Revisiting the Decline in the Exchange Rate Pass-Through: Further Evidence from Developing Countries

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Abstract

This paper revisits the Taylor (2000) proposition which stipulates that exchange rate pass-through in industrial countries has known a considerable fall during the 1990s caused by a low inflation environment. By adopting an empirical methodology based on structural change and cointegration tests, we show that the same phenomenon occurs for several developing countries. The main explanation for this finding is that these countries experienced a significant inflation fall during the 1990s induced by a shift in their monetary policy caused by exchange rate targeting or inflation targeting.

Key-words: Exchange rate pass-through, developing countries, structural changes, cointegration tests

JEL classification: C20, F21, F30

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1 Introduction

The extent to which exchange rate movements are passed-through to import prices is an important issue in international economics and in the debate about appropriate monetary and exchange rate policies. Recently, theoretical and empirical researches on the exchange rate pass-through to aggregate import prices have attracted a renewed attention. In this context, the focus has shifted to the stability of exchange rate pass-through over time. Indeed, some empirical researches show that the degree of exchange rate pass-through has declined in the 1990s. In this context, Taylor (2000) argues that the recently observed declines in the pass-through to aggregate prices are the result of low inflation environments. His argument is that it is increasingly difficult for firms to fully pass on exchange rate movements to their export prices in the context of recent economic environment characterized by low and stable inflation. The Taylor point of view is based on the link between pass-through and inflation. The pass-through depends on the policy regime: a credible low inflation regime will automatically achieve a low exchange rate pass-through.

To verify the Taylor hypothesis, most of the empirical studies that examine a recent decline in the pass-through coefficient, have focused on import or consumer prices into industrialized countries, rather than into developing countries. For example, Otani et al. (2003) find a decline in pass-through for imports into Japan, which they attribute to a decline in global inflation. Campa and Goldberg (2004) find a decline in the pass-through coefficient in the 1990s, which they attribute to changing commodity composition more than to a less inflationary environment, but their data set again consists solely of industrialized countries. The apparent decline in the pass-through coefficient in developing countries in the 1990s has not been extensively documented and explained. More precisely, there are a few studies as that of Choudri and Hakura (2001) who investigate a sample of 71 countries, including developing countries. Borensztein and De Gregorio (1999), and Goldfajn and Werlang (2001) study the low pass-through of recent large devaluations in developing countries. Saiki (2004) includes two developing countries in his study of whether a switch in monetary regime to inflation targeting is associated with a fall in the pass-through coefficient. But, all these studies focus on the CPI and don't consider the import prices.

The main purpose of this paper is to revisit the Taylor (2000) proposition to a broad sample including some developing countries in order to examine the reported decline in their pass-through coefficients and possible explanations for it. To that effect, we propose an empirical approach based on structural breaks detection. Indeed, we adopt some selection procedures for the number of breaks and their locations based on tests of structural changes and information criterion. We also use Gregory and Hansen's (1996) tests of cointegration allowing one break at unknown time in order to test for possible exchange rate pass-through decline. The use of these tests is motivated by the fact that there are cointegration relationships in our pass-through equation.[1]

This paper is structured as follows. Section 3, briefly reviews some empirical research on exchange rate pass-through decline. In section 3, we present our pass-through equation. Section 4 investigates our empirical methodology and presents the results. Section 5 concludes

[1] Based on the micro-foundations of firms pricing, we consider exchange rate pass-through as an equilibrium profit maximizing strategy for firms (a long-run phenomenon) rather than a short-run phenomenon.

the paper. Appendix A provides a description of the data used in the empirical illustration and Appendix B presents the empirical results.

2 A review of empirical literature

Recently, there has been a growing interest in examining the relationship of the exchange rate pass-through with monetary policy behavior and inflation environment. We can divide these different works in two categories. The first one is based on finding significant subsample dummy variables, when these are applied to the coefficient on the exchange rate in a univariate pass-through equation such as Gagnon and Ihrig (2001), Choudri and Hakura (2001) and Devereux and Yetman (2001).

Gagnon and Ihrig (2001) study the link between consumer prices and monetary policy in a sample of 20 industrialized countries over the period from 1971 to 2000. Using a cross-sectional approach; they test whether the pass-through declines in each country in the sample following a change in the inflation regime. One regime change is identified for each country using a combination of causal inspection of the data and judgment. For each country in the sample, pass-through equations are then estimated on two subsamples. In all cases, the pass-through coefficients are smaller in the second subsample, which they interpret as evidence that exchange rate pass-through has declined in industrialized countries, and this decline is attributable to the change in the inflation regime.

Choudri and Hakura (2001) test the Taylor (2000) hypothesis for a sample of 71 countries, including developing countries, by using a Dynamic General Equilibrium (DGE) model with imperfect competition and staggered contract. In their model, a low inflation regime reduces exchange rate pass-through because the pass-through reflects the expected effect of monetary shocks on current and future costs. They find strong evidence that the relation between exchange rate pass-through and the average inflation rate is positive and significant across regimes. Their results are based on comparisons of regimes across countries and periods.

Devereux and Yetman (2001) study the link between exchange rate pass-through and monetary policy in the context of a DGE model. In their theoretical model, pass-through is determined by the frequency of prices changes of importing firms (this frequency is a function of the monetary policy regime). They stipulate that firms in countries where monetary policy is more credible (i.e. the mean inflation is lower) will tend to change their prices less frequently, and thus they lead to a lower exchange rate pas-through. In addition to their theoretical contribution, Devereux and Yetman (2001) investigate the role of inflation variables in accounting for cross-country differences in exchange rate pass-through in a large sample of countries. Their approach is based on estimating a first-regression for each country in their sample in order to obtain an average pass-through elasticity estimate over the time period (30 years).

The second category is based on the structural break tests, in particular the Campa and Goldberg (2004), and Baillu and Fujii (2004) studies. Campa and Goldberg (2004) performed two types of structural change tests on the pass-through elasticities. Firstly, they assume an exogenously imposed break point in the middle of the sample and performed the Chow tests for parameter stability. They find that exchange rate pass-through coefficients have been

declined in both short-run and long-run for many OECD countries. Secondly, they use the methods proposed by Andrews (1993), and Andrews and Ploberger (1994). They show that pass-through coefficients have declined in the short-run but not in the long-run.

In their empirical study, Baillu and Fujii (2004) investigate the question of whether a transition to a low-inflation environment induced by a shift in monetary policy results in a decline in the degree of pass-through of exchange rate movements to consumers prices. As mentioned above, this empirical work focuses on the identification of changes in the inflation environment and uses a panel data approach for 11 industrialized countries over the period from 1977 to 2001. Their results suggest that pass-through to import producer and consumer price inflation declined following the inflation stabilization that occurred in many industrialized countries in the early 1990s.

Note that the Devereux and Yetman (2001) empirical analysis is different to the Baillu and Fujii (2004) study. The latter imposes a dummy based on a known break, while the former incorporates dummies based on estimated break dates.

3 Exchange rate pass-through equation

Exchange rate pass-through empirical studies were interested in the extent to which exchange rate movements are transmitted to traded goods prices, versus absorbed in producer profit margins or markups. According to Goldberg and Knetter (1997), exchange rate pass-through is defined as the percentage change in the local currency import prices resulting from a one percent change in the exchange rate between the exporting and importing countries. The exchange rate pass-through into import prices studies are empirically implemented as a statistical relationship of the elasticity of import prices to exchange rates. Testing this relationship is based on the following equation:

$$\Delta p_t = \gamma \Delta e_t + \varepsilon_t, \tag{1}$$

where p_t is the natural logarithm of import price, e_t is the nominal exchange rate, ε_t is an error term, and γ is the exchange rate pass-through coefficient. The extent of exchange rate pass-through coefficient is based on the value of γ. A one to one response of import prices to exchange rate is known as a complete exchange rate pass-through and $\gamma = 1$, while less than exchange rate pass-through coefficient ($\gamma < 1$) is known as partial or incomplete exchange rate pass-through. However, Campa and Goldberg (2004) criticize this specification because it only represents a non-structural statistical relationship and lacks an economic interpretation. They argue that a correct specification should include, additionally, controls to capture exporter's costs associated with local inputs and demand conditions in the destination country. Recent empirical studies[2] on exchange rate pass-through into import prices use an approach based on micro-foundations of pricing behavior by exporting firms.

In this paper, the equation used to estimate the degree of the exchange rate pass-through into import prices is similar to the equation found in the literature in this area (Hooper and Mann (1989), Goldberg and Knetter (1997), Campa and Goldberg (2003), and Barhoumi (2005, 2006)). We consider that exchange rate pass-through into import price is determined

[2] Campa and Goldberg (2003), and Eiji Fuji (2004).

by a combination of nominal effective exchange rate, the price of the competing domestic product, the exporters costs and domestic demand conditions. Additionally, we adopt the assumption of imperfectly competitive market structures, we concentrate on the micro-foundations of firms' pricing and we regard exchange rate pass-through as an equilibrium profit maximizing strategy for firms (a long-run phenomenon) rather than a short-run phenomenon caused by the contractual stickiness of prices. Our exchange rate pass-through equation is the following:[3]

$$pm_t = \mu + \beta_1 e_t + \beta_2 c_t^* + \beta_3 p_t + \beta_4 y_t, \tag{2}$$

where all the variables are expressed in logarithm, pm_t is the import unit value in domestic currency, e_t is the nominal effective exchange rate, c_t^* is the marginal cost of production of foreign firm, p_t is the Producer Price Index (PPI), and y_t is the Industrial Price Index (IPI). Note that the coefficient β_1 is the long-run exchange rate pass-through.

4 Empirical methodology

The main problem in empirical studies on developing countries is data availability. Because of the difficulty to find some variables such as the nominal effective exchange rate, we only consider a sample of 8 developing countries, namely Bolivia, Botswana, Chile, Colombia, Indonesia, Singapore, Uruguay and Venezuela. The data are quarterly and span the period 1980:2-2003:4 (yielding 95 observations). They are obtained from International Financial Statistics.

A formal investigation of the Taylor's hypothesis requires a comparison of pass-through estimates under inflation environment, where the shift can result from an inflation decline. In order to analyze the decline of exchange rate pass-through in some developing countries, we use the approach suggested by Gregory and Hansen (1996) which has the advantage of testing for cointegration relationship between the variables of our equation in the presence of structural break, and it is suitable in the single equation case. Based on the Taylor (2000) proposition that the decline of pass-through coefficient is caused by a fall in inflation, we employ two types of structural break tests. Firstly, we analyze the inflation decline in some developing countries by using the Bai and Perron (1998) tests. Then, we investigate the decline of exchange rate pass-through by using Gregory and Hansen (1996) tests. Our empirical approach consists in comparing the break dates detected for the inflation series and nominal effective exchange rate series especially those identified near the 1990s. Once we have determined the dates, we investigate the sign of their coefficient and see whether it corresponds to a decline of both inflation and pass-through.

4.1 Structural break approach without cointegration

We review the structural change approach of Bai and Perron (1998, 2003) and present a selection procedure based on an information criterion. These are designed, in this paper, to detect regime-shifts in the considered series.

[3] Details of the exchange rate pass-through equation are available in Appendix A.

We first present the structural change model with m breaks, $(T_1, \ldots, T_m)$:

$$y_t = z_t' \delta_j + u_t, \tag{3}$$

where $t = T_{j-1} + 1, \ldots, T_j$ for $j = 1, \ldots, m+1$, $T_0 = 0$ and $T_{m+1} = T$. y_t is the observed dependent variable, $z_t \in \mathbb{R}^q$ is the vector of regressors, δ_j is the corresponding vector of regression coefficients with $\delta_i \neq \delta_{i+1}$ $(1 \leq i \leq m)$, and u_t is the error term. We observe that all the coefficients are subject to change, and thus, the model is described as a pure structural change model.[4] Let $\delta = \left(\delta_1', \delta_2', \ldots, \delta_{m+1}'\right)'$.

The estimation method considered is that based on the ordinary least-squares (OLS) method proposed by Bai and Perron (1998). The method first consists in estimating the regression coefficients δ_j by minimizing the sum of squared residuals $\sum_{i=1}^{m+1} \sum_{t=T_{i-1}+1}^{T_i} (y_t - z_t' \delta_i)^2$. Once the estimate $\hat{\delta}(T_1, \ldots, T_m)$ is obtained, we substitute it in the objective function and denote the resulting sum of squared residuals as $S_T(T_1, \ldots, T_m)$. The estimated break dates $\left(\hat{T}_1, \ldots, \hat{T}_m\right)$ are then determined by minimizing $S_T(T_1, \ldots, T_m)$ over all partitions $(T_1, \ldots, T_m)$ such that $T_i - T_{i-1} \geq [\varepsilon T]$,[5] where ε is an arbitrary small positive number. Finally, the estimated regression coefficients are such that $\hat{\delta} = \hat{\delta}\left(\hat{T}_1, \ldots, \hat{T}_m\right)$. In our empirical computations, we use the efficient algorithm developed in Bai and Perron (2003), based on the principle of dynamic programming, to estimate the unknown parameters.

Bai and Perron (2003) propose a test-based selection procedure (TBSP) to estimate the number of breaks. Indeed, they suggest to first look at the results of two parameter instability tests, UD max or WD max,[6] to see if at least a structural break exists. The number of breaks is then determined based upon a sequential examination of a test, $\sup F_T(l+1|l)$.[7] We then choose m break dates such that the test $\sup F_T(l+1|l)$ is not significant for any $l \geq m$. Bai and Perron (2003) conclude that this method leads to the best results and is recommended for empirical applications.

We can also use the Bayesian information criterion, BIC, proposed by Yao (1988) to detect the number of structural breaks in the data.[8] Unlike this criterion, the test-based selection procedure takes into account the effect of serial correlation in the errors and heterogenous variances across regimes.[9] Note that the existence of breaks in the variance could be exploited to increase the precision of the break date estimates (e.g., Bai and Perron, 2003).

[4] Note that the constraint of absence of change must be imposed when it is known that some coefficients of the model are not time-varying. Thus, the model is described as a partial structural change model.

[5] From Bai and Perron (2003), if the estimation is the sole concern for the study, then the minimal number of observations in each regime $[\varepsilon T]$ can be set to any value greater than q.

[6] Note that these tests are called *double maximum tests*. They allow us to test the null hypothesis of no structural break versus an unknown number of changes given a maximum permitted number of breaks M for m. The significance of these tests does not provide enough information about the exact number of breaks but means that one break is at least present (for more details, see Bai and Perron (1998)).

[7] Note that this test allows to test the null hypothesis of l breaks against the alternative that an additional break date exists. It is based on the difference between the sum of squared residuals obtained with l breaks and that obtained with $(l+1)$ breaks (for more details, see Bai and Perron (1998)).

[8] The estimated number of break points is obtained by minimizing the criterion BIC given an upper bound M for m.

[9] For more details on the different versions of the structural change tests that can be obtained depending on the assumptions made with respect to the distributions of the errors across regimes, the readers are referred to Bai and Perron (2003).

Jouini and Boutahar (2005) use the above-mentioned selection methods to explore the empirical evidence of the instability by uncovering structural breaks in some U.S. time series. To that effect, they pursue a methodology composed of different steps and propose a modelling strategy to implement it. Their results indicate that the time series relations have been altered by various important facts and international economic events such as the two Oil-Price Shocks and changes in the International Monetary System.

4.2 Structural break approach with cointegration

We now present an alternative approach for the study, namely that of Gregory and Hansen (1996). This approach consists in testing the null hypothesis of no cointegration against the alternative of cointegration with one unknown break date.[10] To that effect, Gregory and Hansen (1996) consider the following regime-shift model:

$$y_t = \mu_1 + \mu_2 \varphi_{t\tau} + \beta_1' x_t + \beta_2' x_t \varphi_{t\tau} + e_t, \qquad t = 1, \ldots, T, \tag{4}$$

where y_t is real-valued, x_t is a vector of regressors, e_t is an $I(0)$ process, and $\varphi_{t\tau}$ is a dummy variable defined as follows:

$$\varphi_{t\tau} = \begin{cases} 0, & \text{if} \quad t \leq [\tau T], \\ 1, & \text{if} \quad t > [\tau T], \end{cases} \tag{5}$$

with $\tau \in (0, 1)$ the timing of the break date ($[\cdot]$ denotes integer part of argument). From this model, we see that the structural change affects the intercept μ and the slope vector β. This permits the equilibrium relation to rotate as well as shift parallel.

In the context of standard cointegration, the methods used to test the null hypothesis of no cointegration are residual-based. The same approach could be used for the model (4), if the break date were known a priori. Gregory and Hansen (1996) present three test statistics Z_α^*, Z_t^* and ADF^*, that don't require to know the timing of the break, by adopting a similar solution to that of Banerjee, Lumsdaine, and Stock (1992), and Zivot and Andrews (1992). Indeed, they compute the cointegration test statistic for each possible regime shift τ, and take the smallest value across all possible break points.[11]

4.3 Inflation environment

In order to attribute the declining pass-through to a lower-inflation environment as in Taylor (2001), one needs to document the extent to which inflation did indeed decline in our sample. In their empirical analysis, Tytell and Wei (2003) find that the average inflation declined in developing countries from 25% in the second half of the 1970s to 13% in the second half of the 1990s. Figure 1 depicts the aggregate inflation in developing countries during the period 1980:2-2003:4. We observe that inflation decreased in particular during the 1990s. To be sure that the considered series is characterized by the presence of breaks in its structure, we apply

[10] This is because in some empirical applications, it may be desirable to think of cointegration as holding over some period of time, and then shifting to a new long-run relationship.

[11] Note that $\tau \in \Upsilon$, where $\Upsilon = (0.15, 0.85)$ seems a reasonable suggestion according to Gregory and Hansen (1996).

the structural change techniques defined above; the test-based selection procedure and the information criterion. To that effect, we first consider that $z_t = 1$ in the model (3), which implies that our approach is directly oriented at the issue of looking for multiple structural changes in the mean of the series. We then look for multiple breaks in the level and the persistence of the series, that is, $z_t = (1, y_{t-1})'$ in the model (3). For the two cases, we impose the minimum structure on the data by allowing for different distributions of both the regressors and the errors in the different segments since the graph of the series shows different variability in different periods.[12] Hence, we investigate the stability of the process allowing for different variances for the residuals across segments.[13] The maximum permitted number of breaks is set at $M = 5$ and we use a trimming $\varepsilon = 0.15$ which means that the minimal number of observations in each segment is set at $[\varepsilon T]$, with T the sample size. From Table 1, we observe that aggregate inflation in developing countries has declined in two periods 1983:1 and 1992:4. For the individual inflation series (see, Tables 2.1 and 2.2), the structural change procedures[14] show that the bulk of the break dates occur in the middle of the 1980s and during the 1990s.

4.4 Pass-through decline

We test for stationarity by using the tests ADF and KPSS. The results (not reported here) indicate that, in most cases, the tests have argued unanimously on the order of integration to be 1. Tacking into account the presence of structural breaks, we cannot reject the null hypothesis of unit root for all variables by employing the Perron (1990), and Banerjee, Lumsdaine and Stock (1992) tests.

In order to verify the Taylor (2000) proposition, we apply the cointegration tests Z_α^*, Z_t^* and ADF^* to the long-run exchange rate pass-through relation represented by the model (2) by considering the case where all the coefficients change over time as shown by the model (4). By inspecting the results of Table 3, we note that with the exception of Colombia and the Venezuela, we reject the null of no cointegration for the other countries. Besides, the break dates associated to the nominal effective exchange rate are between the middle of the 80s and during the 90s. For these countries, the results given in Table 4.1 and Table 4.2 indicate that the inflation decline is occurred from 1983:2 for Bolivia to 1998:1 for Indonesia.[15] We now investigate whether the observed inflation decreases are associated with a decline in the pass-through coefficient in the 1990s. Indeed, the results presented in Table 5 confirm this fact. This decline can be explained by a change in the monetary policy of the 6 countries. More precisely, in order to reduce their inflation rate, these countries started to adopt different monetary policies such as exchange rate targeting and/or inflation targeting. Table 6 provides

[12] Note that different versions of the tests can be obtained depending on the assumptions made with respect to the distributions of the regressors and the errors across subsamples. These relate to different specifications in the construction of the estimate of the limiting covariance matrix $V\left(\hat{\delta}\right)$ given by (**??**) (Bai and Perron, 2003a, 2004).

[13] Note that the existence of breaks in the variance could be exploited to increase the precision of the break date estimates (e.g., Bai and Perron, 2003a).

[14] Note that the conditions imposed on the distributions of the regressors and errors across segments to determine the breaks in the aggregate inflation series are applied for the individual inflation series.

[15] Note that the selected dates reported on Table 4 are the sole dates associated with an inflation decline.

further details about the monetary shift in these countries. In the following, we provide the explanations for the pass-through decline in such countries.

- For Bolivia, inflation declined late 1988 and was accompanied by a significant fall of the pass-through coefficient in 1989 (ADF^* test statistic) or 1990 (Z_t^* test statistic). The inflation decrease in Bolivia is explained by the adoption of a New Economic Policy, namely the ESAF arrangement in 1988.

- Botswana's inflation fell early 1996, this fall was accompanied by a significant exchange rate pass-through decline in 1997 (ADF^* test statistic) or in 1998 (Z_t^* and Z_α^* test statistics). This corresponded to Botswana adopting the Exchange Rate Targeting in June 1994.

- Inflation in Chile declined twice, first early 1991 and then in 1993, leading to a significant pass-through decline at the end of 1993. In September 1990, the Central Bank of Chile decided to announce an inflation target for the year 1991 after the oil price shock due to the 1990 Gulf War and expansionary policies in 1989 causing high inflation rates.

- Indonesia incurred two falls in inflation rate as well, the first one in 1994 and the second in 1998. The exchange rate pass-through decline coincides with the second date (1998:3). After the Financial Asian Crisis 1997-98, Indonesia adopted an inflation target in May 1998.

- For Singapore, we observe two inflation falls, the first one in 1990:2 and the second in 1997:2. The second one corresponds to a decline in the pass-through. Throughout the Asian Crisis 1997-98, the Singapore Dollar (SGD) appreciated moderately because the Monetary Authority of Singapore (MAS) managed the exchange rate within an undisclosed band.

- Uruguay's inflation decreased late 1988 and was followed by an exchange rate pass-through decline in 1990. Uruguay shifted to a floating exchange rate in 1985. At the end of the 1980s, Uruguay adopted an exchange rate target (Adjustable peg/Exchange Rate Band).

5 Conclusion

Using a data set of 8 developing countries over the period $1980 : 2003$, we find evidence to support that exchange rate pass-through into import prices declined in the 1990s in 6 developing countries. Our empirical finding, based on some structural change and cointegration tests, gives support to Taylor's (2000) view that exchange rate pass-through decline is caused by a low inflation environment in the 1990s. Indeed, we show that such pass-through declines in these developing countries are associated with inflation decreases in the period from the 1980s to the 1990s. The principal explanations for this finding are that the change in the monetary policy regimes in these countries caused a shift to a low inflation environment and hence an exchange rate pass-through decline.

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Appendix A: Data Descriptions

- Aggregate inflation for developing countries is obtained from the percentage change of the CPI expressed in log (database IFS, series_code 20064..XZF).
- Countries inflation series are calculated as the percentage change of the CPI expressed in log.
- Nominal effective exchange rate (EXCHANGE RATE INDEX 1995=100), real effective exchange rate (based on real consumer price, INDEX 1995=100), Producer Price Index (Index Numbers (1995=100)), Industrial Price Index (INDUST PRODUCTION,SEAS. ADJ) and import unit values series (Index Numbers (1995=100)) are obtained from International Financial Statistics.

Appendix B: Empirical Results

Table 1. The decline in aggregate inflation in developing countries

$z_t = 1$		$z_t = (1, y_{t-1})'$	
$BIC = 2$	1983:1$^+$; 1992:4$^+$ $\hat{\delta}_1 = 0.0084$ $\hat{\delta}_2 = 0.0041$ $\hat{\delta}_3 = 0.0017$	$BIC = 1$	1995:4 $\hat{\delta}_{1,1} = 0.0012$ $\hat{\delta}_{1,2} = 0.0019$ $\hat{\delta}_{2,1} = 0.7100$ $\hat{\delta}_{2,2} = -0.2870$
$TBSP = 2$	1983:1$^+$; 1992:4$^+$ $\hat{\delta}_1 = 0.0084$ $\hat{\delta}_2 = 0.0041$ $\hat{\delta}_3 = 0.0017$	$TBSP = 1$	1995:4 $\hat{\delta}_{1,1} = 0.0012$ $\hat{\delta}_{1,2} = 0.0019$ $\hat{\delta}_{2,1} = 0.710$ $\hat{\delta}_{2,2} = -0.287$

Note: $^+$ In these dates, there is an aggregate inflation decline in the developing countries.

Table 2.1 Structural break tests results

	$z_t = 1$	
	BIC	$TBSP$
Bolivia	$\widehat{T}_1$: 1983:4; $\widehat{T}_2$: 1998:4	$\widehat{T}_1$: 1983:4; $\widehat{T}_2$: 1988:4; $\widehat{T}_3$: 1995:3
Botswana	$\widehat{T}_1$: 1996:1	$\widehat{T}_1$: 1996:1
Chile	$\widehat{T}_1$: 1993:2	$\widehat{T}_1$: 1991:1; $\widehat{T}_2$: 1995:2
Colombia	$\widehat{T}_1$: 1997:4	$\widehat{T}_1$: 1997:4
Indonesia	$\widehat{T}_1$: 1994:2; $\widehat{T}_2$: 1998:3	$\widehat{T}_1$: 0
Singapore	$\widehat{T}_1$: 1987:4	$\widehat{T}_1$: 1987:4; $\widehat{T}_2$: 1990:2; $\widehat{T}_3$: 1997.2
Uruguay	$\widehat{T}_1$: 1985:2; $\widehat{T}_2$: 1988:4; $\widehat{T}_3$: 1991:2; $\widehat{T}_4$: 1994:4	$\widehat{T}_1$: 1985:2; $\widehat{T}_2$: 1992:1; $\widehat{T}_3$: 1995.4
Venezuela	$\widehat{T}_1$: 1986:3; $\widehat{T}_2$: 1992:4; $\widehat{T}_3$: 1996.2	$\widehat{T}_1$: 1987:4; $\widehat{T}_2$: 1992:4

Table 2.2 Structural break tests results

	$z_t = (1, y_{t-1})'$					
	BIC			$TBSP$		
Bolivia	$\widehat{T}_1$			$\widehat{T}_1$		
	1985:1			1985:1		
Botswana	$\widehat{T}_1$			$\widehat{T}_1$		
	0			1996:1		
Chile	$\widehat{T}_1$			$\widehat{T}_1$	$\widehat{T}_2$	
	1991:2			1983:2	1991:2	
Colombia	$\widehat{T}_1$			$\widehat{T}_1$		
	1997:4			1997:4		
Indonesia	$\widehat{T}_1$	$\widehat{T}_2$		$\widehat{T}_1$	$\widehat{T}_2$	
	1994:2	1998:3		1994:2	1998:3	
Singapore	$\widehat{T}_1$			$\widehat{T}_1$		
	0			1985:1		
Uruguay	$\widehat{T}_1$			$\widehat{T}_1$	$\widehat{T}_2$	
	1983:2			1983:2	1985:2	
Venezuela	$\widehat{T}_1$	$\widehat{T}_2$	$\widehat{T}_3$	$\widehat{T}_1$	$\widehat{T}_2$	
	1989:3	1992:2	1995.4	1989:3	1992:3	

Table 3. Gregory and Hansen tests results

	ADF^*		Z_t^*		Z_α^*	
	Observed value	Break date	Observed value	Break date	Observed value	Break date
Bolivia	-6.73	1989:1	-6.73	1990:2	-61.17	1989:1
Botswana	-5.13	1997:4	-8.55	1998:1	-83.69	1998:1
Chile	-7.86	1993:4	-7.90	1993:4	-76.17	1993:4
Colombia	-4.37	1991:1	-7.90	1994:2	-34.87	1994:2
Indonesia	-6.95	1998:3	-7.30	1999:2	-68.67	1999:2
Singapore	-6.14	1997:2	-6.58	1985:1	-60.34	1985:1
Uruguay	-10.35	1990:3	-10.42	1990:4	-99.60	1990:4
Venezuela	-5.52	1989:4	-5.55	1990:1	-47.72	1990:1

Note: The asymptotic critical values of the tests ADF^*, Z_t^* and Z_α^* are respectively (-6.92 (1%), -6.41 (5%) and -6.17 (10%)), (-6.92 (1%), -6.41 (5%) and -6.17 (10%)) and (-90.35 (1%), -78.52 (5%) and -72.56 (10%)).

Table 4.1 Inflation decline in developing countries in the 1990s

	$z_t = 1$	
	BIC	$TBSP$
Bolivia	$\widehat{T_1} = 1983{:}4^+$ $\hat{\delta}_1 = 0.201$ $\hat{\delta}_2 = 0.018$	$\widehat{T_2} = 1988{:}4^+$ $\hat{\delta}_1 = 0.201$ $\hat{\delta}_2 = 0.017$
Botswana	$\widehat{T_1} = 1996{:}1^+$ $\hat{\delta}_1 = 0.011$ $\hat{\delta}_2 = 0.008$	$\widehat{T_1} = 1996{:}1^+$ $\hat{\delta}_1 = 0.011$ $\hat{\delta}_2 = 0.008$
Chile	$\widehat{T_1} = 1993{:}2^+$ $\hat{\delta}_1 = 0.019$ $\hat{\delta}_2 = 0.005$	$\widehat{T_1} = 1991{:}1^+$ $\hat{\delta}_1 = 0.011$ $\hat{\delta}_2 = 0.004$
Indonesia	$\widehat{T_1} = 1994{:}2,\ {}^+\widehat{T_2} = 1998{:}3^+$ $\hat{\delta}_1 = 0.024$ $\hat{\delta}_2 = 0.007$ $\hat{\delta}_3 = 0.0048$	no inflation deline
Singapore	$\widehat{T_1} = 1987{:}4^+$ $\hat{\delta}_1 = 0.0048$ $\hat{\delta}_2 = 0.0037$	$\widehat{T_1} = 1987{:}4,\ {}^+\widehat{T_3} = 1997.2^+$ $\hat{\delta}_1 = 0.00026$ $\hat{\delta}_2 = 0.000028$ $\hat{\delta}_3 = 0.002632$
Uruguay	$\widehat{T_2} = 1988{:}4,\ {}^+\widehat{T_3} = 1991{:}2^+$ $\hat{\delta}_1 = 0.072$ $\hat{\delta}_2 = 0.043$ $\hat{\delta}_3 = 0.0624$	$\widehat{T_1} = 1985{:}2,\ {}^+\widehat{T_2} = 1992{:}1^+$ $\hat{\delta}_1 = 0.0729$ $\hat{\delta}_2 = 0.0437$ $\hat{\delta}_3 = 0.233$

Table 4.2 Inflation decline in developing countries in the 1990s

	$z_t = (1, y_{t-1})'$	
	BIC	$TBSP$
Bolivia	$\widehat{T_1} = 1985{:}1^+$ $\hat{\delta}_{1,1} = 0.066$ $\hat{\delta}_{1,2} = 0.007$ $\hat{\delta}_{2,1} = 0.797$ $\hat{\delta}_{2,2} = 0.332$	$\widehat{T_1} = 1985{:}1^+$ $\hat{\delta}_{1,1} = 0.066$ $\hat{\delta}_{1,2} = 0.007$ $\hat{\delta}_{2,1} = 0.797$ $\hat{\delta}_{2,2} = 0.332$
Botswana	no inflation decline	$\widehat{T_1} = 1996.1^+$ $\hat{\delta}_{1,1} = 0.066$ $\hat{\delta}_{1,2} = 0.007$ $\hat{\delta}_{2,1} = 0.797$ $\hat{\delta}_{2,2} = 0.332$
Chile	$\widehat{T_1} = 1991{:}2^+$ $\hat{\delta}_{1,1} = 0.0184$ $\hat{\delta}_{1,2} = 0.0027$ $\hat{\delta}_{2,1} = 0.0916$ $\hat{\delta}_{2,2} = 0.5328$	no inflation decline
Indonesia	$\widehat{T_1} = 1994{:}2^+$, $\widehat{T_2} = 1998{:}3$ $\hat{\delta}_{1,1} = 0.0084$ $\hat{\delta}_{1,2} = 0.0058$ $\hat{\delta}_{1,3} = 0.0061$ $\hat{\delta}_{2,1} = 0.0704$ $\hat{\delta}_{2,2} = 0.932$ $\hat{\delta}_{2,3} = 0.216$	$\widehat{T_1} = 1994{:}2^+$, $\widehat{T_2} = 1998{:}3$ $\hat{\delta}_{1,1} = 0.0084$ $\hat{\delta}_{1,2} = 0.0058$ $\hat{\delta}_{1,3} = 0.0061$ $\hat{\delta}_{2,1} = 0.0704$ $\hat{\delta}_{2,2} = 0.932$ $\hat{\delta}_{2,3} = 0.216$
Singapore	no inflation decline	no inflation decline
Uruguay	$\widehat{T_1} = 1983{:}2^+$ $\hat{\delta}_{1,1} = 0.031$ $\hat{\delta}_{1,2} = 0.001$ $\hat{\delta}_{2,1} = 0.018$ $\hat{\delta}_{2,2} = 0.095$	no inflation decline

Note: $^+$ In these dates, there is an inflation decline for each country.

Table 5. Pass-through decline in developing countries in the 1990s

	Pass-through break point		
	ADF^*	Z_t^*	Z_α^*
Bolivia	1989:1$^+$ $\beta_1 = -0.380$ $\beta_1' = -0.392$	1990:2$^+$ $\beta_1 = 0.387$ $\beta_1' = -0.285$	1989:1$^+$ $\beta_1 = 0.387$ $\beta_1' = -0.285$
Botswana	1997:4$^+$ $\beta_1 = -0.177$ $\beta_1' = -0.987$	1998:1$^+$ $\beta_1 = 0.902$ $\beta_1' = 0.357$	1998:1$^+$ $\beta_1 = 0.902$ $\beta_1' = 0.357$
Chile	1993:4$^+$ $\beta_1 = -0.855$ $\beta_1' = -0.937$	1993:4$^+$ $\beta_1 = -0.855$ $\beta_1' = -0.937$	1993:4$^+$ $\beta_1 = -0.855$ $\beta_1' = -0.937$
Indonesia	1998:3$^+$ $\beta_1 = -0.381$ $\beta_1' = -0.971$	1999:2 $\beta_1 = -0.951$ $\beta_1' = 0.985$	1999:2 $\beta_1 = -0.951$ $\beta_1' = 0.985$
Singapore	1997:2$^+$ $\beta_1 = -0.120$ $\beta_1' = -0.803$	1985:1 $\beta_1 = -0.318$ $\beta_1' = -0.059$	1985:1 $\beta_1 = -0.3184$ $\beta_1' = -0.059$
Uruguay	1990:3$^+$ $\beta_1 = 0.829$ $\beta_1' = -0.170$	1990:4$^+$ $\beta_1 = 0.854$ $\beta_1' = -0.162$	1990:4$^+$ $\beta_1 = 0.854$ $\beta_1' = -0.162$

Note: $^+$ denotes a significant exchange rate pass-through decline or a significant inflation decline. Note that $\beta_1$and β_1' are respectively the values of the pass-through coefficient before and after the change point.

Table 6. Identified changes in the monetary policy

Country	Break date	Monetary policy change
Bolivia	1988:4	Exchange rate targeting
Botswana	1996:1	Exchange rate targeting
Chile	1991:1 and 1993:2	Inflation targeting
Indonesia	1998:3	Inflation targeting
Singapore	1997:2	Exchange rate targeting
Uruguay	1985:2 and 1988:4	Exchange rate targeting