global primary commodities. The advantage of this broad index is that it is unlikely to be driven by the market power of any of the individual exporting countries considered in this study. Furthermore, since crude oil is the main export product of Mexico and Indonesia, we also perform tests using the all inclusive commodity price index for these two countries. To ensure that we have real price indices, we deflate both series using the US consumer price index.

Based on the composition of exports, all of our emerging floaters but South Korea could potentially have commodity currencies.[16] Table A2 in the appendix shows that, with the exception of Indonesia and Thailand, the Phillips-Perron unit root test does not allow to reject the null hypothesis for both the commodity price and the real exchange rate indices at the 10% confidence level. At the same time, however, the Kwiatwoski, Phillips, Schmidt and Shin (1992) test does lead to the rejection of the null of no unit root in all cases, except for the real exchange rate of the Philippines. Given the well-known low power of unit root tests for small samples and the fact that 15 of the 18 tests suggest the presence of a unit root at the 10% confidence level we therefore conclude that all series are non-stationary.[17] In first differences, the presence of a unit root is consistently rejected for all variables.

[Table 7 about here]

The first three columns of Table 7 report the results of the Granger-causality and the Phillips-Ouliaris (1990) residual based-tests for cointegration. The Granger-

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structural break in their cointegration relationships over the period when they use the Gregory and Hansen (1996) residual-based cointegration test.

[16] See Cashin, Cespedes and Sahay (2004).

causality tests indicate that there is at least a 90% probability that causality runs from real commodity prices to real exchange rates in the cases of Thailand and Indonesia (irrespective of whether the non-fuel or the all inclusive real commodity price index is used). The Phillips-Ouliaris Zt and Zα tests do not flag cointegration for any of the emerging floaters except for Thailand and Indonesia at the 10% confidence level.

It could well be that the reason why the cointegration tests above do not lead to the rejection of the null of no cointegration in the remaining countries is because these tests do not allow for the possibility of structural breaks. Gregory and Hansen (1996) derived the asymptotic distribution of the test statistics for an alternative test in which the null of no cointegration is checked against the alternative of cointegration in the presence of a possible (single) regime shift of unknown timing. We therefore proceed to check whether the richer Gregory and Hansen test leads to the more frequent rejection of the null of no cointegration. Essentially, if the standard cointegration relation with no structural change is

$$rer_t = \alpha_0 + \beta_0 p_t + \varepsilon_t, \tag{2}$$

where p_t is I(1) and ε_t is I(0), and α_0 and β_0 are coefficients to be estimated, we test whether there is cointegration according to the more flexible relation

$$rer_t = \alpha_0 + \alpha_1 z_t + \beta_0 p_t + \beta_1 p_t z_t + \varepsilon_t, \tag{3}$$

where z_t is an indicator variable that takes the value 1 after the structural break. If α_1 and β_1 are forced to be zero we have a standard cointegration test. If only β_1 is set to zero, we are allowing for a one time parallel shift in the relationship, whereas in the unrestricted case the structural break may involve both a change in the intercept and a change in the slope of the cointegration relationship. The main advantage of this method is that the timing of the structural break does not have to be known a priori. The test involves the computation of the test statistics for each and every possible break point. For computational purposes we used the standard restriction that the break point has to

[17] A similar judgment is made by Chen and Rogoff (2003) among others.

lie within the interval of observations ([.15n],[.85n]), where n is the length of the time series (see Gregory and Hansen (1996)).

The Gregory-Hansen Zt* test statistics for each country are shown in the right half of Table 7. As one might expect, allowing for a unique regime shift does indeed lead to additional rejections of the null of no cointegration: in addition to the cases already mentioned earlier, the Zt* tests now suggest a cointegration relationship between both the Indonesian Rupiah and the Mexican Peso and the real non-fuel commodity price index. Note that the rejection of the null of no cointegration does not depend on whether the parameter β_1 is set to zero or not (i.e. whether the structural change affects the constant only or both the constant and the slope). Moreover, the tests consistently fail to reject the null of no cointegration in the cases of Brazil, South Africa, the Czech Republic and the Philippines. We therefore consider that the currencies of these countries are not commodity currencies as is the case for the Australian or the New Zealand Dollar (see Chen and Rogoff (2003)).

5.2. Cointegration Vectors

Since the above tests flag the existence of cointegrating vectors in the cases of Mexico, Indonesia and Thailand, we now set out to pin these relations down.[18] To estimate the cointegrating relationships in each case we employ the fully modified estimation method of Phillips and Hansen (1990) (FM-OLS). This approach corrects for small sample bias in the OLS estimations. Since Indonesia's main export is crude oil and as our previous analysis has led to the rejection of the null of no cointegration when we used the all inclusive real commodity price index as well, we compute two estimates of the elasticities for this country. Only the non-fuel index is used in the case of the Mexican Peso and the Thai Baht.

[Table 8 about here]

The estimated elasticities for each of the countries with commodity currencies are reported in Table 8. In the case of the Rupiah, the non-fuel based elasticity is .43,

[18] Mexico's main commodity exports are crude oil and copper, Indonesia's crude oil and natural gas, while Thailand's are rice and natural rubber.

whereas the all inclusive elasticity is estimated to be .35. For the case of the Baht we obtain an elasticity of .28, whereas this parameter is .66 for the Mexican Peso. The latter estimate should however be taken with caution for reasons to be explained further below.

[Figure 2 about here]

Table 8 also lists the results of stability checks of the cointegration vectors. Note that the Gregory and Hansen (1996) tests of Table 7 suggested several possible structural breaks in the cointegrating vectors. To test for possible parameter instability in the estimated equations we use all 3 tests suggested by Hansen (1992): SupF, MeanF and LC. As in that paper, the bandwith selection is left to automatically follow Andrews (1991). The stability test statistics make it clear that the null hypothesis of constant parameters cannot be rejected for the Thai Baht and the Indonesian Rupiah vectors, implying that these cointegration relationships are indeed stable. In the case of the Mexican Peso, however, parameter stability is clearly rejected at usual confidence levels. The plots of the evolution of the F-statistics over time are reported in the Appendix for all cases. Whereas in the cases of Thailand and Indonesia the F-statistic remains much below the critical thresholds throughout the sample, the case of Mexico suggests not one but various structural breaks. Figure 2 compares the real exchange rates and real commodity price indices for Thailand and Indonesia, where stable cointegration relationships were identified. Note that, to ease comparison, the real exchange rate scale has been inverted. The Thai Baht seems to be dancing to the tune of commodity prices.

6. Conclusion

The main aim of this paper was to provide answers that are more robust to the Lucas critique to two major questions of international macroeconomics in the context of developing countries. First, what is the extent and the dynamics of exchange rate pass-through to consumer prices and, second, whether the exchange rate movements are related to commodity prices in a stable way, i.e., whether some of these countries have *de facto* commodity currencies.

In contrast to previous literature, we find that pass-throughs to consumer prices have typically been moderate in the 8 major developing countries that have had a floating exchange rate regime for at least a decade - even though there is considerable cross-country variation. We noted that, even though simple correlations suggest a positive association between pass-throughs and inflation rates, the ranking of countries according to their anti-inflation performance clearly does not match with the ranking of countries according to pass-throughs. Pass-through coefficients seem to be related to volatility of the exchange rates and the composition of trade flows in this group of countries. Furthermore, even though higher exchange rate volatilities seem to be associated with higher pass-throughs, we do not find any evidence that price variations have fed into new rounds of exchange rate adjustments.

We also found that for most emerging floaters there is no evidence that the variations in exchange rates are tied to variations in international commodity prices. The noteworthy exceptions are the Mexican Peso, the Indonesian Rupiah, and the Thai Baht. Only for the latter two we find cointegrating vectors that have been stable throughout the period of floating exchange rate regimes.

In our view, further research should try to further disentangle the relation between exchange rate volatility and exchange rate pass-through, perhaps exploring the possibility of asymmetries in price adjustments. Empirically, one could explore the role of trade composition in explaining pass-through in the context of developing countries. Finally, when it comes to explaining exchange rate variations themselves, our results suggest that a closer look at commodity prices may turn out to be a promising venue for some currencies.

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APPENDIX

Data Sources:

IMF IFS:

CPI: National consumer price indices (64..ZF). For Japan the source was the OECD Main Economic Indicators; y: Industrial production index - s.a. (66..CZF). Manufacturing production for Chile; pnfc : price index of non-fuel primary commodities. (00176NFDZF) and pcomm: price index of all primary commodities. (00176ACDZF).

BIS:

Nominal and real effective exchange rate (inverted).

The World Bank (WDI):

Trade openness (taken as the average of the years 1995, 2000 and 2005); share of food and energy products in trade flows and size of GDP in U$ Dollars.

[Table A1 about here]

[Table A2 about here]

Table 1

Nominal Effective Exchange Rate Volatilities

country	ex. rate regime		p(|Δe/e|)>.01	p(|Δe/e|)>.02	p(|Δe/e|)>.05
US	float	since 94/2	0.379	0.107	0.006
Eurozone	float	since 99/2	0.462	0.154	0
Japan	float	since 94/2	0.605	0.345	0.062
Brazil	float	since 99/1	0.805	0.559	0.170
South Korea	float	since 97/12	0.550	0.260	0.084
Mexico	float	since 94/12	0.539	0.281	0.048
Indonesia	float	since 97/8	0.681	0.496	0.315
South Africa	float	since 95/3	0.732	0.476	0.098
Thailand	float	since 97/7	0.471	0.169	0.066
Czech Rep.	float	since 97/5	0.449	0.203	0.016
Philippines	float	since 97/12	0.585	0.244	0.022
China	peg	since 94/2	0.362	0.085	0
Brazil	peg	94/7 - 98/12	0.370	0.148	0.037
Russia	peg	since 99/12	0.355	0.121	0
India	peg	since 94/2	0.441	0.141	0.006
South Korea	peg	94/2 - 97/11	0.131	0.022	0.022

Note: Based on nominal effective exchange rates published by the BIS. Numbers represent fraction of time that variation exceeded threshold.

Figure 1
Real and Nominal Effective Exchange Rates

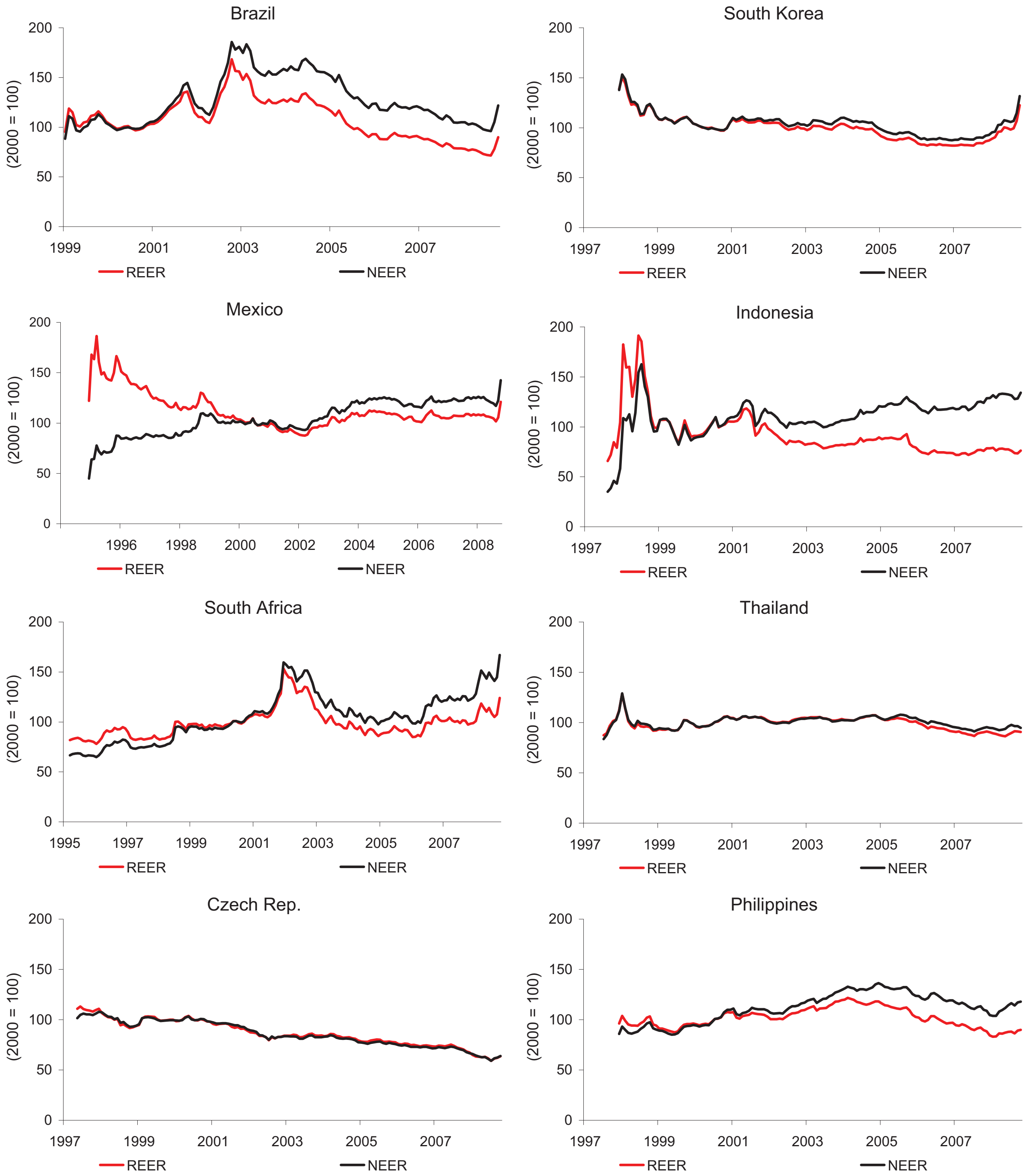


Table 2

Estimated Pass-Throughs - Bivariate VARs

	horizon			
	t=3	t=6	t=9	t=12
Brazil	0.069	0.203	0.302	0.321
	0.022	0.047	0.078	0.108
South Korea	0.071	0.161	0.154	0.206
	0.032	0.056	0.074	0.096
Mexico	0.033	0.134	0.125	0.133
	0.026	0.051	0.065	0.076
Indonesia	0.190	0.314	0.515	0.599
	0.022	0.050	0.082	0.122
South Africa	0.083	0.144	0.180	0.247
	0.016	0.035	0.052	0.065
Thailand	0.042	0.086	0.175	0.159
	0.028	0.052	0.064	0.077
Czech Rep.	0.103	0.214	0.350	0.512
	0.047	0.081	0.110	0.142
Philippines	0.048	0.058	0.061	0.093
	0.049	0.086	0.114	0.137
Average EM	0.081	0.180	0.254	0.284
Median EM	0.069	0.161	0.175	0.206

Notes:

1) VARs between Δe and Δp, where p is the log of the CPI.

2) Using 12 months (lags).

Table 3

Granger Causality Tests

	obs.	causal direction $\Delta e \rightarrow \Delta cpi$ (prob.)	$\Delta cpi \rightarrow \Delta e$ (prob.)
Brazil	105	0.999	0.746
South Korea	118	0.955	0.485
Mexico	154	0.970	0.142
Indonesia	135	1.000	0.465
South Africa	162	0.999	0.111
Thailand	136	0.944	0.767
Czech Rep.	138	0.958	0.871
Philippines	131	0.701	0.135

Note: Using 12 lags with monthly data.

Table 4

Estimated Pass-Throughs - VARs with output

	horizon			
	t=3	t=6	t=9	t=12
Brazil	0.065	0.195	0.291	0.325
	0.024	0.051	0.085	0.120
South Korea	0.056	0.172	0.154	0.224
	0.036	0.062	0.087	0.117
Mexico	0.034	0.154	0.158	0.165
	0.026	0.053	0.069	0.080
South Africa	0.082	0.140	0.179	0.245
	0.017	0.037	0.055	0.069
Czech Rep.	0.105	0.233	0.370	0.542
	0.050	0.089	0.122	0.148
Philippines	0.044	0.072	0.093	0.097
	0.052	0.093	0.133	0.163
Average EM	0.052	0.174	0.201	0.238
Median EM	0.056	0.172	0.158	0.224

Notes:

1) VARs between Δy, Δe and Δp, where y is the log of industrial production and p is the log of the CPI.

2) Using 12 months (lags).

Table 5
Potential determinants of pass-through

	6 months pass-through	12 months pass-through	inflation (average)	inflation (median)	inflation (variance)	NEER volatility	trade openess	food&energy trade	GDP 2005 ($ bn)
Brazil	0.203	0.321	0.073	0.060	0.004	0.559	0.172	0.256	882
South Korea	0.161	0.206	0.036	0.035	0.004	0.260	0.612	0.161	791
Mexico	0.134	0.133	0.127	0.076	0.031	0.281	0.561	0.134	768
Indonesia	0.314	0.599	0.172	0.074	0.147	0.728	0.541	0.323	287
South Africa	0.144	0.247	0.066	0.057	0.005	0.650	0.428	0.175	242
Thailand	0.086	0.159	0.033	0.025	0.005	0.359	0.980	0.171	176
Czech Rep.	0.214	0.512	0.029	-0.048	0.179	0.452	1.051	0.108	125
Philippines	0.058	0.093	0.059	0.044	0.006	0.398	0.850	0.141	99

Notes:
1) Coefficients in the first two columns are taken from Table 2.
2) Inflation figures refer to yearly CPI rates.
3) NEER volatility is the proportion of months in which variation exceeded 2% (from Table 1).
4) Openness is defined as the sum of export and import values divided by GDP.
5) Food and energy represents the share of trade flows that is made up by food and energy products.

Table 6
Correlations

	π (avrg)	π (median)	π (var)	NEER volatility	trade	food&energy	GDP
Simple correlations							
6 months PT	0.737	0.646	0.812	0.599	-0.569	0.869	0.010
12 months PT	0.673	0.472	0.865	0.697	-0.352	0.963	-0.306
Spearman correlations							
6 months PT	-0.031	-0.188	-0.031	0.063	-0.656	0.125	0.188
12 months PT	-0.125	-0.313	-0.063	0.188	-0.656	0.281	0.000

Note: π refers to annual consumer price inflation.

Table 7
Cointegration and Granger-Causality tests

real exchange rate and real price of non-fuel commodities

	N	Granger causality pnfc → e (probab.)	Phillips - Ouliaris 1) Zt	Phillips - Ouliaris 1) Zα	Gregory Hansen Zt* test constant	sugg. break	constant & slope	sugg. break
Brazil	118	0.885	-2.879	-8.481	-3.71	-	-3.69	-
Mexico	165	0.689	-2.297	-5.778	-4.93**	2005:12	-5.65***	2002:12
Indonesia	135	0.905	-2.715	-4.217	-4.89**	1999:03	-5.19**	2002:06
South Africa	162	0.772	-1.703	-2.923	-3.37	-	-3.04	-
Thailand	136	0.999	-3.266*	-12.12*	-4.53*	2006:04	-5.25**	2000:07
Czech Rep	138	0.379	-2.319	-6.348	-2.97	-	-2.45	-
Philippines	131	0.801	-1.542	-3.813	-3.42	-	-3.58	-

real exchange rate and real price of all *commodities*

	N	Granger causality pcom → e (probab.)	Phillips - Ouliaris Zt	Phillips - Ouliaris Zα	Gregory Hansen Zt* test constant	sugg. break	constant & slope	sugg. break
Mexico	165	0.049	-1.586	-0.878	-4.03	-	-4.35	-
Indonesia	135	0.922	-3.698***	-9.909	-4.87**	1999:03	-4.73*	2001:10

*, **, *** denote statistical significance at the 10, 5 and 1% confidence level respectively.

1) Critical values obtained for these tests were taken from MacKinnon (1991) [P-O Zt], Huag (1992) [P-O Zα].

Table 8
Cointegration relationships and Hansen parameter stability tests

Thailand

non-fuel commodittties

	rer_t	=	5.889	$-.275{*}pnfc_t$	N=136
			(.497)	(.105)	
Bandwith			4.068		
SupF			1.983	p > 0.20	
MeanF			1.129	p > 0.20	
LC			0.111	p > 0.20	

Indonesia

non-fuel commodittties

	rer_t	=	6.507	$-.425{*}pnfc_t$	N=135
			(1.791)	(.227)	
Bandwith			3.145		
SupF			2.649	p > 0.20	
MeanF			1.100	p > 0.20	
LC			0.116	p > 0.20	

all commodittties

	rer_t	=	6.187	$-.354{*}pcomm_t$	N=135
			(.667)	(.140)	
Bandwith			3.096		
SupF			2.658	p > 0.20	
MeanF			1.253	p > 0.20	
LC			0.163	p > 0.20	

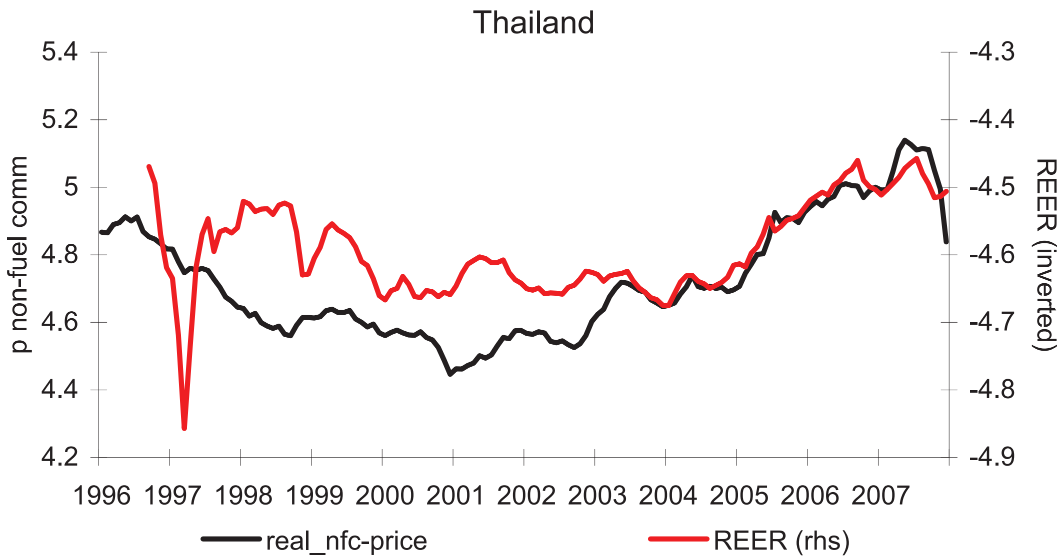
Mexico

non-fuel commodittties

	rer_t	=	1.576	$-.659{*}pnfc_t$	N=167
			(.894)	(.188)	
Bandwith			1.757		
SupF			15.51**	p=0.011	
MeanF			7.48***	p=0.010	
LC			1.293***	p=0.010	

Critical values obtained from Hansen (1992).

Figure 2
Real Effective Exchange Rates and Real Commodity Prices



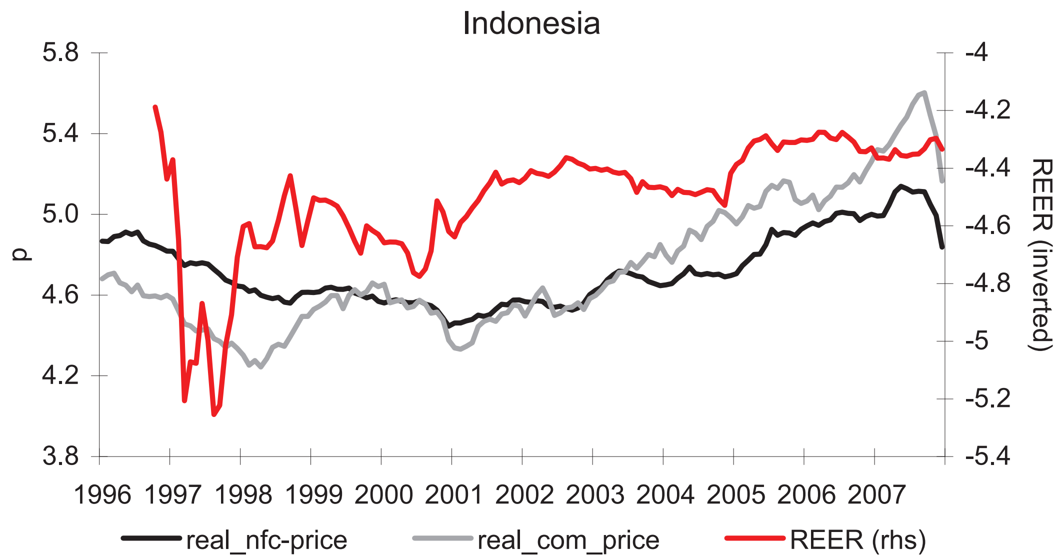


Table A1

Variance Decomposition - Bivariate VARs

		[Δe and Δp] ordering		[Δp and Δe] ordering	
		t=6	t=12	t=6	t=12
Brazil	*e* as part of *cpi*	0.400	0.486	0.428	0.511
	cpi as part of *e*	0.050	0.054	0.082	0.094
South Korea	*e* as part of *cpi*	0.197	0.211	0.208	0.220
	cpi as part of *e*	0.049	0.066	0.046	0.070
Mexico	*e* as part of *cpi*	0.210	0.220	0.174	0.183
	cpi as part of *e*	0.015	0.031	0.039	0.060
Indonesia	*e* as part of *cpi*	0.584	0.615	0.563	0.643
	cpi as part of *e*	0.053	0.088	0.093	0.115
South Africa	*e* as part of *cpi*	0.246	0.293	0.239	0.290
	cpi as part of *e*	0.019	0.035	0.018	0.034
Thailand	*e* as part of *cpi*	0.121	0.186	0.086	0.154
	cpi as part of *e*	0.092	0.135	0.129	0.175
Czech Rep.	*e* as part of *cpi*	0.105	0.189	0.105	0.190
	cpi as part of *e*	0.014	0.081	0.014	0.081
Philippines	*e* as part of *cpi*	0.043	0.074	0.024	0.084
	cpi as part of *e*	0.024	0.084	0.044	0.075

Notes:

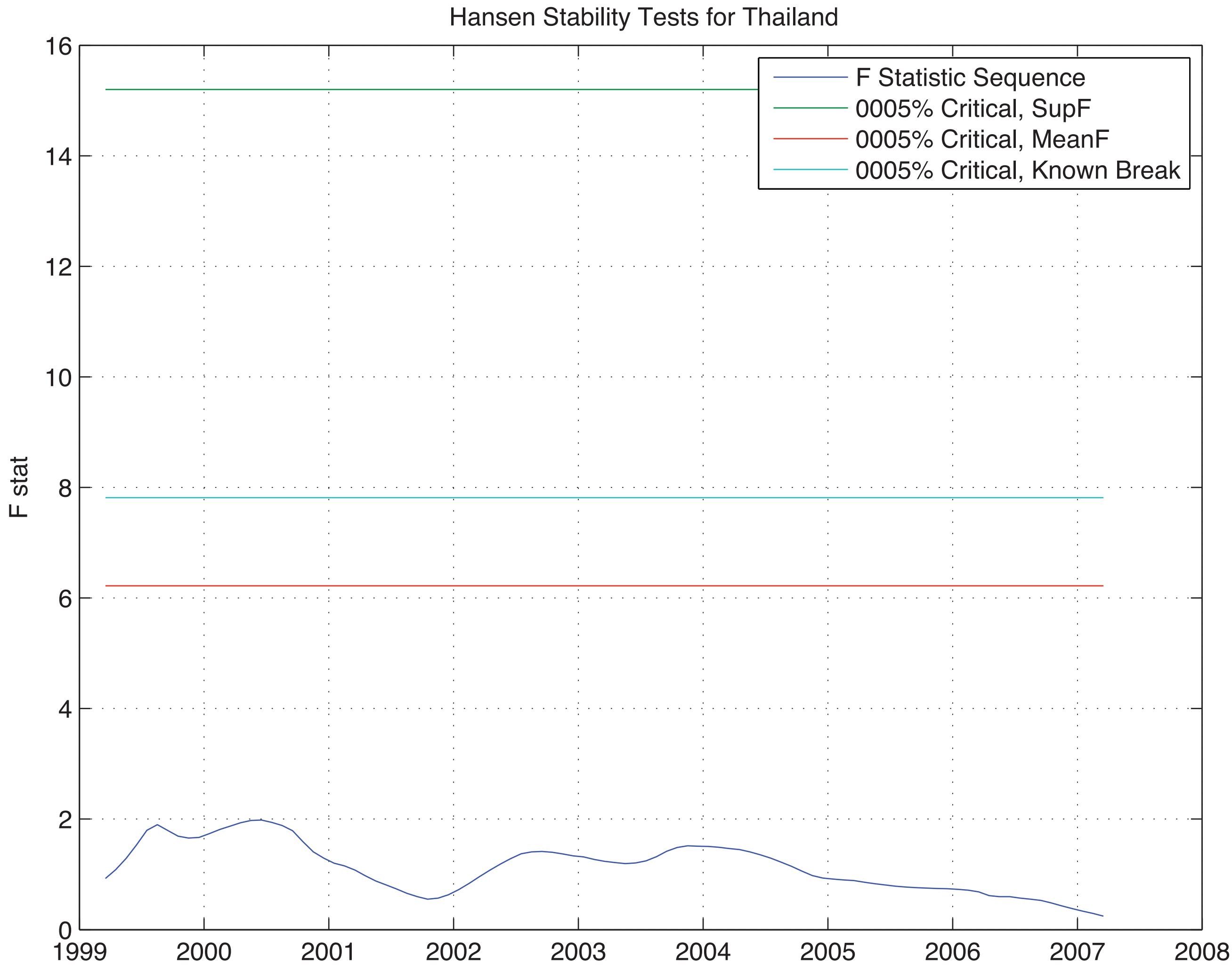
1) VARs between Δe and Δp, where p is the log of the CPI.

2) Using 12 months (lags).

Table A2
Unit Root Tests

	N	Phillips - Perron		Kwiatwoski *et. al.*	
		level	diff	level	diff
real exchange rate:					
Brazil	118	-2.054	-6.205***	0.619**	0.247
Mexico	165	-1.921	-10.052***	0.803***	0.110
Indonesia	135	-2.832*	-9.301***	0.957***	0.135
South Africa	164	-1.886	-9.360***	0.409*	0.079
Thailand	136	-4.802***	-9.665***	0.440*	0.295
Czech Rep	138	-0.069	-9.957***	1.417***	0.077
Philippines	131	-1.371	-8.537***	0.300	0.290
real prices:					
non-fuel commodities	165	-1.301	-7.054***	0.404*	0.182
all commodities	165	-0.734	-8.971***	1.010***	0.148

Bandwith selection method: Newey-West using Bartlett kernel



Hansen Stability Tests for Thailand
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0005% Critical, MeanF
0005% Critical, Known Break
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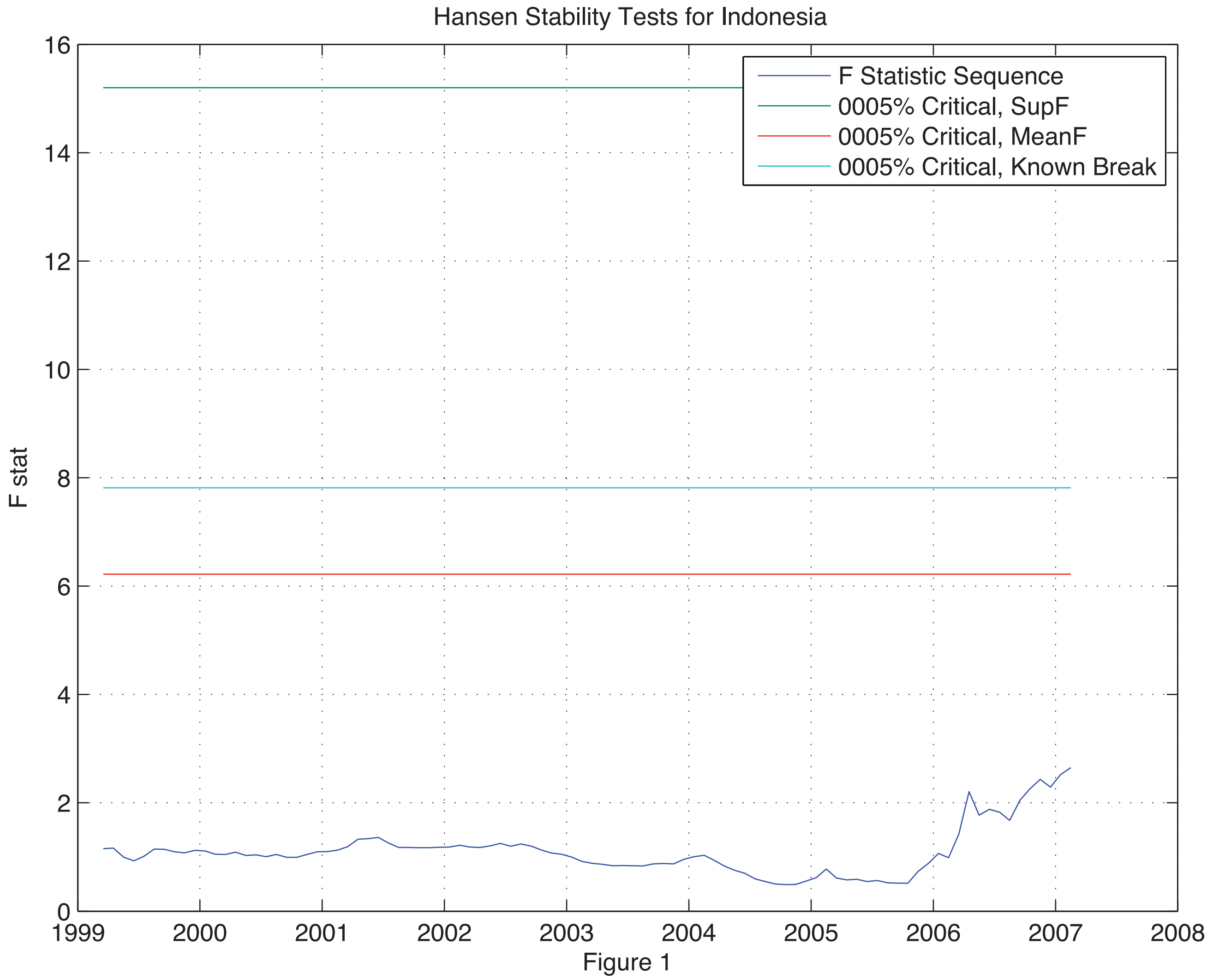
Figure 1


Hansen Stability Tests for Indonesia
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0005% Critical, MeanF
0005% Critical, Known Break
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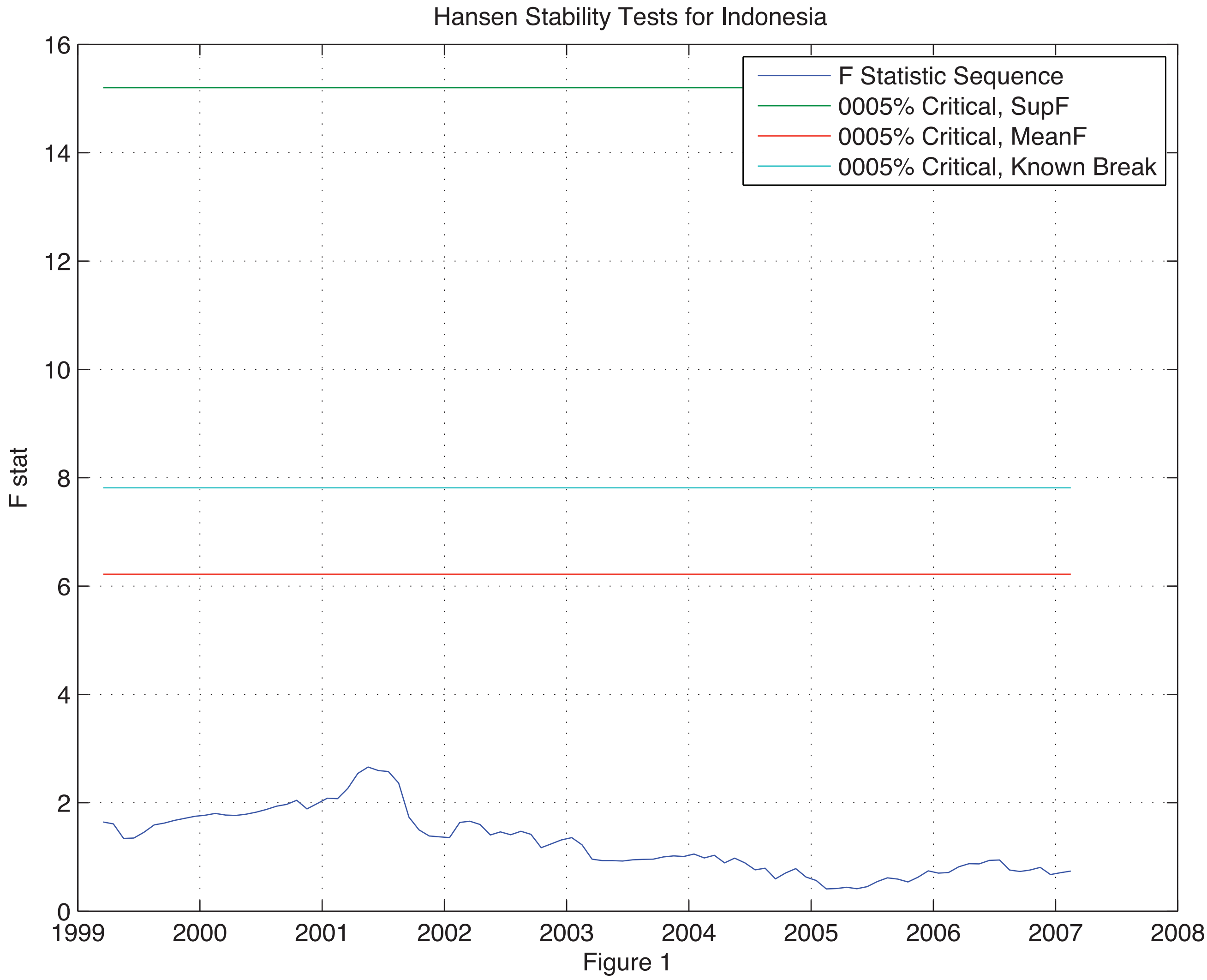
Hansen Stability Tests for Indonesia
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Figure 1

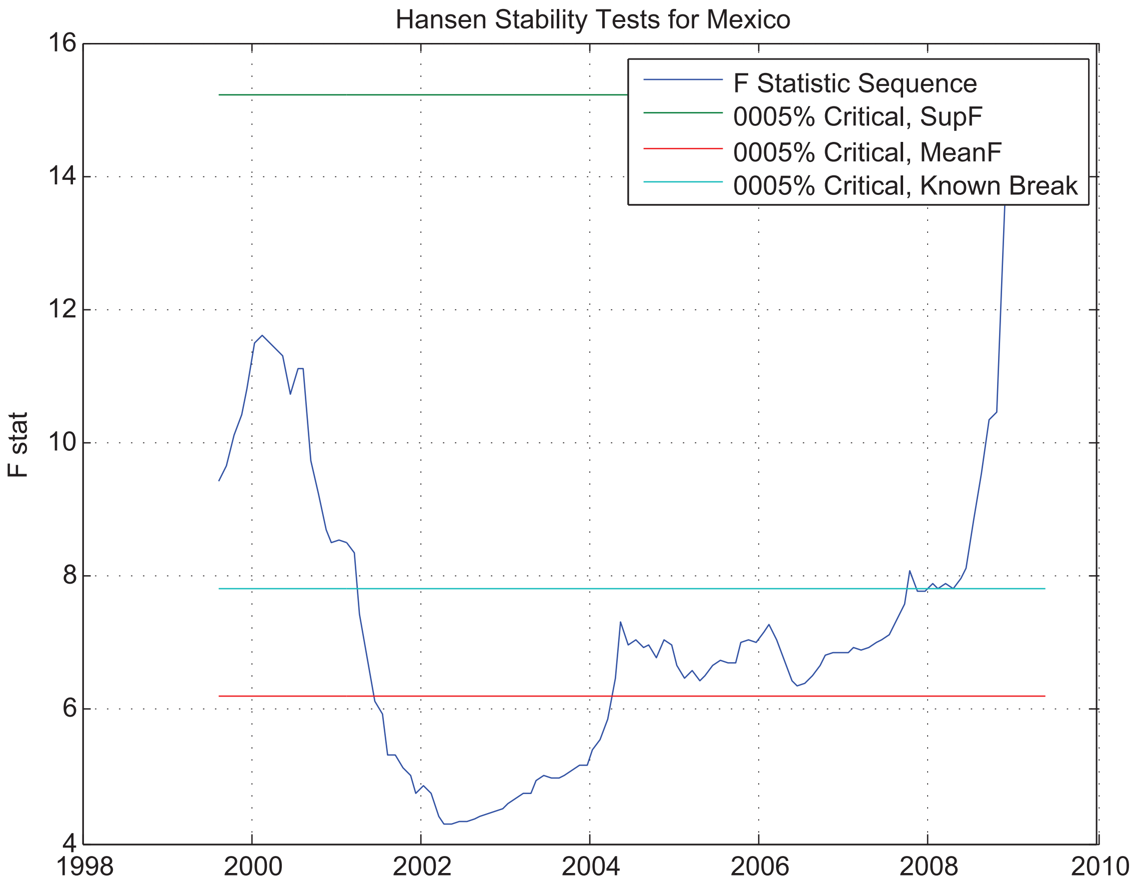
Hansen Stability Tests for Mexico
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Figure 1

Banco Central do Brasil

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