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Countercyclical Capital Buffer: The Case of Uruguay

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Country Risk, Macroeconomic Fundamentals and Uncertainty in Latin American Economies

Abstract

This paper analyzes the relation between the country risk and its macroeconomic determinants for Argentina, Brazil, Mexico and Venezuela, during the 1998-2013 period, using a Markov-switching SUR model estimated by Bayesian techniques. Two independent regimes for each country were identified. The first one, associated with periods of stability and favorable international conditions, in which the variables under consideration behave as reported in the literature. On the other hand, the second regime temporarily coincides with periods of high domestic and international uncertainty. Our findings suggest that the changes in the analyzed relation depend on the origin of the uncertainty. If the uncertainty's source is associated with external shocks, such as international crises, the financial markets volatility gains relevance, while the solvency and liquidity variables are less relevant; if the causes of uncertainty are domestic, the latter are the key variables to explain the sovereign risk.

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1. INTRODUCTION

The impact of macroeconomic fundamentals on sovereign default risk has been studied in the traditional literature (Sachs, 1985; Edwards, 1986; González-Rozada, 2006; Uribe and Yue, 2006; Hilscher and Nosbusch, 2010) through linear models. Interest has grown recently in exploring nonlinear relations between sovereign default risk, its macro determinants and global variables, in different types of economies.

In advanced economies, discussion on the high levels reached by sovereign debt and the sustainability of fiscal policy, have demonstrated the importance of a nonlinear relation between debt size and borrowing costs, as well as the nonlinearities caused by uncertainty regarding the type of economic policy coordination designed to address the deterioration in fiscal accounts. Troy et al. (2010) studied the consequences of rising public debt in developed countries in an environment with a fiscal limit, concluding that uncertainty regarding the way economic policies are combined generates a nonlinear relation between debt and inflation. Huixin (2012) presents a study on the interactions between sovereign risk premia and fiscal policy under conditions of fiscal limit in developed countries, finding a nonlinear relation between sovereign risk premia and the level of public debt in line with the empirical evidence. Greenlaw (2013) analyzed the tipping points of sovereign debt markets for 20 advanced countries during the period 2000-2011. The authors find evidence of nonlinearities in the relation between borrowing rates on sovereign debt and its proportion on GDP of the economies studied. These authors point out that sovereign interest rates rise much more quickly when debt levels are high.

In emerging economies, linear models have encountered some difficulties in explaining the evolution of country risk over the last two decades in terms of their macroeconomic fundamentals and global variables due to factors such as political or economic uncertainty, contagion effects, among others. Acosta, Barráez and Urbina (2014) proposed a Markov regime switching model (Hamilton, 1989) to study the case of Venezuela. These authors suggest that in the process of forming expectations regarding a country's capacity to pay its debts, agents will not weigh the different macroeconomic determinants constantly over time. They identified two temporary regimes in which the linear relation between the fundamentals and sovereign default risk clearly varies regarding the temporal regime. These regimes temporarily coincide with periods of high and low economic uncertainty.

A number of the research papers mentioned above point to uncertainty as one possible cause of the nonlinear relation between sovereign debt and fundamentals. This empirical work studies the break in the linear relation between sovereign default risk and its determinants (macroeconomic fundamentals and global variables) for Argentina, Brazil, Mexico and Venezuela during the 1998-2013 period. To this end, we focus on exploring how uncertainty influences this break, according to the event that generates it, whether it is domestic (corresponds to specific events in each economy) or external (linked to international type events).

For this purpose, a Markov switching regime model is implemented, which unlike that the one presented by Acosta, Barráez and Urbina (2014), has a Seemingly Unrelated Regression (SUR) structure estimated with Bayesian simulation techniques (Kim and Nelson, 1999). The proposed model allows temporary states or regimes to be specified for each country, while carrying out the estimatation jointly, this enables to exploit the correlation that might exist between the shocks to the sovereign default risk in each country.

This paper verified the presence of nonlinearities between sovereign default risk and its determinants, and identified two temporary regimes in each country similar to that reported by Acosta, Barráez and Urbina (2014): a first regime, linked to periods of relative stability or *low uncertainty*, where the relation between country risk and fundamentals behaves according to the reported in the literature, and a second regime associated to periods of *high uncertainty*. The most significant finding of this research demonstrates that the changes occured in the relation between the risk and its explanatory variables depends on the causes of uncertainty in both regimes. If the source of uncertainty is associated with external events, such as international crises, financial market volatility gains relevance, while solvency and liquidity variables are less relevant, such as in the case of Mexico and Brazil. If the causes of uncertainty are of domestic origin, the opposite occurs, such as observed in Argentina and Venezuela. In the case of the latter, the results coincide with the findings of Acosta, Bárraez and Urbina (2014).

It is important to mention that the *subprime* crisis is the only common shock in the high uncertainty regime for all the countries, except Brazil, in whose case the relation of sovereign default risk with its determinants remained in the stability regime. Such behavior is probably explained by economic policy measures adopted in this country to address the crisis.

From the model estimated in this paper we obtained country risk elasticities with respect to their determinants in each regime. These elasticities are useful because they allow for assessing economic policies aimed at reducing sovereign default risk.

The paper is divided as follows. Sections 2 and 3 describe the main aspects of the data and econometric methodology. The fourth section presents the empirical model estimated. The fifth shows the results. Finally, the conclusions are given.

2. DATA

The database used for the estimation contains quarterly information for the period 1998-2013, for Argentina, Brazil, Mexico and Venezuela.

The EMBI+ index calculated by JP Morgan, obtained from Bloomberg, was used as a measure of sovereign default risk for each country included in the study. The variables considered as country risk determinants are divided into three groups: *macroeconomic fundamentals, solvency and liquidity variables, and global indicators.* The first group consists of the growth rate of real GDP, inflation and exchange rate variations. The second group includes international reserves, commodity prices and external debt as percentage of GDP. The third group of variables includes *global indicators*, such as market volatility and international interest rates.

In the case of macroeconomic fundamentals, data for GDP growth rate, inflation and the exchange rate are taken from IMF statistics for Argentina, Brazil and Mexico. In the case of Argentina, the price index registered by PriceStats was also used. In the case of Venezuela, these variables were obtained from Central Bank statistics, except for the parallel market exchange rate employed for calculating the spread with respect to the official rate as a measure of exchange rate imbalance, which is obtained from different sources.

Regarding liquid and solvency indicators, international reserves were expressed as months of imports obtained from IMF statistics. Data related to external debt was obtained from the statistics of the respective ministries of finance and statistics institutions of each country. This variable was expressed as a percentage of gross domestic product.

With respect to global indicators, market volatility is captured through the Chicago Board Options Exchange Volatility Index (VIX). The 3-month United States Treasury bill rate, obtained from Federal Reserve statistics, was used as the measure of international interest rates. Commodity prices were incorporated via the commodity price index, obtained from the IMF for Brazil and Mexico. The commodity price index published by the Banco de la República Argentina was used for Argentina, while for Venezuela the price series of the Venezuelan oil basket, obtained from the Venezuelan Energy and Oil Ministry, was employed. These criteria for selecting the indexes were based on the structure of exports, taking into account the most representative commodities of each country.

Before starting the estimation, unit root tests were carried out to detect stationarity in the series. Thus, the test of Levin, Lin and Chu (to verify the existence of common unit root processes), and those of Pesaran and Shin, W-Stat, ADF Fisher and PP Fisher, were employed to prove the existence of individual unit root processes. All the variables were transformed into logarithmic differences, except the coefficients (external debt/GDP, reserves/imports) and the interest rate, which are assumed at stationary in levels.

Selection of these economies was made considering the most representative Latin American countries in terms of the size of the economies for which the EMBI+ is elaborated (JP Morgan calculates the EMBI+ for 16 countries, six of which belong to Latin America). The study period was chosen taking into account the availability of statistical data.

3. METHODOLOGY

The multiple structural changes in Latin American economies would seem to suggest that a linear model for explaining default risk for each of the countries considered would be an inappropriate simplification. Thus, nonlinear Markov regime switching models seem to be more appropriate for adjusting to this type of behavior.

The instability in regression models is frequently associated to changes experienced by the equation's parameters from one sample period (regime) to another. If knowledge is available on when such regime changes occur, and the subgroups of the sample are well defined, the Chow F-test can be applied to prove the existence of the structural change hypothesis. However, in many cases very little information is available about the occurrence of such structural changes, then, in addition to the estimation of the model's parameters, the structural breaks of the equation must also be inferred as unobservable variables.

The SUR methodology was used in order to jointly estimate the Markov regime switching model, this provides information aboutthe correlation between the shocks to which risk is exposed in each considered country.

The Markov-switching SUR model can be written as follows:

1 $$y_{i,t} = x_{i,t}\ \beta_{i,s_t} + e_{i,t}\ ,$$

with $t = 1,\ldots,T$ observations for each of the $i = 1,\ldots,N$ equations (countries). $y_{i,t}$ denotes the sovereign default risk observation at time t of equation i, $x_{i,t}$ is a $1 \times k_i$ vector that contains the explanatory variables of equation i at time t, β_{i,s_t} represents the respective vector coefficients of equation i at time t, which has the following structure:

$\beta_s = \beta_i^0 \left(1 - s_{i,t}\right) + \beta_i^1 s_{i,t}$, $s_{i,t} = 0$ or 1 (regime 0 or 1).

$s_{i,t}$ is the unobservable variable that governs the regime change of equation i, during regime 0 the parameters of this equation are given by β_i^0, while during regime 1 they would be given by β_i^1.

Up until now nothing has been said regarding the characteristics of random errors in the model. $e_t = (e_{1,t}, e_{2,t}, \ldots, e_{N,t})'$ is defined to allow error correlation between cross-section units, we should assume that $e_t \sim N(\mathbf{0}_N, \Sigma)$ for $t = 1,\ldots,T$, where Σ is a co-variance matrix $N\times N$. Then, the likelihood function to maximize is given by:

2 $$\ln\left(L\right) = \sum_{t=1}^{T}\sum_{s_1=0}^{1}\sum_{s_2=0}^{1}\cdots\sum_{s_N=0}^{1} f(y_t \mid s_{1,t}, s_{2,t}, \ldots, s_{N,t}, \psi_{t-1}) \prod_{i=1}^{N} f(s_{i,t} \mid \psi_{t-1})$$

where

$$f\left(y_t \mid s_{1,t}, s_{2,t}, \ldots, s_{N,t}, \psi_{t-1}\right) = \frac{1}{(2\pi)^{\frac{N}{2}} \left|\Sigma\right|^{\frac{1}{2}}} \exp\left\{-\frac{1}{2}\left(y_t - x_t\beta_{s_t}\right)' \Sigma^{-1}\left(y_t - x_t\beta_{s_t}\right)\right\},$$

$$y_t = \left(y_{1,t}\ y_{2,t} \cdots y_{N,t}\right)',\ x_t = \begin{pmatrix} x_{1,t} & 0 & \cdots & 0 \\ 0 & x_{2,t} & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & x_{N,t} \end{pmatrix}, \beta_{s_t} = \begin{pmatrix} \beta_{1,s_t} \\ \beta_{2,s_t} \\ \vdots \\ \beta_{N,s_t} \end{pmatrix}, s_t = \begin{pmatrix} s_{1,t} \\ s_{2,t} \\ \vdots \\ s_{N,t} \end{pmatrix}.$$

Ψ_{t-1} represents the data available at time $t - 1$.

Finally, an assumption must be imposed on the stochastic behavior of the unobservable variables $s_{i,t}$, which will allow to determine $f(s_{i,t} \mid \psi_{t-1})$. If it is assumed that these follow a first-order Markov random process, the specification of a Markov regime switching model will have been completed. Inference of these variables is carried out through the Hamilton filter (1989).

When the model presented in Equation 1 depends on multiple cross-section units, each with explicative variables, the number of parameters to be estimated increases considerably and maximization of the likelihood function expressed in Equation 2 by classical methods becomes a very complicated task. Moreover, Bayesian methods provide several important advantages that avoid difficulties related to numerically maximizing the likelihood function with restrictions on the parameters imposed by economic theory. The use of prior densities, in addition to including information not contained in the sample into the estimation process, allows for working with smaller sized samples than those required by frequentist methods, which is of particular interest in our case. Regarding the estimation technique, the Bayesian simulation algorithms proposed by Kim and Nelson (1998) were employed for estimating the model. The idea is to use Gibbs sampling to obtain the posterior distribution of the parameters β_i^0, β_i^1, Σ, $i=1,\ldots,N$ and the state vectors $s_{i,t}$ from which their mean and variance can be inferred, thereby avoiding direct maximization of the likelihood function.

Gibbs sampling only requires posterior simulation of the conditional distributions of each parameter. Assuming a multivariate normal prior distribution for the vector of parameters $\beta = \left[\beta_1^{0\prime} \beta_2^{0\prime} \ldots \beta_N^{0\prime} \beta_1^{1\prime}\ \beta_2^{1\prime} \ldots \beta_N^{1\prime} \right]' \sim N\left(B_0, V_0\right)$ the posteriorconditional distribution $f(\beta \,|\, \Psi_T, \Sigma, s_{1,t}, s_{2,t}, \ldots, s_{N,t})$ will be given by $\beta \sim N(B_1, V_1)$, where:

$$V_1 = \left(V_0^{-1} + \mathbb{X}'\Sigma^{-1}\mathbb{X}\right)^{-1},$$

$$B_1 = V_1\left(V_0^{-1}B_0 + \mathbb{X}'\bar{\Sigma}^{-1}\mathbb{Y}\right),$$

$\bar{\Sigma} = \Sigma \otimes I_T$ ($\otimes$: Kronecker product operator),

$$\mathbb{Y} = \begin{pmatrix} Y_1 \\ Y_2 \\ \vdots \\ Y_N \end{pmatrix}, \; Y_i = \begin{pmatrix} y_{i,1} \\ y_{i,2} \\ \vdots \\ y_{iT} \end{pmatrix}, \; i = 1, \ldots, N$$

$$\mathbb{X} = \begin{pmatrix} X_1 & 0 & \cdots & 0 \\ 0 & X_2 & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & 0 & X_N \end{pmatrix} \odot \left(\iota_K \otimes S \right),$$

($\odot$: element-wise product operator)

$$X_i = \begin{pmatrix} x_{i,1} \\ x_{i,2} \\ \vdots \\ x_{iT} \end{pmatrix}, \; S = \begin{pmatrix} S_1 \\ S_2 \\ \vdots \\ S_N \end{pmatrix}, \; S_i = \begin{pmatrix} s_{i,1} \\ s_{i,2} \\ \vdots \\ s_{iT} \end{pmatrix}, \; K = \sum_{i=1}^{N} k_i, \; i = 1, \ldots, N,$$

ι_K denotes a $K \times 1$ vector that only contains ones.

To simulate the variance-covariance matrix Σ, an inverse Wishart priori distribution $\Sigma \sim IW(D_0, \delta_0)$ will be assumed, where D_0 and δ_0 represent a $N \times N$ scale matrix and degrees of freedom, respectively. The resulting posterior conditional distribution $f(\Sigma \mid \Psi_T, \beta, s_{1,t}, s_{2,t}, \ldots, s_{N,t})$ is of the same functional form: $\Sigma \sim IW\left(D_1, \delta_1\right)$, where:

$$D_1 = D_0 + E'E$$

$$E = \left(E_1 \quad E_2 \quad \cdots \quad E_N \right), \; E_i = \begin{pmatrix} e_{i,1} \\ e_{i,2} \\ \vdots \\ e_{i,T} \end{pmatrix}, \; i = 1, 2, \ldots, N, \; \delta_1 = \delta_0 + T.$$

To simulate the posterior distribution $f(S_i \mid \beta, \Sigma, \Psi_T)$ we use the Carter and Kohn (1994) result, which indicates that:

$$f\left(S_i \mid \beta,\Sigma,\Psi_T\right)=f\left(s_{i,T} \mid \beta,\Sigma,\Psi_T\right)\prod_{t=1}^{T-1} f\left(s_{i,t} \mid s_{i,t+1},\beta,\Sigma,\Psi_T\right),\ i=1,\ldots,N\ ,$$

where each of these distributions are obtained by implementing the Hamilton (1989) filter [for further details on this result see Carter and Kohn (1994) or Kim and Nelson (1999)].

4. EMPIRICAL MODEL

The base model estimation is given by:

$$EMBI_{i,t} = \theta_{S_t}^0 + \theta_{S_t}^1 \Delta PIB_{it} + \theta_{S_t}^2 \pi_{it} + \theta_{S_t}^3 R_{it} + \theta_{S_t}^4 Tc_{it} + \\ + \theta_{S_t}^5 Vix_{it} + \theta_{S_t}^6 Ti_{it} + \theta_{S_t}^7 D_{it} + \theta_{S_t}^8 \Delta PMP_{it} + \varepsilon_{it,S_t} ,$$

$$\varepsilon_{it,s_t} \sim N(0,\sigma_{s_t}^2)\ ,$$

$$\theta_{S_t}^i = \theta_0^i \left(1-S_t\right) + \theta_1^i S_t\ ,$$

$$S_t = 0 \text{ o } 1 \text{ (regime 0 or 1),}$$

where subindexes i and t denote the country and the time respectively, ΔPIB_{it} represents real GDP growth rate; π_{it} the inflation rate; R_{it} is international reserves expressed in months of imports; Tc_t is the variation of the exchange rate;[3] Vix_{it} is the CBOE volatility index; Ti_{it} is the three-month US Treasury bill interest rate; D_{it} is external debt as percentage of GDP and ΔPMP_{it} is the variation of commodity prices.

5. RESULTS

The results of the parameter estimation are shown in Table 1 of the Annex. Two regimes were identified for each country. The first regime, which we will call regime L (*low uncertainty*), is related to periods of stability, economic growth and favorable international conditions. The second one is the regime H (*high uncertainty*), which temporarily coincides with periods of domestic and international turbulence. The methodology employed allows regimes to be independent between countries and they do not necessarily coincide in temporality.

In regime L, for Argentina, Brazil and Mexico all of the determinants considered are statistically significant and the signs of the coefficients were as expected, except for the GDP growth rate in the case of Argentina. Out of the macroeconomic fundamentals, GDP growth rate has a negative sign, while inflation and the exchange rate have positive ones. Out of solvency and liquidity variables, debt has a positive sign, while international reserves and commodity prices have negative ones. Out of the global variables, the VIX has a positive sign. In this regime, the countries' sovereign default risk behaves as expected in the literature.

Unlike the rest of the countries, in Venezuela during regime L, risk is mainly determined by a small number of variables, being the most important oil prices and financial market volatility, confirming the results obtained by Acosta, Barráez and Urbina. This finding reflects the importance of oil revenues for the Venezuelan economy and the sensitivity of the yield curve of debt instruments to oil price shocks (Chirinos and Pagliacci, 2015): in periods of low uncertainty, sovereign debt risk perception is essentially linked to oil prices.

In contrast to regime L, regime H temporarily coincides with periods of high uncertainty where disturbances of international scope are present, such as the Russian crisis, the Argentine debt crisis and the subprime crisis, in addition to domestic events that adversely affected risk premia. In the case of Mexico and Brazil, the periods of high uncertainty are mainly associated with major external disturbances, while in Argentina and Venezuela this regime basically coincides with domestic type events.

Now we are going to analyze the results for each of country in regime H. In Mexico regime H is observed during the periods 1998Q1-1998Q3 and 2007Q2-2009Q2, coinciding with the Russian and subprime crises, respectively. Negative economic growth rates, depreciation of the Mexican peso and increased sovereign default risk were recorded in both periods.

With respect to the latter period, it is important to mention that out of the countries of the region, Mexico was most affected due to the synchronization of its business cycle with that of the United States.

With respect to the coefficients of regime H, inflation, GDP and the exchange rate cease to be significant. The international reserves and VIX coefficients are greater than those estimated for this variables in regime L. Unlike regime L, the sign of the coefficient for the United States Treasury bill ratesis negative in regime H, which reveals the importance of United States monetary policy in investors' valuation of Mexican debt.

In the case of Brazil, regime H, covering the period 2002Q2-2004Q2, was characterized by a significant deterioration in the terms of trade due to a decline in trade flows with Argentina as a consequence of the debt crisis affecting that country. In addition, the burst of the speculative bubble in 2000 and the events of September 2001 generated volatility in international markets. This unfavorable international environment caused a slowing of economic activity. During this period sovereign bond spreads surpassed 2,000 basis points (bp) and the real suffered a sharp depreciation.

With respect to the coefficients of regime H as in the case of Mexico, a group of variables ceased to be significant: inflation, external debt, GDP and United States interest rates. The exchange rate and the VIX increased their weight as risk determinants compared to those obtained by regime L. While the coefficients of international reserves and commodity prices are similar to those of regime L, the signs of the coefficients are as expected a priori.

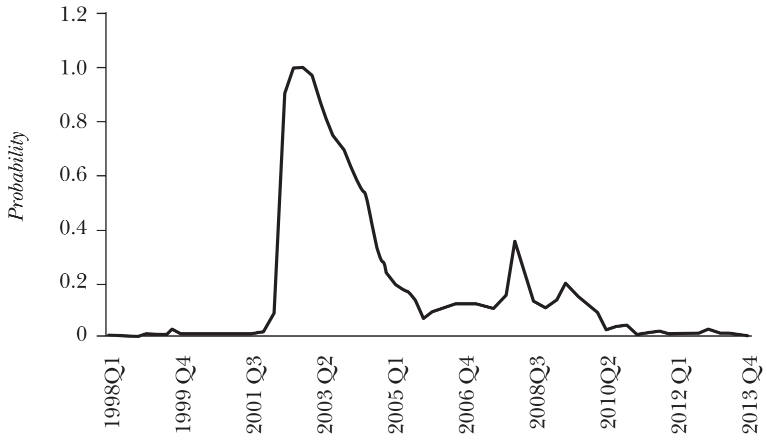
During 1998Q4-1999Q2 the presence of macroeconomic imbalances were seen after the collapse of the Plan Real, which increased risk premia. However, the methodology employed did not associate this period with regime H, given that this regime depends on the behavior of global indicators.

During the subprime crisis, no change of regime was observed either in sovereign default risk for Brazil, which remained in regime L, despite the increase in the regime switching probability

(Figure 1). Such behavior could be explained by the effectiveness of economic policy measures (mainly monetary and fiscal) mitigating the impact of the crisis.

Figure 1

PROBABILITY OF REGIME SWITCHING FOR BRAZIL



Source: own elaboration.

To assess whether monetary policy in Brazil influenced the evolution of sovereign default risk during the subprime crisis, a Taylor rule was estimated.

$$i-\bar{\pi}=\bar{r}+a\left(\pi-\bar{\pi}\right)+b\left(y-\bar{y}\right)+\varepsilon,$$

where i is the monetary policy interest rate of the Banco Central do Brasil (SELIC), $\bar{r}$ is the long-term interest rate, $\bar{\pi}$ is the inflation target, $\pi-\bar{\pi}$ is the difference between the actual inflation rate and the target, $y-\bar{y}$ is the GDP gap and ε is the monetary policy shock.

In order to test whether the policy mesures influenced the fact that soveriegn default risk remained in the regime of low uncertainty during the crisis period, the residuals of the Taylor rule (which express the orientation and magnitude of monetary policy) were captured for estimating a logistic model on the regime switching probability.

Figure 2 shows the probability of a change in country risk regime during the implementation of an expansive monetary policy measure (Figure 2a), comparing it with a counterfactual exercise assuming the implementaion of a neutral monetary policy (null shocks in the Taylor rule), which is shown in Figure 2b. In this regard, it can be seen how the regime change probabilityduring the crisis is higher in the abscence of monetary policy measures, i.e., monetary policy contributed to stay in the low uncertainty regime during the subprime crisis.

It is important to point out that fiscal policy actions were included in the logistic model, and counterfactual exercises were carried out similarly to those mentioned above, employing variables such as tax revenues and expenditures. However, these variables were not statistically significant, i.e., no statistical evidence was found to support the premise that fiscal policy influenced the presence of the Brazilian economy in the low uncertainty regime during the crisis.

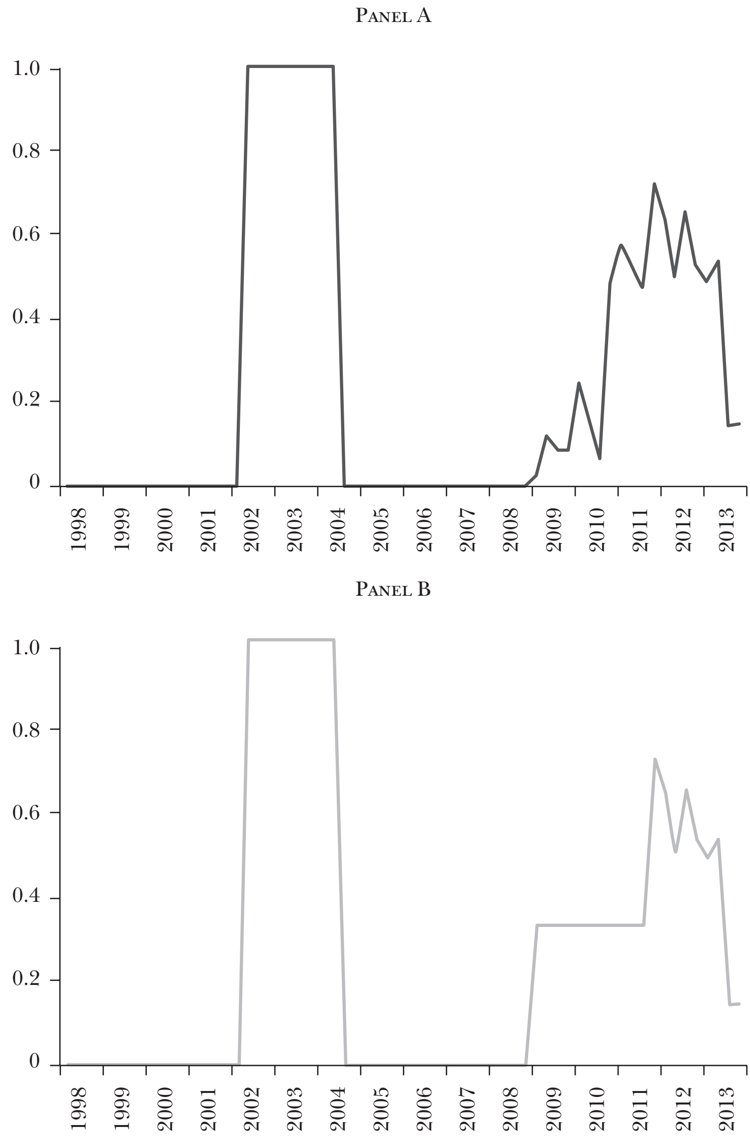
On the other hand, regarding Mexico and Brazil, international reserves is only variable of the solvency and liquidity group whose coeffcient increased during this regimeis in the case of Mexico. The other two of this group remained unchanged or even lost their significance, such as in the case of debt in Brazil, while the coefficient of the global variable VIX is higher in this regime for both economies.

In the cases of Argentina and Venezuela, regime H consists of three periods, mainly associated with adverse domestic economic and political events.

In Argentina, the first period (2001Q4-2005Q2) was marked by the public debt crisis of December 2001 and the subsequent social and political events that led to the resignation of the then president. The economy suffered a substantial contraction during this period, accompanied by a significant fall in

Figure 2

PROBABILITY OF BRAZIL COUNTRY RISK REGIME SWITCHING, GIVEN THE IMPLEMENTATION OF ECONOMIC POLICY MEASURES (PANEL A) OR NOT (PANEL B)



Fuente: own elaboration.

international reserves, depreciation of the exchange rate and a cessation of external public debt payments.

The second period (2008Q4-2009Q2) coincides with the outbreak of the subprime crisis that affected various countries in the region. The international crisis caused a slowing of economic activity, a decline in the terms of trade and a depreciation of the currency in Argentina. During this period, fears in the financial markets increased with respect to the Argentine government's ability to meet debt and interest payment commitments maturing in 2009. Thus, although the initial disturbance was of external origin, it passed through the domestic economy, affecting fundamentals, and solvency and liquidity variables.

The third period (2012Q1-2013Q4) was characterized by the application of economic policy measures, the most important ones being those related to renationalization of a majority share in the oil company Repsol YPF, foreign currency controls on domestic operations (mainly in the real estate sector) and the reduction of foreign currency hoarding by residents.

With respect to the coefficients, those for debt and the exchange rate cease to be significant in regime H. The most important changes are expressed in the size of the coefficient of commodity prices and the constant term, which represent almost double and triple the estimates for regime L, respectively. This reflects the growing importance that agents give to this liquidity indicator in response to the drop in international reserves.

For Venezuela, the first period that took place during 1998 coincides with the collapse of the fixed exchange rate regime and capital controls implemented since 1994, and the start of a system of exchange rate bands in July 1996. In the international context, several events occured during this period, such as the Asian crisis in July 1997, the Russian crisis of 1998 and the fall of international oil prices to historically low levels.

During the second period (2002Q1-2003Q2), important events such as the attempted coup d'état of April 2002 and the subsequent oil strike in December of the same year, which had economic and political repercussions. In the economic field, the substantial fall in international reserves led to the application of a new fixed exchange rate regime and capital controls.

The third and final period (2005Q3-2013Q4) was characterized by high risk premia deriving from domestic events such as socialist economic initiatives (nationalization of private companies: steelmakers, cement producers, and food processing firms, among others). In the international context, 2008 saw the default of Ecuador and the subprime crisis, which led to a contraction in the global economy, significant market volatility and a decline in oil prices. All the aforementioned considerably increased Venezuela's risk premium.

In regime H, Venezuela maintains its atypical behavior; all the variables are significant, except GDP and the exchange rate. It should be mentioned that the external debt and international reserves in regime L are not significant, while in this regime the former of these is the variable with the largest coefficient. The absolute values of the coefficients for the VIX and for commodities are smaller than in regime L. The behavior of Venezuela's sovereign default risk obeys to the specific characteristics of an oil economy, in periods of low uncertainty agents focus on oil prices and international market volatility to form their risk perceptions, while in periods of high uncertainty they consider other variables, besides those mentioned early.

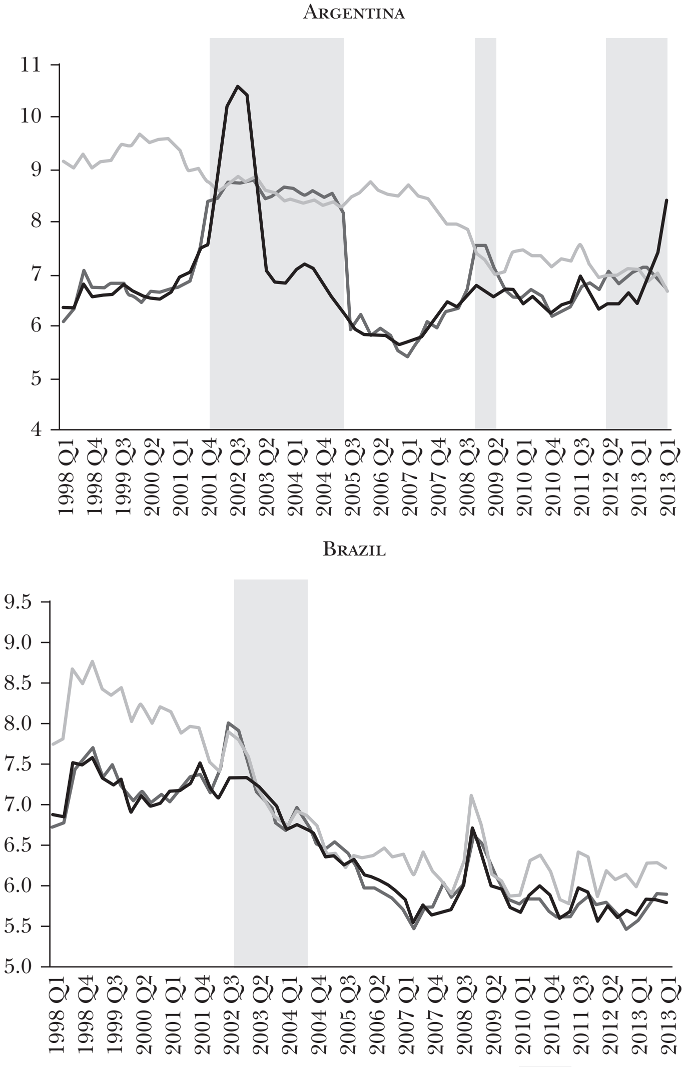
For Argentina and Venezuela, where the high uncertainty regime is associated with domestic events, the coefficient of the VIX decreased compared to those of the regime L in both cases. Regarding the solvency and liquidity group of variables, only the coefficientes of oil prices for Venezuela and debt for Argentina decreaseor lose their significance, the others increase their weight or remain the same. In Argentina's case, although two different measures of inflation were used, none of them were significant, regardless of the regime. The same was also observed for Venezuela.

The results obtained allow extracting some characteristics that are common to all the economies. In general terms, the results suggest that a change of regime in the relation between country risk and its determinants depends on the origin of the uncertainty. If the uncertainty's source is associated with external shocks (such as international crises), financial market volatility gains relevance, whereas the solvency and liquidity variables

Figure 3

MARKOV REGIME SWITCHING BY COUNTRY

(1998Q1-2013Q4)



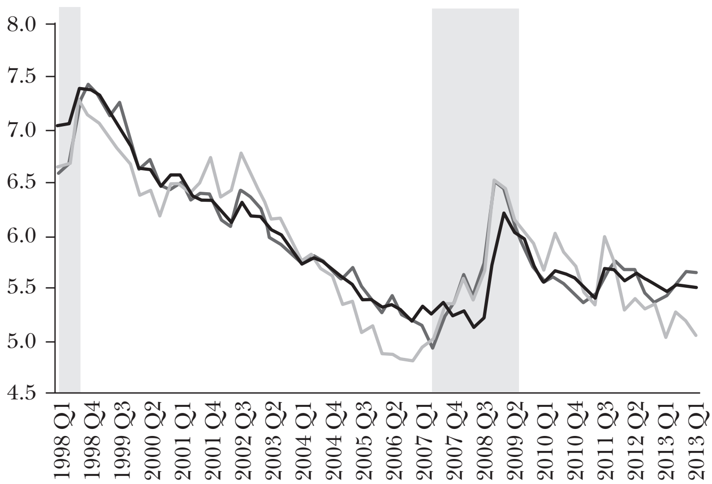
Note: lines in dark grey corresponds with the observed EMBI+ by each country, lines in light grey and the black ones correspond to the high and low uncertainty regimes, respectively. Shaded area is associates to high uncertainty and allows observing the regime switching easily.
Source: own elaboration.

Figure 3 (cont.)

MARKOV REGIME SWITCHING BY COUNTRY

(1998Q1-2013Q4)

Mexico



Venezuela

Note: lines in dark grey corresponds with the observed EMBI+ by each country, lines in light grey and the black ones correspond to the high and low uncertainty regimes, respectively. Shaded area is associates to high uncertainty and allows observing the regime switching easily.
Source: own elaboration.

are less relevant.While, if the triggers of uncertainty are of domestic origin, the latter are the key variables.

On the other hand, to assess the robustness of the base model, alternative models were estimated that took into consideration other control variables such as the degree of openness, government effectiveness, political stability/absence of violence and regulatory quality. The first is measured as the ratio between total imports plus total exports and GDP, while the rest of them are indexes prepared by the World Bank.

The model specificationthat includes degree of openness is the same as in the baseline model, but excludes the ratio of international reserves to imports and the external debt as a percentage of GDP because of problems of collinearity. The results of the estimation of this model are shown in Table 2 of the Annex. It can be seen that the model is robust after this variable is incorporated given that the regime changes registered and the majority of the parameters do not change significantly compared to the baseline model. This measure of openness was significant for Argentina and Brazil, with a positive sign in regime H and a negative one in regime L for both countries. This indicates that the more open these economies are during periods of high uncertainty, the more sovereign default risk is affected due to fears of contagion.

With respect to the other variables considered for estimating the alternative models, none of them were significant except for government effectiveness in the case of Argentina in regime H, with a negative sign as would be expected.

The model estimated allows for sovereign default risk elasticities to be derived with regard to their determinants in each regime. By simulating percentage increases in the respective exogenous variable, the resulting percentage variations in the endogenous variable are counted in order to obtain the desired elasticity. These elasticities, shown in Table 3 of the Annex, are useful for elaborating policies aimed at mitigating the impact of crises on sovereign default risk. The Table shows, for instance, how an increase of 1% in the exchange rate leads to an increase of 0.49% in sovereign default risk for the case of Mexico in the low uncertainty regime.

6. CONCLUSIONS

The results of this research point to the fact that the relation between sovereign default risk and its determinants for the countries considered has been disturbed by different types of events. In the international context, these events are related to the economic and financial crises that occurred during the study period: Russian crisis, Argentine debt crisis and the subprime crisis. In the domestic environment, these events are linked to macroeconomic imbalances, political instability and socialconflicts. The nonlinearity associated with this behavior was captured by estimating a Bayesian Markov-switching SUR model. This methodology allowed two independent regimes to be identified for each country.

The first regime, named regime L (*low uncertainty*), is related to periods of stability, economic growth and favorable international conditions. The second, regime H (*high uncertainty*), temporarily coincides with periods of international and domestic turbulence.

The results suggest that in the period of high uncertainty, agents give more importance to some key variables for forming their risk expectations. Such variables depend on the causes of the uncertainty. If the source of uncertainty is associated with external events, such as international crises, financial market volatility becomes important, such as in the case of Mexico and Brazil. If the triggers of uncertainty are of domestic origin, the key variables are the liquidity and solvency indicators of the country in question, as observed in Argentina and Venezuela. In the case of Venezuela, the results coincide with the findings of Acosta, Barráez and Urbina (2014), despite the differences with respect to the frequency of the statistical data used.

It should be pointed out that the subprime crisis is the only common event in regime H for all the economies, except Brazil, in whose case the relation between sovereign default risk and its determinants remained stable in regime L as a result of effective economic policy measures (mainly monetary and fiscal).

Annex A

Table 1

RESULTS OF ESTIMATES FOR THE BASE MODEL					
		Regime H		Regime L	
Country	*Coefficient*	*Posterior mean*	*Posterior 90% confidence bands*	*Posterior mean*	*Posterior 90% confidence bands posterior*
Argentina	Constant	13.5524	(12.52;14.57)	7.19	(4.36; 10.01)
	GDP	–	–	–	–
	IR/I	–	–	−0.28	(−0.40; −0.14)
	ED/GDP	–	–	2.85	(2.31; 3.38)
	VIX	0.38	(0.27; 0.49)	0.46	(0.37; 0.54)
	Treasury Bill	0.21	(0.16; 0.26)	−0.15	(−0.17; −0.12)
	IPMP	−1.35	(−1.46; −1.23)	−0.64	(−0.86; −0.41)
	ER	–	–	2.29	(1.66; 2.91)
Brazil	Constant	5.81	(2.40; 9.07)	7.95	(7.76; 8.13)
	π	–	–	3.23	(2.27; 4.22)
	GDP	–	–	−1.21	(−1.91; −0.51)
	IR/I	−0.21	(−0.31; −0.11)	−0.22	(−0.24; −0.20)
	ED/GDP	–	–	2.24	(1.53; 2.96)
	VIX	0.66	(0.41; 0.92)	0.27	(0.23; 0.31)

	Treasury bill	–	–	−0.06	(−0.07; −0.05)
	IPMP	−0.60	(−0.89; −0.29)	−0.59	(−0.62; −0.56)
	ER	0.66	(0.38; 0.93)	0.28	(0.18; 0.38)
Mexico	Constant	4.98	(3.78; 6.19)	4.44	(4.23; 4.66)
	π	–	–	3.77	(3.06; 4.49)
	GDP	–	–	−1.15	(−1.75; −0.56)
	IR/I	−0.88	(−1.41; −0.36)	−0.31	(−0.42; −0.19)
	ED/GDP	1.78	(0.58; 3.01)	3.18	(2.82; 3.53)
	VIX	0.85	(0.65; 1.05)	0.36	(0.30; 0.41)
	Treasury bill	−0.10	(−0.14; −0.06)	0.03	(0.02; 0.04)
	IPMP	−0.4 1	(−0.59; −0.24)	−0.30	(−0.34; −0.26)
	ER	–	–	0.49	(0.31; 0.66)
Venezuela	Constant	7.37	(7.01; 7.71)	6.24	(5.80; 6.67)
	GDP	–	–	–	–
	IR/I	−0.18	(−0.26; −0.10)	–	–
	ED/GDP	0.82	(0.40; 1.23)	–	–
	VIX	0.31	(0.25; 0.34)	0.56	(0.46; 0.66)
	Treasury bill	−0.31	(−0.32; −0.29)	0.03	(0.007; 0.04)
	IPMP	−0.29	(−0.33; −0.24)	−0.44	(−0.49; −0.37)
	ER	–	–	0.24	(0.11; 0.36)

Table 2

RESULTS OF ESTIMATIONS FOR THE MODEL INCORPORATING DEGREE OF OPENNESS

		Regime H		*Regime L*	
Country	*Coefficient*	*Posterior mean*	*Posterior 90% confidence bands*	*Posterior mean*	*Posterior 90% confidence bands*
Argentina	Constant	13.42	(12.81; 14.04)	9.54	(8.47; 10.57)
	GDP	–	–	–	–
	VIX	0.36	(0.25; 0.48)	0.8	(0.71; 0.89)
	Treasury bill	0.22	(0.17; 0.28)	−0.18	(−0.20; −0.16)
	IPMP	−1.47	(−1.61; −1.34)	−0.95	(−1.18;− 0.71)
	ER	−0.1	(−0.19; −0.01)	1.04	(0.20; 1.88)
	Degree of openness	1.01	(0.01; 2.02)	−2.83	(−3.83; −1.84)
Brazil	Constant	4.23	(3.32; 5.13)	6.46	(6.03; 6.90)
	π	–	–	8.09	(5.55; 10.61)
	GDP	–	–	–	–
	VIX	1.12	(0.91; 1.32)	0.33	(0.25; 0.40)
	Treasury bill	−0.08	(−0.12; −0.04)	0.03	(0.008; 0.04)
	IPMP	−0.35	(−0.57; −0.11)	−0.51	(−0.59; −0.41)
	ER	0.53	(0.27; 0.80)	0.32	(0.10; 0.53)
	Degree of openness	3.76	(0.00; 7.70)	−1.54	(−2.77; −0.32)

Mexico	Constant	4.87	(3.70; 6.06)	4.83	(4.52; 5.11)
	π	9.27	(6.10; 12.31)	8.93	(7.94; 9.86)
	GDP	-	-	−0.94	(−1.67; −0.22)
	VIX	0.41	(0.27; 0.55)	0.54	(0.46; 0.60)
	Treasury bill	−0.08	(−0.12; −0.04)	0.01	(0.00; 0.02)
	IPMP	-	-	−0.44	(−0.53; −0.34)
	ER	0.82	(0.28; 1.37)	-	-
	Degree of openness	-	-	-	-
Venezuela	Constant	7.88	(7.61; 8.13)	5.73	(5.17; 6.30)
	GDP	-	-	-	-
	VIX	0.25	(0.19; 0.30)	0.69	(0.55; 0.81)
	Treasury bill	−0.33	(−0.34; −0.31)	0.02	(0.00; 0.04)
	IPMP	−0.38	(−0.44; −0.31)	−0.41	(−0.53; −0.29)
	ER	-	-	0.26	(0.09; 0.43)
	Degree of openness	-	-	-	-

Table 3

COUNTRY RISK ELASTICITIES RELATIVE TO THEIR DETERMINANTS

Country	Variable	*Percentage change of EMBI + after a 1% increase in the variable*	
		Regime H	*Regime L*
Argentina	GDP	–	–
	Ir/I	–	−0.28
	D/GDP	–	0.93
	VIX	0.38	0.46
	Treasury bill	0.02	−0.01
	IPMP	−1.35	−0.63
	ER	–	2.32
Brazil	CPI	–	3.23
	GDP	–	−1.21
	Ir/I	−1.06	−1.28
	D/GDP	–	0.41
	VIX	0.66	0.26
	Treasury bill	-	−0.004
	Brent	−0.61	−0.58
	ER	0.66	0.28

Mexico	CPI	–	3.76
	GDP	–	−1.15
	Ir/I	−1.42	−0.50
	D/GDP	0.75	1.34
	VIX	0.85	0.35
	Treasury bill	−0.01	0.003
	Brent	−0.41	−0.29
	ER	–	0.49
Venezuela	GDP	–	–
	Ir/I	−0.17	–
	D/GDP	0.14	–
	VIX	0.31	0.56
	Treasury bill	−0.02	0.002
	Brent	−0.29	−0.43
	TC	–	0.24

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Jorge Ponce

Fundamentals for the Price of Housing in Uruguay

Abstract

This paper proposes a model of fundamentals for the price of housing. The model is calibrated with data from Uruguay. The main findings are: Real housing prices fluctuate more than justified by fundamentals; the misalignment was statistically significant just before the 2002 crisis; a fall in fundamental prices anticipates the crisis; and, in the recent period fundamental prices follow a stable trend of positive growth while real housing prices fluctuate around it.

Keywords: price of housing, model of fundamentals, financial stability, Uruguay.

JEL classification: G28.

1. INTRODUCTION

Deviations in the prices of some assets, particularly those of housing, from their equilibrium path can have significant consequences for financial system stability. When real estate (housing) prices remain higher than justified by their fundamentals for prolonged periods can result

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in excessive indebtedness, disproportionate risk taking and overvalued guarantees, all of which make the appearance of sudden and costly corrections more likely. In a global environment characterized by loose international liquidity conditions, substantial capital flows toward emerging economies and high commodity prices, as occurred after the international financial crisis of 2008-2009, assessing the degree of misalignment of housing prices from their fundamentals became particularly important.[1]

This paper proposes a model for estimating housing prices according to their fundamentals, which exploits the fact that a home can be considered an investment as well as a good that provides services. The model is calibrated with data from Uruguay. A comparison of real prices with those justified by fundamentals provides an indicator for the direction, size and duration of possible imbalances.

The main outcomes can be summarized as follows. First, real housing prices fluctuate more than justified by fundamentals, implying that periods of undervalued or overvalued housing prices are observed across the historical series. Second, prices according to calibrated fundamentals began a downward trend four years before the emergence of the 2002 crisis. The intensity of the fall increases during the year immediately prior to the crisis, showing there is a significant misalignment from real prices, which remain relatively stable. The overvaluation of real prices as compared to fundamental prices becomes statistically significant during that period. This fact, although derived from just one crisis event, argues in favor of using fundamental housing prices as a forward-looking financial fragility indicator. Third, in the most recent period, fundamental prices follow a stable trend of positive growth, while real housing prices in Uruguay fluctuate around it without exhibiting any statistically significant changes. Thus, there are no signs

[1] The following section presents a review of the literature analyzing the links between the international environment and domestic asset prices.

of imbalances between real housing prices and those justified by fundamentals. Nevertheless, it is important to make the following observations. The fact that real housing prices are aligned with fundamentals does not mean the former cannot fall in the future. As mentioned above, real prices fluctuate more than fundamental prices. In addition, prevailing international economic conditions lend weight to the hypothesis that fundamentals for housing prices (particularly income variables) might be overvalued. This point should be the subject of further study. Fourth, housing supply and construction series, as well as estimates for other important variables such as the deprecation rate, maintenance costs and risk premium, have been generated as a byproduct of the calibration exercise.

The rest of the paper is structured as follows. The next section briefly describes the related literature. Section 3, provides a summary of Uruguay's real estate sector. Section 4, shows the model of fundamentals for the price of housing. Section 5 calibrates the model with data for Uruguay. Section 6 offers some concluding remarks.

2. LITERATURE REVIEW

A set of recent contributions have addressed the links between global imbalances, capital flows, international liquidity conditions and asset prices. Hirata et al. (2012) show that housing prices in developed countries tend to move together (they are synchronized), and how this synchronization has increased over time. Among the determinants of global house price fluctuations, the authors find that innovations to the global interest rate (or loose monetary policy) have a significant impact on housing prices. Aizenman and Jinjarak (2009) also found evidence of an increase in the synchronization of prices in world real estate markets. Their paper also shows there is a robust and strong positive association between current account deficits and the real appreciation of housing. This association is stronger in deeper financial markets. Taguchi (2011) studied the response of asset prices to capital flows in East Asian countries. In every case they found a positive response of share prices to portfolio flows. Moreover, this effect is magnified by the indirect impact of monetary policy in countries with a fixed exchange rate regime. Vásquez-Ruiz (2012) studied a panel of 46 countries and found a strong positive association between housing prices and portfolio flows. The exchange rate regime also affects the strength of this relation. Kim and Yang (2011) found that capital flows to Asian countries have contributed to the appreciation of equities and land. Meanwhile, Favilukis et al. (2012) found that capital flows only have a limited effect on housing prices when the endogenous effects on risk premia and the expected supply of housing are considered. On the other hand, changes that modify access to mortgage credit have a large impact on prices.

The literature mentioned in the preceding paragraph does not study the direction of causality between current account deficits and asset prices. Laibson and Mollerstrom (2010) found evidence suggesting that causation runs from asset price bubbles to current account deficits. In particular, movements in asset prices explain more than 50% of the current account deficits in OECD countries. Gete (2010) formally showed how an increase in the demand for housing can generate a current account deficit. Jinjarak and Sheffrin (2011) analyzed the causal relations between the current account deficit and real estate prices. They found little evidence that the former cause the latter.

From a financial stability perspective, it is necessary to be able to identify when a series of asset prices is misaligned from their fundamentals or equilibrium path. Garriga et al. (2012) found that the behavior of housing prices can be correctly described by using standard asset pricing formulas. However, the pricing model required is highly nonlinear. Hott (2009) provided a nonlinear model for calibrating housing prices according to their fundamentals. Both papers conclude that observed prices vary more than that predicted by their fundamentals. Borraz et al. (2012) estimated a dynamic panel data model for 32 countries between 1990 and 2011, studying the relation between housing prices and global variables. The methodology allows periods of overvaluation to be identified for each country when using the estimated relation as a fundamental for housing prices. The results are in line with those found by Jara and Olaberría (2012), who used the methodology of Olaberría (2011).

Olaberría (2011) studied the link between capital flows and financial asset price booms measured as a deviation of current prices from a Hodrick-Prescott trend, finding a strong and significant association between these variables for emerging countries. The author also found no evidence to support that capital controls reduce such association. Hott and Jokipii (2012) used the fundamentals model of Hott (2009) for identifying misalignments in housing prices, finding a significant and positive link between low interest rates and house price bubbles. In addition, the empirical evidence supports the hypothesis that interest rates that have remained low for long periods magnify the impact. Cubeddu et al. (2012) studied a sample of Latin American countries and analyzed whether housing prices were aligned with their fundamentals and up to which point the growth of mortgage credit was excessive with respect to long-term trends. The authors concluded that there were no significant misalignments of housing prices from their fundamentals (estimated through a Hodrick-Prescott trend), but these could materialize if current trends persist.

Price indexes are elaborated by different methodologies, according to the literature. Several methods are also employed for estimating the part of observed prices that can be explained by the evolution of fundamentals in order to identify potential misalignments. The intended goal, the relative complexity of the methodologies and the availability of data determine the choice of one over another. Ponce and Tubio (2013) systematized the methodologies for elaborating housing price indexes and evaluated their applicability to the case of Uruguay. On the one hand, simple methodologies such as that of *repeat sales* do not appear to be applicable given low housing rotation in the Uruguayan market. On the other, highly complex methodologies such as *hedonic models* would not be applicable either, at least for obtaining a relatively long series, due to the lack of disaggregated data.[2] Thus, data availability determines the use in this paper of a mix-adjusted price index (see section 4.2.1). Meanwhile, the methodology proposed for identifying periods of overvalued housing prices is based on microeconomic fundamentals (see section 4). It therefore contributes with an estimation that is complementary to those carried out using Hodrick-Prescott filters (e.g., Jara and Olaberría, 2012) or dynamic panel models (see Borraz et al., 2012, who include the case of Uruguay).

3. THE RESIDENTIAL REAL ESTATE SECTOR IN URUGUAY

The Uruguayan economy has exhibited significant strength in recent years. The average annual growth rate of Uruguay's gross domestic product (GDP) was almost 6% during the last decade. In this context of strong economic growth and loose international financing conditions, real estate sector activity expanded considerably. In fact, the greater income and improved financing possibilities that characterize positive economic cycles also benefit activity in the real estate sector. In particular, construction has exhibited significant strength since 2003, doubling levels of activity as compared to those prevailing in the recession of 2002-2003. In recent years, the sector has shown considerable dynamism mainly driven by housing construction in Punta del Este and the Montevideo coastal zone. According to data from the National Institute of Statistics, the constructed surface area in Montevideo totaled 240,000 square meters in 2012, as compared to around 60,000 square meters in 2002. For several years, the sector's activity was concentrated in building tower blocks and apartments on the coastal zone, but this segment has exhibited some signs of saturation over the last few years.

[2] Landaberry and Tubio (2015) recently elaborated a hedonic price index for housing in Uruguay for the period 2009-2013 by compiling a new database that incorporates data from different sources.

The growth in housing construction and supply was accompanied by a significant expansion in investment. It is estimated that in 2011 and 2012 investment in housing construction amounted to around 1.6 billion USD per year. This represented an important increase from the figures recorded in 2008 and 2009, when investment in housing was a little under 900 million USD. Meanwhile, investment in construction has captured over a quarter of the foreign investment received by the country. As mentioned previously, the luxury segment of Montevideo and Punta del Este has been the main focus for such investments, especially by Argentine investors.

On the demand side, figures from the National Institute of Statistic's Continuous Household Survey show that around 60% of Uruguayan households own their own home. Furthermore, if household income levels are included in said analysis, it can be seen how the percentage difference of households with own homes among the lowest and highest quintile is 20%, meaning on average 50% of lower income households own the homes they live in. Slightly over 80% of total households with homes have already paid for them. Thus, housing represents the most important asset of the average Uruguayan household.

With respect to financing for home purchases, Uruguay is characterized by low mortgage credit penetration. Although in nominal terms mortgage credit in Uruguay has registered significant growth in recent years and constitutes the average household's main liability, in terms of gross domestic product it has remained relatively stable at around 4%, which is a rather low level compared to international values.

Rental agreements are freely negotiated in Uruguay, meaning the parties are who decide on the term, currency and payment readjustment system. This flexibility has allowed market prices to respond to fundamentals, which function as appropriate signals for agents in the sector to make decisions. Under the context of the significant economic growth registered during the last few years, the rental market has exhibited substantial strength, above all in Montevideo, where there has been a continuous increase in the number of contracts signed.

4. THE MODEL

This section describes the model of domestic fundamentals for housing prices. It is inspired by Hott (2009) and exploits the fact that a home can be considered as an investment asset as well as a good that serves utility. The price of a home can therefore be considered in two complimentary ways: as an outcome of the market for housing services or the equilibrium in an asset market.

The model of fundamental prices presented here combines both interpretations. The asset view of housing is the first to be considered: the house prices are defined as the present value of future rents. In second place, future rents are calculated over a housing market equilibrium, i.e., housing is considered as a consumer good. Finally, fundamental housing prices are calculated by replacing the imputed future rent in the present value equation by its fundamental value.

4.1 Asset: Current Value of Future Imputed Rents

When considering housing as an asset, its price is defined as the present value of future rents. The arbitrage condition implies that in equilibrium any individual must remain indifferent as regards purchasing or renting a home. Calibration of the model will exploit this condition. The cost per period of renting a home is *rental*, M_t, while the cost per period when purchasing a home is the imputed rent, H_t.

To calculate imputed rent the following factors are considered:

1) Financial cost (opportunity cost): the costs of financing for purchasing a home (or the opportunity costs due to the unavailability of own money) are represented by $m_t P_t$, where m_t is the interest rate in period t and P_t is the price of housing during the same period.[3]

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[3] The assumption implicit in this representation of financial costs is that the whole value of the home, the share financed with loans

2) Maintenance costs and a risk premium: Maintenance costs and a risk premium are modelled as a fixed proportion of the price of housing, ρP_t.

3) Expected net capital gains or losses: The expected capital gains or losses are given by the variation of housing prices from one period to another. This expected capital variation is modeled in net terms from the depreciation, δ the house suffers over time: $(1-\delta)E_t\left(P_{t+1}\right)-P_t$.

Imputed rent (H_t) is obtained by including these items as follows:

1 $$H_t = \left(m_t + \rho + 1\right)P_t - \left(1-\delta\right)E_t\left(P_t + 1\right).$$

When defining discount factor $R_t = 1 + \rho + m_t$, P_t is taken out of equation 1, and forward iteration then implies the following for the price of housing:[4]

2 $$P_t = E_t\left[\sum_{i=0}^{\infty} \frac{\left(1-\delta\right)^i H_{t+i}}{\prod_{j=0}^{i} R_{t+j}}\right].$$

4.2 Consumer Good: Imputed Rent

Equation 2 implies that the price of a house is equal to the current expected value of future imputed rents. The equilibrium condition in the housing market is used to calculate these rents. In particular, the sequence of imputed rents, H_t, must determine a demand for housing equal to their supply.

The supply of housing in a specified period, S_t, is determined by the supply in the preceding period net of depreciation and by the construction of new units, B_{t-1}:

as well as that financed with own funds, is discounted at the same interest rate m_t.

4 This is one particular solution (without rational bubbles) for the equation in finite differences where convergence is ensured by assuming that $\delta > 0$.

3 $$S_t = (1-\delta)S_{t-1} + B_{t-1} = (1-\delta)^t S_0 + \sum_{j=1}^{t}(1-\delta)^{j-1}B_{t-j},$$

where S_0 is the initial supply of housing.

To determine the demand for housing, D_t, it is assumed that there is a finite number of identical individuals in the economy. Each individual has Cobb-Douglas type consumption preferences for housing and other goods in such way as the proportion of aggregate income, Y_t, allocated to consumption is represented by parameter α. The demand for housing is therefore given by:

4 $$D_t = \alpha \frac{Y_t}{H_t} \cdot$$

The supply and demand for housing are made equal and by rearranging the following equation for imputed rent is obtained:

5 $$H_t = \alpha \frac{Y_t}{(1-\delta)^t S_0 + \sum_{j=1}^{t}(1-\delta)^{j-1} B_{t-j}} \cdot$$

4.3 Fundamental House Prices

By placing the value of imputed rents from equation 5 into the price equation (equation 2), and using the supply of housing in equation 3, the following fundamental house price equation is obtained:

6 $$P_t^* = E_t\left[\sum_{i=0}^{\infty} \frac{(1-\delta)^i \alpha Y_{t+i}}{S_{t+i}\prod_{j=0}^{i} R_{t+j}}\right].$$

5. CALIBRATION

Calibration of the model's parameters is carried out in two stages. The first of these exploits the arbitrage condition according to which any individual must be indifferent between renting and purchasing a home. The parameters are chosen in a way that minimizes the square differences between actual rents, M_t, and the value of imputed rents, H_t. In the second stage the parameters are calibrated in a way that minimizes the square differences between house prices and fundamental prices. This calibration process allows for exploring the nonlinearity of the fundamental price model, which according to Garriga et al. (2012) is necessary for correctly describing housing prices.

5.1 First Stage: Rent and Imputed Rent

The parameters of equation 5 are calibrated in the first stage. The parameter of the preference for housing, α ; the depreciation rate, δ ; and the initial supply of housing units, S_0. It is also necessary to calibrate a fourth parameter, the initial construction of new housing units B_0, given the conditions of consistency demanded by the series generated for the supply of housing, S_t, and the construction of new units B_t. In particular, the series for the construction of new housing units is made to follow the evolution (variation) of the index of gross physical capital formation in residential buildings:

$$B_t = B_0 \frac{I_t}{I_0} ,$$

where I_k is the index of gross physical capital formation in residential buildings in period k. Meanwhile, the series for the supply of housing is forced to replicate the number of housing units reported by the relevant censuses of 1996, 2004 and 2010 with a five percent margin of error.

The calibration of the parameters solves the following optimization problem:[5]

$$\min_{\alpha_1,\delta,S_0,B_0} = \sum_{t=0}^{T}\left[\alpha_1 \frac{Y_t}{(1-\delta)^t S_0 + \sum_{j=1}^{t}(1-\delta)^{j-1} B_{t-j}} - M_t\right]^2, \tag{7}$$

where α_1 is the parameter of the preference for housing, α, corrected by the different scales of the variables employed. Calibration takes into account the following restrictions to parameter values: $\alpha_1 > 0$, $0 < \delta \leq 0.025$, $S_0 \geq 988{,}525$ and $B_0 \geq 0$. Given that α is the parameter of preferences for housing, α_1 has to be positive. The restriction for the parameter of depreciation, δ is introduced to guarantee that the minimum lifecycle of a housing unit is ten years and the maximum is potentially unlimited (remembering that quarterly data will be used for the calibration). The restriction for initial housing levels, S_0, is given by the total supply of housing units reported in the census of 1985 (remembering that the calibration is carried out with a data set starting in the first quarter of 1988). Meanwhile, construction of new units, B_0, must necessarily be non-negative.

5.1.1 Data

Quarterly data for Uruguay is used for calibrating. The period of analysis runs from the first quarter of 1998 to the second quarter of 2011.

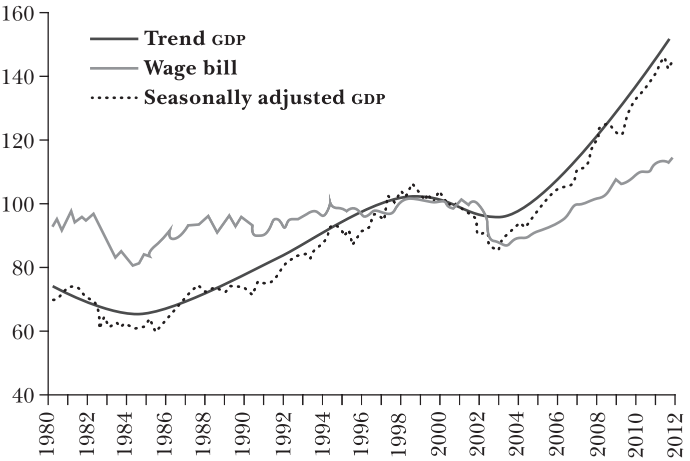
Three variants are used for aggregate income, Y_t: *i)* the index of the physical volume of seasonally adjusted gross domestic product, *ii)* the Hodrick-Prescott trend for the index of the

[5] The assumption that there are no financial market frictions is implicit in this model, meaning the arbitrage condition should always comply equally. This is particularly important in the case of Uruguay given the shallowness of the real estate market and the existence of special housing development plans throughout the last few decades. Many of these plans used development mechanisms other than interest rates (m_t) and it is therefore not possible to incorporate their impact on mortgage credit in this paper.

Figure 1

AGGREGATED INCOME, Y_t

(average 2000=100)



physical volume of gross domestic product, and *iii)* the wage bill calculated as real average wages multiplied by the number of workers. Figure 1 shows these variables. Rents correspond to the National Institute of Statistics' real price series for rents. A description of the series employed is provided in the annex.

5.1.2 Results

Table 1 shows the results of the calibration.[6] As can be seen, the results do not differ substantially when different variants are used for aggregate income, Y_t. In particular, the 0.66% quarterly depreciation parameter implies that, on average, a home fully depreciates in 40 years. Meanwhile, the results obtained by using gross domestic product as aggregate income imply an estimate for the initial supply of housing (for the first

[6] Calibration of parameter α is omitted given that assumptions on the factor of scale contained in calibrated parameter be made in order for it to be identified.

quarter of 1988), S_0, only slightly above the 1985 census value. Calibration of the initial supply of housing obtained by using the wage bill as the aggregate income variable is in line with the data from the 1985 census, as well as with the depreciation rate and initial construction of new units that were calibrated. Figure 2 shows the supply of housing arising as a result of the calibration exercise and compares it with census data. Figure 3 shows the series for the construction of new housing units arising from the calibration and compares it with the index of gross physical capital formation in residential buildings.

Table 1

CALIBRATED PARAMETERS

Y_t	δ	S_0	B_0
Seasonally adjusted GDP	0.0066	990,438	7,112
Trend GDP	0.0063	988,747	7,127
Wage bill	0.0066	1,017,982	6,893

Figure 4 shows the historical series for real rents, comparing it with imputed rents when the wage bill is used as the aggregate income variable. The results of the use of GDP as aggregate income variable are presented in the Annex. In all cases, real rents were higher than their fundamental values during the 1990s. The situation is reversed with the advent of the crisis in 2002. Toward the end of the series real rents have a null misalignment (when gross domestic product is used as a variable for aggregate income), or of around 10% above the fundamental value (when the wage bill is used as an aggregate income variable).

Figure 2

CALIBRATED HOUSE SUPPLY, S_t, AND CENSUS DATA

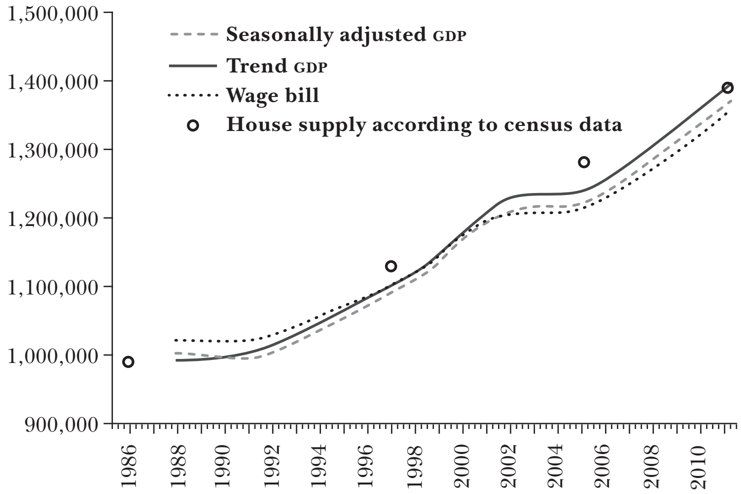


Figure 3

CALIBRATED NEW HOUSES CONSTRUCTION (B_t) AND INDEX OF PHYSICAL GROSS CAPITAL FORMATION IN RESIDENTIAL BUILDINGS (I_t)

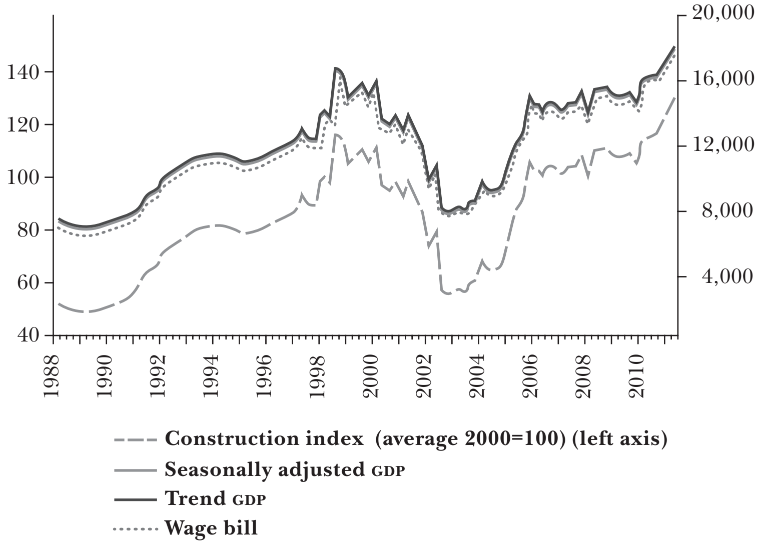
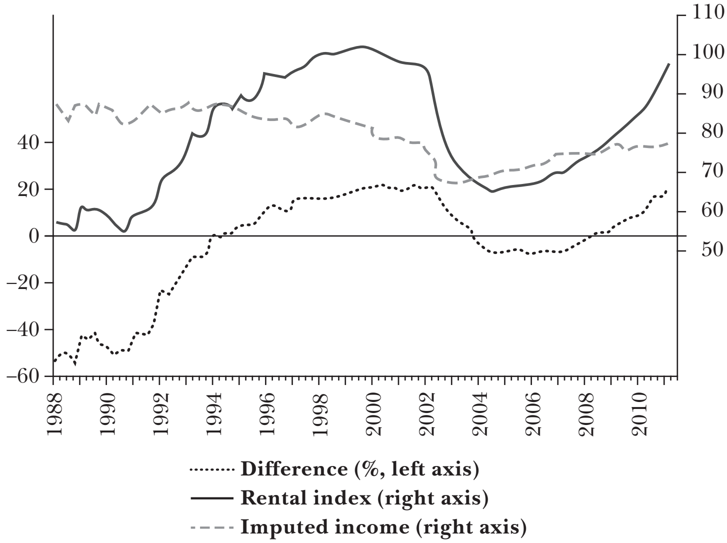


Figure 4

IMPUTED INCOME (Y_t = WAGE BILL), RENTAL INDEX AND DIFFERENCE



5.2 Second Stage: Real and Fundamental Prices

Once rents for housing have been imputed it is possible to calibrate the remaining parameters, particularly the risk premium ρ, through Equation 6 (or equivalently with Equation 2 by replacing H_t with the values calibrated in the first stage), and obtain a fundamental price series for housing. Equation 6 can be written as follows:

$$P_t^* = \sum_{t=0}^{T} \frac{H_t + \delta P_{t+1}^*}{1 + \rho + m_1} . \tag{8}$$

To use Equation 8 it is necessary to calibrate the future value of fundamental prices $\left(P_{T+1}^*\right)$, for which assumptions must be made on the future evolution of fundamentals. Equation 6 shows that fundamental housing prices are determined by the ratio of aggregate income to housing supply, as well as a discount factor. For the sake of simplicity, the following assumptions are made: *1)* the ratio of aggregate income to housing supply evolves following a trend, $\frac{Y_{t+1}}{S_{t+1}} = (1+w)\frac{Y_t}{S_t}$; and *2)* the interest rate as of period $T+1$ remains constant at its average value $\bar{m}$. Hence, the future value of fundamental prices, P^*_{T+1} , can be written as:

9

$$P^*_{T+1} = \alpha \frac{(1+w)Y_T}{(\rho + \bar{m}\delta - w + \delta w)S_T},$$

which introduces two extra parameters, w and $\bar{m}$, that must be calibrated.

Calibration in this second stage is made by solving the following optimization problem:

10

$$\min_{\alpha_2, \delta, w, \bar{m}} \sum_{t=0}^{T} \left[\frac{\alpha_2 H_t + \delta P^*_{t+1}}{1+\rho+m_t} - P_t \right]^2,$$

where α_2 is a parameter of scale. Calibration is made taking into account the following restrictions on parameter values: $\alpha_2 > 0$, $0 < \rho \leq 0.03$ and $0 < \bar{m} \leq 0.025$. The restriction to the parameter of maintenance costs and risk premium, ρ, is the same as that used in Hott (2009). Meanwhile, the restriction to the real average interest rate implies a maximum in annualized terms of approximately 10 percent.

5.2.1 Data

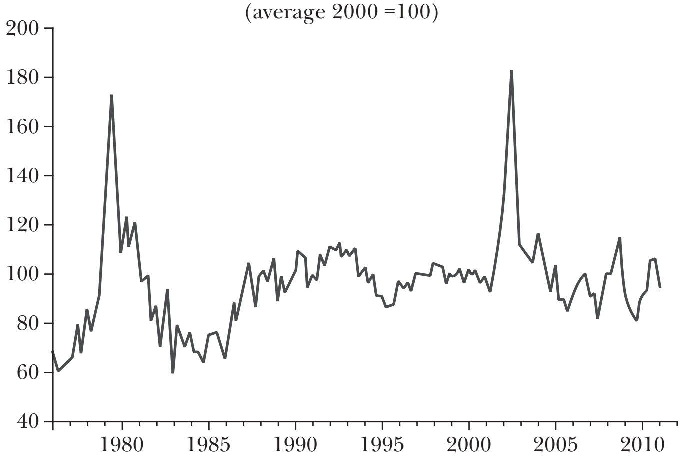
Quarterly data is used for calibrating. The period of analysis runs from the first quarter of 1988 to the second quarter of 2011.

In addition to the imputed rents calculated in the previous stage, a real housing price index is employed (see Figure 5). The latter arises from linking the National Institute of Statistics's housing price index with the pre-1999 data calculated

Figure 5

REAL HOME PRICE INDEX

(average 2000 =100)



by Carlomagno and Fernández (2007), following the methodology proposed by Grau et al.,1987.[7] Finally, the real interest rate series was generated from the banking system's active rate series for the non-financial sector by removing the inflationary component.

5.2.2 Results

Table 2 shows the calibration results. As can be seen, the inclusion of different aggregate income indicators does not substantially affect the results of the calibration. Moreover, only the restriction to the risk premium operates when gross domestic product is considered as an aggregate income variable.

[7] It is important to point out that the linked series offer different coverages. The series taken from the National Institute of Statistics is compiled based on the transactions that were actually made. The other series are based on press releases on the prices demanded by those offering houses for sale. None of the series are adjusted for housing quality or characteristics.

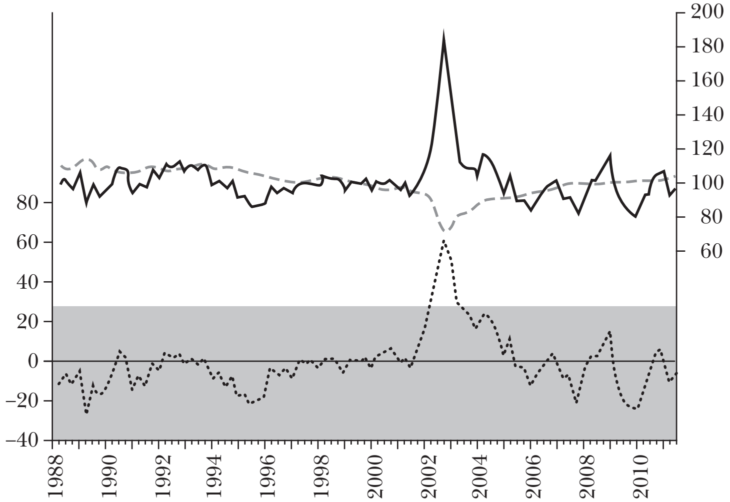
Table 2

CALIBRATED PARAMETERS

Y_t	ρ	$\overline{m}$
Seasonally adjusted GDP	0.030	0.020
Trend GDP	0.030	0.021
Wage bill	0.029	0.018

Figure 6

PRICE ACCORDING ITS FUNDAMENTALS
(Y_t = WAGE BILL), REAL PRICE AND DIFFERENCE
(in gray: interval of confidence at 95%)

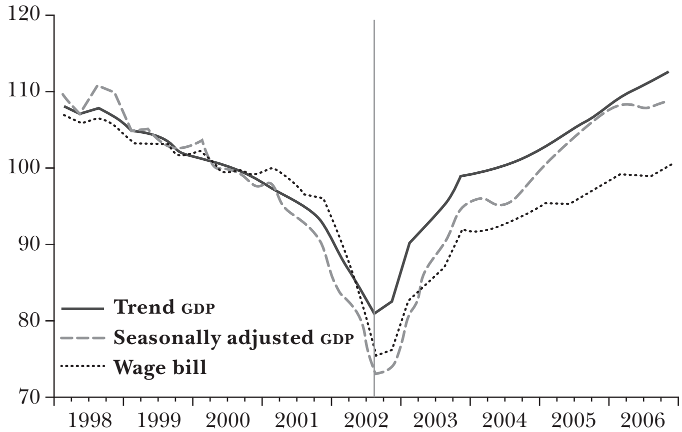


.... **Difference (%, left axis)**
—— **Price (right axis)**
- - - - **Price according its fundamentals (right axis)**

Figure 6 shows historical series for real housing prices, comparing it with imputed prices when the wage bill is used as a variable for aggregate income. The results of using GDP as a variable of aggregate income are presented in the Annex. As can be seen, real housing prices fluctuate more than those justified by their domestic fundamentals.[8] This implies that more or less prolonged periods of significant under or overvaluation are occasionally observed. The latter could indicate the existence of housing price bubbles. In fact, in the quarters leading up to the crisis of 2002, the difference between real and fundamental housing prices was statistically significant.[9] On the other hand, during the most recent period, real housing prices fluctuated around fundamental prices, or were slightly undervalued as compared to said prices arising from the model.

Figure 7

INDEX OF PRICES ACCORDING THEIR FUNDAMENTALS AROUND THE 2002 CRISIS
(average 2000 =100)



8 See Hott (2009) for an analysis of the reasons why real housing prices fluctuate more than fundamental prices.

9 The confidence interval of 95% was calculated under the assumption that the difference between real and fundamental prices follows a normal probability distribution. The data does not allow for rejecting this hypothesis for the series calculated using gross domestic product as a variable of aggregate income.

In every case there is a statistically significant misalignment of real prices from fundamental prices under the environment of the 2002 crisis. The fundamental prices that emerge from the model show a downward trend starting around four years prior to the outbreak of the crisis determined by a fall in aggregate income variables.[10] The speed of the decline in fundamental prices intensifies during the year immediately before the crisis due to the deepening of the recession the Uruguayan economy was undergoing (see Figure 7). Meanwhile, real prices exhibit downward rigidity. The latter could be explained by the structural characteristics of the Uruguayan economy and its real estate market. For instance, housing is offered and sold in United States dollars, which, under the framework of the fixed exchange rate regime in force at the time, might have meant that adjustments were made more by quantities than by prices.

Although these results only derive from the crisis event, they argue in favor of using fundamental prices as a forward-looking indicator of financial fragility. Moreover, measures that favor the downward flexibility of prices and the de-dollarization of real estate transactions would have to be macroprudential in nature to facilitate smoother corrections in misalignments.

6. FINAL REMARKS

A comparison of real housing prices with prices that can be explained by economic fundamentals is important from the point of view of financial system stability. Prolonged periods of housing prices that are above those justified by fundamentals can lead to over-indebtedness, excess risk taking and the overvaluing of guarantees, all of which make sudden and costly corrections more likely.

In the case of Uruguay, no statistically significant differences were found between the real price of housing and that

[10] For instance, gross domestic product fell by around 10% between 1998 and 2001.

calibrated for the recent period by the model of fundamentals. It is also important to mention that fundamental prices calibrated in this paper follow, in the recent period, a stable growth trend. The same estimated prices began to follow a downward trend four years before the outbreak of the 2002 crisis, a trend that intensified in the year preceding said crisis where the difference between real prices was statistically significant.

In general terms, this paper found evidence that housing prices in Uruguay were overvalued with respect to those justified by fundamentals immediately before the 2002 crisis. On the other hand, no significant evidence is found in the most recent period. In any case, the aforementioned does not mean housing prices cannot fall in the future or even that their fundamentals are overvalued due to the particular conditions prevailing in international markets. These matters, as well as a more detailed analysis of the exposure of financial intermediaries to housing prices, should be the subject of future study.

ANNEXES

Description of the Variables

- P_t *Real housing price index*: real housing prices compiled by Carlomagno and Fernández (2007), updated by the author to the second quarter of 2011. Up until 1998, the variable corresponds to the price per square meter obtained from real estate sales advertisments published in national newspapers. The index is compiled following the methodology proposed by Grau et al. (1987). As of 1999, the variable corresponds to the price of actually registered transactions. This is compiled by the National Institute of Statistics.

- M_t *Real rent index*: Rent component of the consumer price index compiled by the National Institute of Statistics based on surveys of real estate agents and the rental service of the General Accounting Office.

- Y_t *Aggregate income*: Uses three aggregate income variants: *1)* The physical volume of seasonally adjusted gross domestic product obtained from the Banco Central del Uruguay; *2)* the trend through Hodrick-Prescott filter of the physical volume of gross domestic product obtained from the Banco Central del Uruguay; *3)* the wage bill calculated as real average wages multiplied by the number of workers, both obtained from the National Institute of Statistics.

- m_t *Real interest rate*: The interest rate calculated by discounting the inflationary component (consumer price index calculated by the National Institute of Statistics) from a series for the average asset interest rate of the banking system calculated as the ratio between yields from interest and total loans of the system to the non-financial sector (source: Banco Central del Uruguay).

- I_t Gross physical capital formation in residential buildings compiled by the Banco Central del Uruguay.

Figure 8

IMPUTED INCOME (Y_t = SEASONALLY ADJUSTED GDP), RENTAL INDEX AND DIFFERENCE

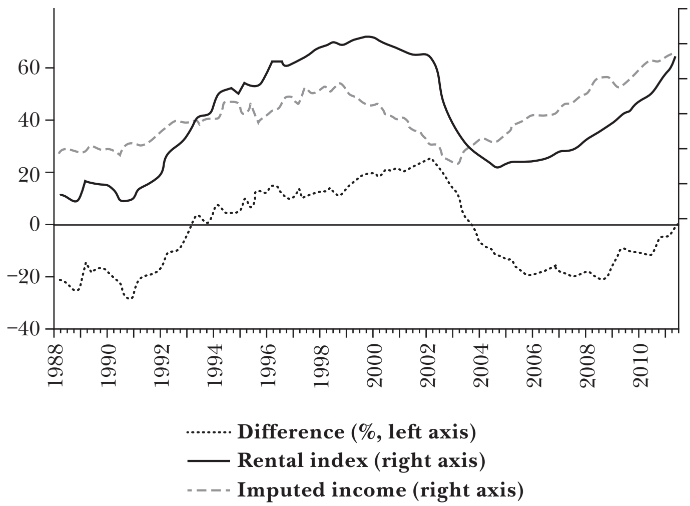


Figure 9

PRICE ACCORDING ITS FUNDAMENTALS (Y_t = SEASONALLY ADJUSTED GDP), REAL PRICE AND DIFFERENCE

(in gray: interval of confidence at 95%)

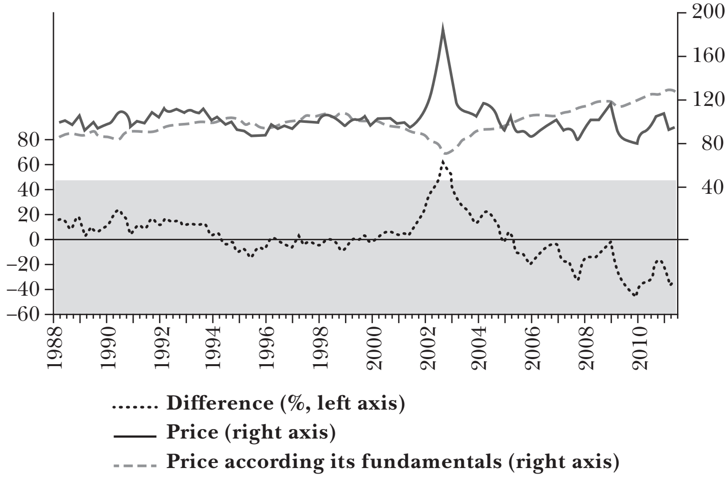


Figure 10

IMPUTED INCOME (Y_t = TREND GDP), RENTAL INDEX AND DIFFERENCE

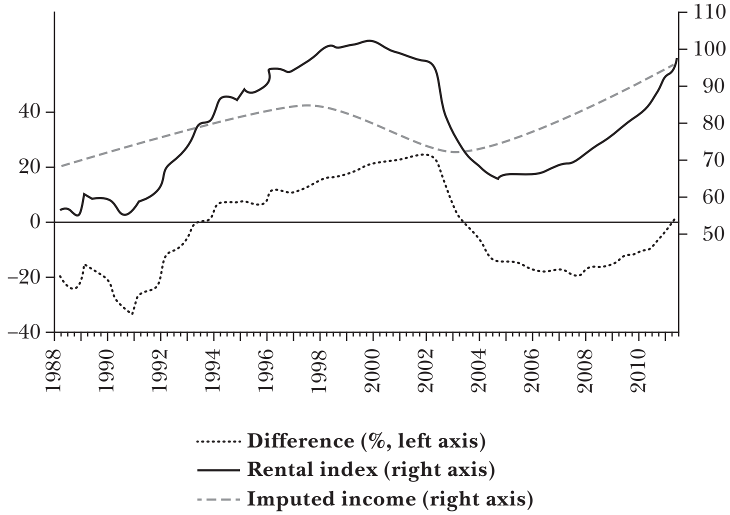
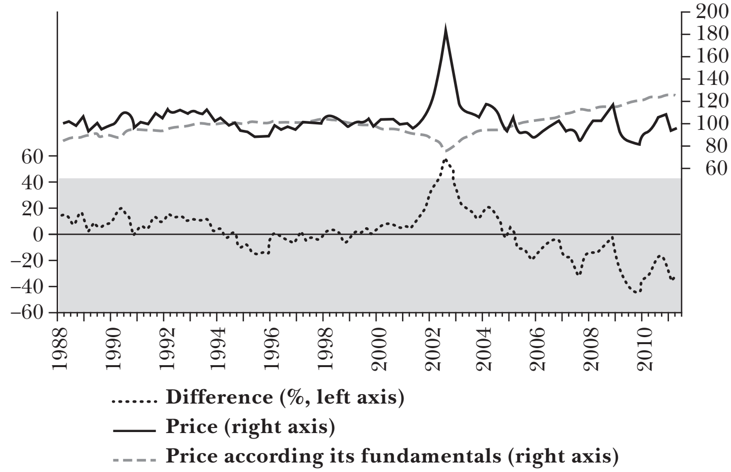


Figure 11

PRICE ACCORDING ITS FUNDAMENTALS (Y_t = SEASONALLY ADJUSTED GDP), REAL PRICE AND DIFFERENCE
(in gray: interval of confidence at 95%)



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Luis Eduardo Arango
Ximena Chavarro
Eliana González

Commodity Price Shocks and Inflation within an Optimal Monetary Policy Framework: The Case of Colombia

Abstract

We estimate food, oil and energy price effects on inflation in a small-open-economy model for Colombia. Such an economy exports and imports commodities and has an inflation-targeter central bank who follows an optimal interest rate rule. We found evidence of small effects of commodity prices shocks on headline inflation once the reaction of monetary authority has been taken into account. Thus, our interpretation is that monetary authority has faced rightly the shocks to commodity

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prices. Inflation expectations are the main determinant of inflation during the inflation targeting regime. Commodity prices movements are to a great extent included in the information set to form expectations.

Keywords: commodity prices, inflation-targeting regime, optimal monetary policy, expectations.

JEL classification: E43, E58.

1. INTRODUCTION

The behavior of commodity prices is matter of permanent concern among its producers, investors, policymakers and economists. The reason is that commodity prices changes potentially may bring about new economic conditions and give signals on the future path of some relevant domestic macroeconomic variables. That is the case of inflation in countries where commodity prices shocks represent important sources of either demand or supply pressures. In consequence, the question whether monetary authorities should react to commodity prices fluctuations effects on domestic inflation is not trivial.

Output disturbances are different in nature and require different policy responses. The basic theory on monetary economics suggests that monetary authority should offset demand shocks but accommodate supply shocks (Clarida, Galí, and Gertler, 1999). Thus, identifying the nature of shocks is just one of the tasks to which monetary authorities are faced to (see Uribe, 2010). In this sense, it is necessary to gauge the magnitude of the effects: If there a significant long-lasting impact exists, then an adequate response will have to be implemented. For example, fluctuations of the commodity exports prices that lead to reactions of aggregate national income may represent an important source of inflation due to demand pressures in countries where these products are the core of the economic activity (IMF, 2008). However, if the country is net importer of commodities the policy reaction could be different depending on the pass-through from import prices to inflation.[1]

[1] Some standard small open economy models link the inflation impact

Regardless of the apparent importance of commodity prices shocks on inflation, there is no much research devoted to the study and estimation of this phenomenon by invoking an economic model in which the inflation process can be derived as we do below. Most research is carried out in the spirit of either general equilibrium (Medina and Soto, 2007) or empirical models (Pedersen and Ricaurte, 2014). In a recent work, Jalil and Tamayo (2011) estimated first and second round effects of food international prices on inflation of Brazil, Chile, Colombia, Mexico and Peru. The authors found that, for Colombia, the effects of commodity price shocks disappear four months after the shock, estimating an elasticity of 0.27 of domestic prices to the international prices. When inflation is decomposed into core inflation without food and food price changes, the elasticities are, on average, 0.194 and 0.477, respectively. With respect to the second round effects, they provided evidence that the effects take place within a period close to four months, though the numerical magnitude is lower than 10% of the first round effects.

On the effects of commodity prices shocks on inflation and inflation expectations in Colombia, recently Arango, Chavarro and González (2013) found evidence of first-round and second-round effects between 1990 and 2010. Their empirical results showed that there is a positive and significant pass-through from food and oil international prices to the domestic prices of some selected items of the CPI and PPI baskets. Nevertheless, the magnitudes of the effects are small: They found an elasticity of domestic prices to international prices between 0.1 and 0.3 on average.[2] The estimated effects on core inflation and inflation expectations are higher, especially in the case of food prices shocks. In particular, a 1% rise in the internation-

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of imports prices to the weight given to imports in the CPI (see for example, Galí et al., 2005), while others, such as McCallum and Nelson (2001), show that the transmission to inflation is limited to the extent to which relative price shocks affect aggregate supply.

[2] For items as cocoa, coffee, sugar, palm-oil, sun-flower oil and soy-oil the elasticity is higher than 0.5.

al price of food brings forth a rise of 0.56% on core inflation and explains about 32% of the changes on inflation expectations in one-month horizon, with an important decline on the latter when the time-horizon is extended. According to these authors, the reduction of the pass-through coefficient of food prices to core inflation since the inflation-target regime was established shows that there have been significant gains by controlling inflation.

However, the approaches of Jalil and Tamayo (2011) and Arango, Chavarro and González (2013) are empirical in essence. None of them present a theoretical setting in which the behavior of monetary authority within a proper framework to face shocks is explicit. This is relevant because Colombia's central bank follows an inflation-targeting strategy and is committed to control inflation to provide conditions for a sustainable economic growth path. In our view, the final pass-through from commodity prices shocks to domestic inflation should be analyzed taking into account the reaction function implicit in the monetary policy rule.

Accordingly, this article is aimed to determine how much of international commodity price shocks are passed through inflation under an optimal monetary policy framework. This subject is important for two reasons. On the one hand, Colombia is a commodity exporter hence changes in commodity world markets may have a direct impact on the economy through channels encompassing gross domestic product growth, exchange rate movements, financial (un)balances, inflation behavior and a higher exposure to the dynamics of aggregate demand in emerging and developed economies. On the other hand, it is worth to evaluate how an optimal monetary policy framework leads to a higher domestic price stability given the commodity prices movements.

The theoretical body we use is based on Walsh (2002) and De Gregorio (2007), which consists of a text-book model used to explain the inflation targeting strategy which is explained below.[3] In that sense, this paper might be though as an empirical attempt to verify the goodness of this simple model to explain the inflation determinants in a small open economy. The theoretical device is enriched with four shocks: Cost-push, demand, and two structural. The demand shock is attempted to capture the idea that movements in commodity prices (oil, coal, etc.) have effects mainly via aggregate demand rather than supply in the economy.[4] As we will see below, the success of the model is not complete tough auspicious. Moreover, the results suggests that commodity price shocks and other demand and supply shocks are of minor importance while expectations are the main determinant of inflation during the inflation targeting regime. Thus, we claim that the monetary authority has faced rightly the shocks to commodity prices during this regime.

The article develops in six sections of which this "Introduction" is the first. The second shows some facts of the recent behavior of commodity prices and inflation. The third section presents and explains the model and provides some intuition. The fourth section is devoted to explain the way in which structural shocks, commodity prices shocks and inflation expectations are obtained. Section fifth shows and discusses the results. Finally, the sixth section draws some conclusions.

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[3] This framework was also used by Vargas and Cardozo (2013).

[4] This conjecture is supported by two facts. First, Colombia is a net oil-exporter country (oil exports represented 34.3% of total exports on average between 2005 and 2010). In addition, the government is the major stockholder of the main oil firm in Colombia (revenues received by the government from oil are about 12% of total government income between 2007 and 2013). Second, the symptoms of Dutch decease undergone by the Colombian economy, associated to the well behavior of the terms of trade. In fact, the industrial sector maintained a 3.1% annual growth rate between 2000 and 2013, interrupted by the crisis occurred in 2008 and 2009 when the growth was 0.6% and −4.1%, respectively. In the aftermath, the annual growth rate of the sector was 1.8%, well below the whole economy (4.4%).

2. SOME FACTS ON RECENT BEHAVIOR OF COMMODITY PRICES

Figure 1 shows nominal food and crude oil price indexes.[5] After a period of relative stability, during the past decade the index of food world prices grew up by more than 110% between January 2000 and December 2013 and by more than 339% in the case of crude oil during the same period. In real terms, between January 2000 and December 2013, the percentage variations of food and crude oil prices were of 52% and 218%, respectively, while from January 1990 to December 1999 the registered percentage variations were of −41% and −9%, respectively. All these price movements, as has been argued in previous research (see Frankel, 2006; Bernanke, 2006), are consequence of both supply and demand shocks. On the one hand, an increasing demand for commodities by large emerging economies as China and India has soared commodity prices. The transition to other types of energy, in particular an increasing demand for biofuels, has raised the price of land and, in turn, increased the cost of food production. Financial developments in commodities markets, climate phenomena and supply shocks in the crude oil market are also among the reasons that explain the upsurge in commodity world prices.

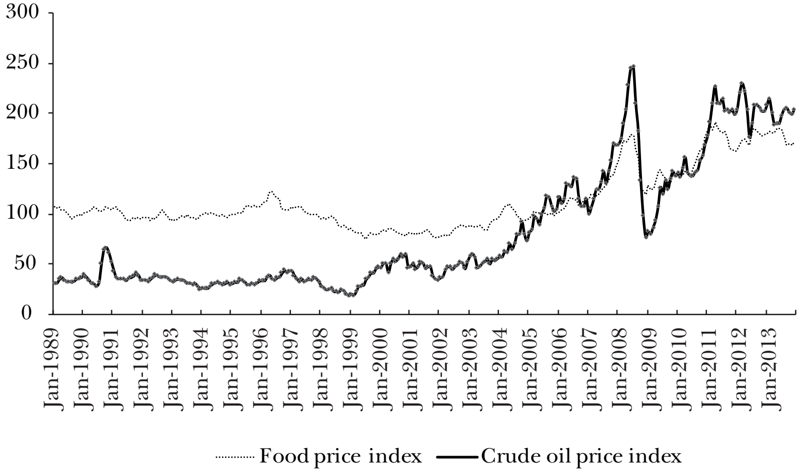
Potentially, commodity price booms bring about first and second round effects on inflation. The former consist of primary or direct effects while the latter are linked to a rise in underlying inflation. This is the case when the increases in food and fuel prices drive up expectations and underlying inflation producing further price increases and demand for higher wages. This is especially important for those economies where commodities account for a large share of final expenditure and where monetary policy has only limited credibility. To the extent that commodity prices shocks are large and persistent, inflation risks increase and second

[5] Our reference prices are food and crude oil international price indexes from the IMF.

Figure 1

FOOD AND CRUDE OIL PRICE INDEXES

(base: Jan-2005=100)



Sources: IMF, National Administrative Department of Statistics (DANE), and authors' calculations.

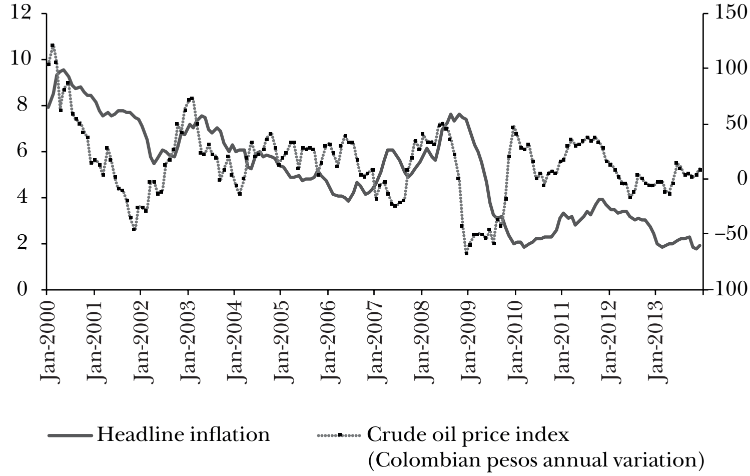
round effects arise, requiring an accurate and timely policy response. In other words, if shocks to commodity prices are transitory, they are expected to damp in the short run with persistent effects on neither expectations nor underlying inflation. In contrast, with large and persistent shocks, as were the cases of recent episodes, monetary authorities face a challenge since the shocks are transmitted to inflation expectations and prices of other goods and services in the economy (Bernanke, 2006).

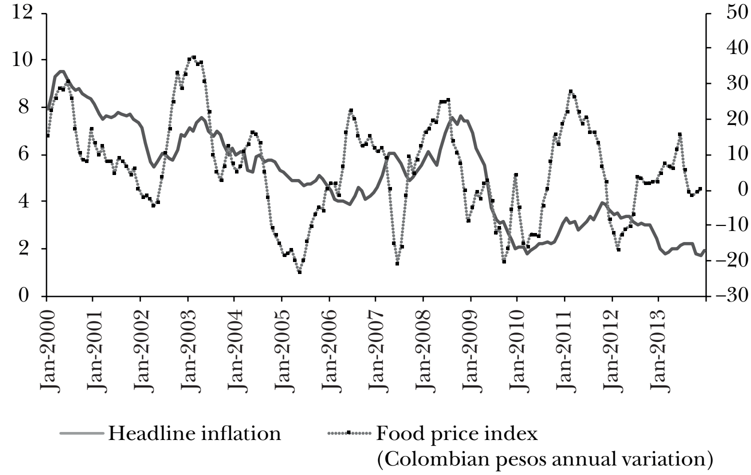
Figure 2 presents some of the variables we use to analyze the effect of commodity price shocks on inflation in Colombia; later we shall introduce the expectations processes. The upper panel presents the headline inflation and the annual variation of crude oil (LHS) and food international prices (RHS), both expressed in Colombian pesos. Two things are worth highlighting. First, there is some coincidence between

Figure 2

COMMODITY PRICES ANNUAL VARIATION, HEADLINE INFLATION AND POLICY RATE

(percent)



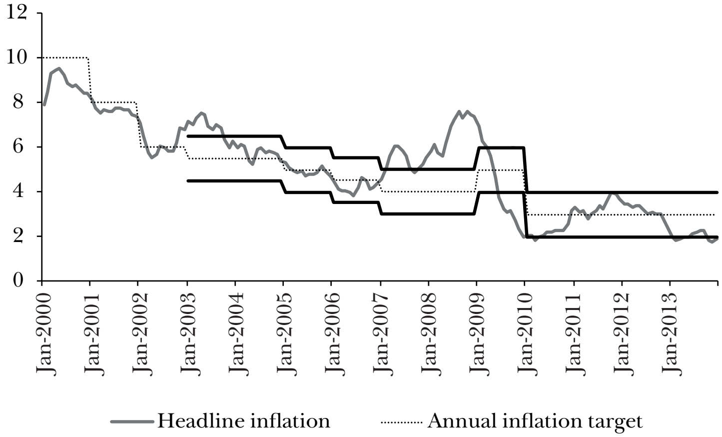


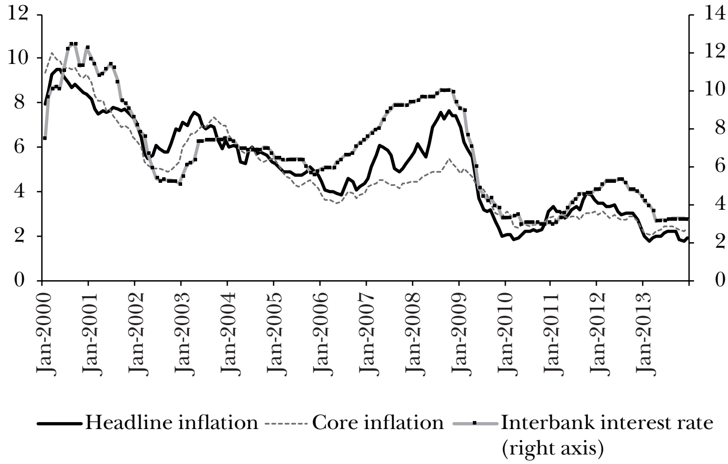
Note: the right axis measures the annual percentage change of international prices.
Source: IMF; National Administrative Department of Statistics (DANE); and author's calculations.

Figure 2

COMMODITY PRICES ANNUAL VARIATION, HEADLINE INFLATION AND POLICY RATE

(percent)





Note: the right axis measures the interbank interest rate.
Source: IMF; National Administrative Department of Statistics (DANE); and author's calculations.

international commodity price movements and domestic headline inflation during the same period.[6] Second, the more recent developments of inflation in Colombia would suggest an effective reaction of monetary authority, as new increases in crude oil and food world prices did not have an impact on inflation which has been maintained within the rank-target (see the LHS lower panel). The lower panel (RHS) of Figure 2 shows the behavior of policy rate (measured by interbank interest rate) and headline inflation and core inflation as measured as inflation without food and regulated prices.

3. THE MODEL

As we mentioned before, the aim of this paper is to analyze the effects of commodity prices shocks on domestic inflation in Colombia based on a model in which monetary authority reacts to deviations of output from its flexible price equilibrium level, deviations of the inflation rate from its target and deviations of real exchange rate from its long run equilibrium value. We formulate a data generating process for inflation,

[6] For example, in 2007, inflation in Colombia reached a level of 5.69%, surpassing the upper limit of the inflation target by 119 basis points. This, as pointed out by the monetary authority of Colombia (see, inflation report December 2007), was mainly due to a food inflation higher than expected with world commodity prices explaining a large part of this increase. At the end of 2008, inflation in Colombia jumped to 7.67%, missing this time the upper limit of the target by 317 basis points. Once again, the monetary authority of Colombia argued that high international oil prices and other commodity prices, not only created inflation pressures on domestic food and fuel prices, but also had a considerable impact on inflation expectations. The length of increasing international commodity prices and its impact on expectations for further prices increases and total inflation was underestimated by some central banks, a situation that also apparently occurred in Colombia. As we will see below, there is no evidence to reject such statement if we include the permanent component of commodity prices into the picture.

π_t, in the context of a small open economy with an inflation-targeting strategy, and estimate the effect on this of commodity prices (demand) shocks,[7] inflation expectations, cost-push shocks,[8] and two structural shocks.

The framework is based on two relations: An expectations augmented Phillips curve and a description of monetary policy behavior, reflecting the policymaker's preferences in trading off output gap, inflation and exchange rate deviations (see Walsh, 2002). The latter implies a central bank that sets its policy instrument to stabilize inflation, output gap and exchange rate. A monetary policy rule (MPR thereafter) emerges when monetary authority balances the marginal costs and benefits of its policy actions. In other words, the MPR shows a relation between the output gap, misalignments of exchange rate, and deviations of inflation from the target consistent with a monetary authority designed to minimize the costs of output and inflation variability.[9]

The MPR shows the reaction of the monetary authority. Once this authority observes the current inflation, decides optimally on the size of the output gap and the exchange rate deviation. In this set up, the long-run equilibrium occurs when the output gap equals zero, current inflation equals the central bank's target and the exchange rate is on the long-run level. As pointed out by Gertler, Galí and Clarida (1999), the policy design problem consists of characterizing how the interest rate should adjust to the current state of the economy.

7 We refer here to shocks to the international prices of oil and energy.

8 As we explain below, a fraction of these corresponds to shocks to the international price of food.

9 Nevertheless, factors other than systematic monetary policy influence aggregate demand and output in ways that monetary authority cannot perfectly foresight. In addition, policymakers may have goals beyond inflation and output gap stabilization that would shift the relation between the output gap and inflation described by the monetary policy rule. A random disturbance variable denoting the net impact on output of those additional factors can then be added to the model.

We follow De Gregorio (2007) and Walsh (2002) to lay out a model consisting of three basic equations: An expectations-augmented Phillips curve, which schedules the aggregate supply of the economy, an IS-type aggregate demand and a MPR derived below from the objective function of a monetary authority assumed to follow an optimum rule.

The Phillips curve and the IS curve are given by:

1 $$\pi = \pi^{e} + \theta\left(y - \overline{y}\right) + \delta\left(q - \overline{q}\right) + \omega\varepsilon^{food} + \varepsilon ,$$

2 $$y - \overline{y} = A - \varphi\left(i - \pi^{e}\right) + \alpha q + \rho\mu^{crude\ oil} + \mu ,$$

where, π stands for the annual rate of inflation, π^{e} are the inflation expectations, y is the output, $\overline{y}$ is the flexible price equilibrium output level, q stands for the real exchange rate, $\overline{q}$ is the long run value of the real exchange rate, A is a composite factor accounting for autonomous spending, i is the nominal interest rate, $\overline{\pi}$ is the annual inflation target, ε^{food} and $\mu^{crude\ oil}$ are the components of commodity prices orthogonal to expectations mechanisms, ε and μ stand for cost-push and demand shocks, respectively, and θ , δ , φ , ω , α , ρ are unknown parameters.

According to Equations 1 and 2, a fraction of cost-push and demand shocks is strictly related to pressures coming from shocks to international food and crude oil prices respectively. Thus, both ε and μ are residuals of a regression of each on both ε^{food} and $\mu^{crude\ oil}$. The way in which ε , μ , ε^{food} and $\mu^{crude\ oil}$ are identified and obtained is explained below.

Following De Gregorio (2007), the optimization problem faced by monetary authority can be written as: $\min \lambda(y - \overline{y})^{2} + (\pi - \overline{\pi})^{2} + \beta(q - \overline{q})^{2}$ subject to 1 and 2. The loss function accounts for deviations of output from its flexible price equilibrium level, inflation rate from its target and deviations of real exchange rate from its long run equilibrium value. The model also includes an uncovered interest parity condition, $r = r^{*} + \overline{q} - q$, and the Fisher equation, $i = r + \pi^{e}$.

From the first order conditions of the optimization problem, the MPR is given by

$$\pi - \bar{\pi} = -\left(\frac{\alpha\lambda}{\alpha\theta+\delta}\right)(y-\bar{y}) - \left(\frac{\beta}{\alpha\theta+\delta}\right)(q-\bar{q}) . \tag{3}$$

This curve reflects the trade-off faced by monetary authority in terms of keeping inflation, output and real exchange rate as close as possible to their target or equilibrium levels. After replacing MPR in the Phillips curve and the IS curve, we find the optimal interest rate rule (IRR),

$$i = \bar{i} + \left(1 + \frac{\theta\alpha+\delta}{\upsilon}\right)\left[\pi^{e} - \bar{\pi}\right] + \frac{\theta(\alpha\theta+\delta)+\alpha\lambda}{\upsilon}(\rho\mu^{crude\ oil} + \mu) + \\ + \frac{(\alpha\theta+\delta)}{\upsilon}(\omega\varepsilon^{food} + \varepsilon), \tag{4}$$

where $\bar{i} = r^{*} + \bar{\pi}$ and $\upsilon = \alpha^{2}\lambda + \varphi\alpha\lambda + \beta + (\theta\alpha+\delta)(\delta+\alpha\theta+\varphi\theta)$.

The IRR shows the reaction of the monetary authority when inflation expectations are different from the target, or commodity price, demand or other cost-push shocks reveal. It is evident that the higher the value of β, the less the reaction of monetary authority to shocks or expectations.[10] Recall that parameter υ contains β, and the former appears in the denominator of each coefficient.

After some algebra manipulation, the inflation process can be written as

$$\pi = \left(\frac{\alpha^{2}\lambda+\varphi\alpha\lambda+\beta}{\upsilon}\right)\pi^{e} + \left[1 - \left(\frac{\alpha^{2}\lambda+\varphi\alpha\lambda+\beta}{\upsilon}\right)\right]\bar{\pi} + \\ + \left(\frac{\beta\theta-\delta\alpha\lambda}{\upsilon}\right)(\rho\mu^{crude\ oil} + \mu) + \left(\frac{\alpha^{2}\lambda+\varphi\alpha\lambda+\beta}{\upsilon}\right)(\omega\varepsilon^{food} + \varepsilon), \tag{5}$$

[10] Parameter β represents the weight of the deviations of real exchange rate from its long run value.

where the sources of inflation process in this model become clear. In first place, we observe that the higher the expectations the higher the annual inflation. At the same time, positive realizations of, almost, all shocks will render a higher inflation rate. However, in the case of shocks to oil prices or structural demand, the inflation reaction will be different depending on which force is greater either $\beta\theta$ or $\delta\alpha\lambda$. The former parameters represent, on the one hand, the weight of deviations of exchange rate from its long run value in the loss function and the parameter linked to the marginal cost in the Phillips curve. On the other hand, the parameters, $\delta\alpha\lambda$, represent the contribution to inflation of deviations of exchange rate from its long run value in the Phillips curve, the coefficient of real exchange rate in the IS equation, and the weights of the gap in the loss function, respectively. Thus, if the value of the product $\beta\theta$ is greater (less) than the product $\delta\alpha\lambda$, any positive shock to oil prices or a demand shock will increase (reduce) inflation. In particular, if the monetary authority expresses a concern on the deviations of the real exchange rate from its long run value, the increase of the nominal interest rate will be less than otherwise. In the extreme case that the monetary authority does not express any concern at all about the real exchange level in the loss function $(\beta = 0)$, any shock to oil prices or demand shock will drive to a reduction in inflation given the reaction condensed in the IRR.

The inflation process can also be written as:

$$\pi - \bar{\pi} = \left(\frac{\alpha^2\lambda + \varphi\alpha\lambda + \beta}{\upsilon}\right)(\pi^e - \bar{\pi}) + \left(\frac{\beta\theta - \delta\alpha\lambda}{\upsilon}\right)(\rho\mu^{crude\ oil} + \mu) + \left(\frac{\alpha^2\lambda + \varphi\alpha\lambda + \beta}{\upsilon}\right)(\omega\varepsilon^{food} + \varepsilon), \quad (6)$$

which is the equation we actually estimate. In essence, it shows that, within this economic framework, deviations of inflation from the target are caused by deviations of inflation expectations from the target, and shocks to commodity prices, demand and other cost-push shocks. In the next section we show how some variables included in the model are built.

4. DATA: COMMODITY PRICE SHOCKS, INFLATION EXPECTATIONS, MONTHLY INFLATION TARGET AND SUPPLY AND DEMAND SHOCKS

The estimation of Equation 6 requires some data that are not readily available. That is the case of inflation expectations mechanisms, trajectories of commodity price shocks orthogonal to inflation expectations, monthly series of inflation target[11] and identified structural shocks. We now consider each in turn.

4.1 Expectations Mechanisms

The first mechanism we use to measure inflation expectations is the break-even inflation (BEI from now on) which uses information of secondary market of debt in Colombia. In this case, by invoking Fisher equation, inflation expectations are computed at different horizons as the difference between the nominal yield of fixed income bonds and the real yield on inflation linked bonds, both issued by the government. Accordingly, BEI rates from 1-year and 2-years yield curves stand for our market-based measures of inflation expectations. The second indicator of inflation expectations is the forward BEI which in essence derives expectations from one and two years BEI forward curves and reflects the expected inflation in a year time for the following year.

Finally, inflation expectations are also obtained by assuming that agents form their expectations about future inflation, one and two-years ahead (s = 12, 24 months), according to a specific process. To this aim, we use an imperfect rational expectations mechanism, wich is a moving average process given by:

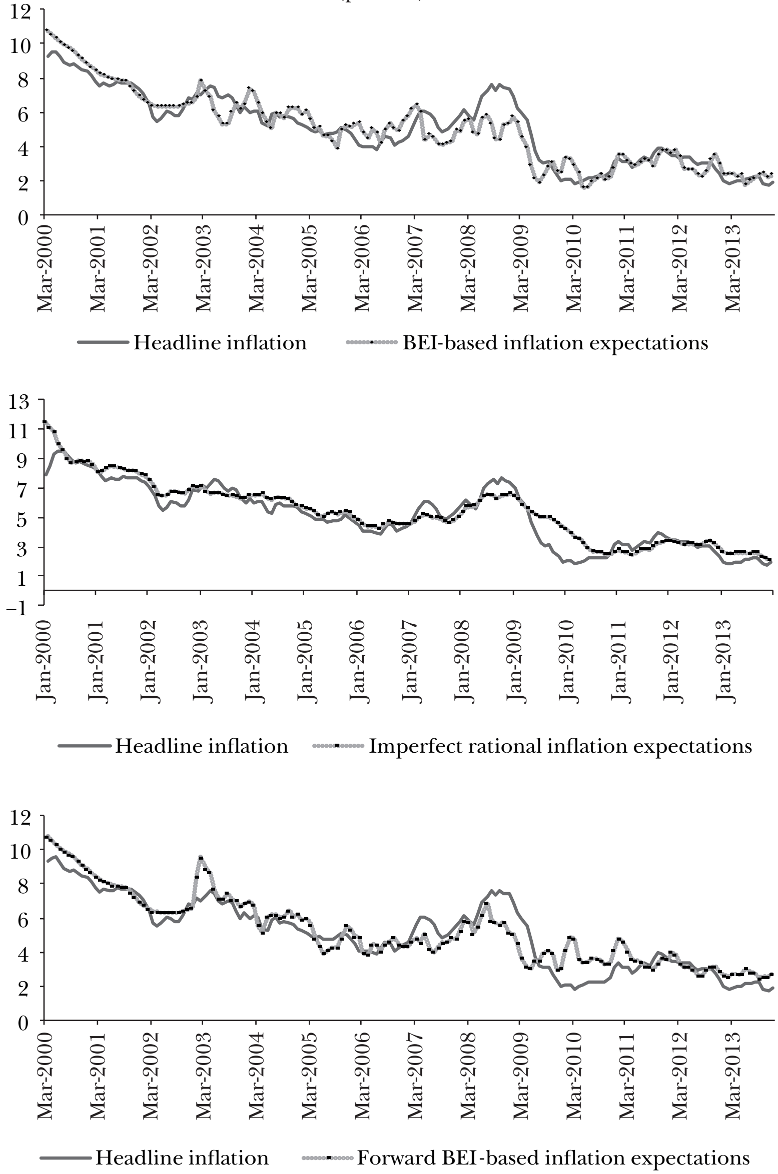
$$\pi^{e}_{t-S,t} = \kappa\pi^{headline}_{t-S} + (1-\kappa)\pi^{headline}_{t}, \tag{7}$$

[11] This is so, because we use monthly data on the estimation. Then, we need an estimation of the monthly target of inflation in order to calculate the deviation of monthly inflation rate from its implicit target.

Figure 3

PANEL A. INFLATION EXPECTATIONS ONE YEAR AHEAD

(percent)

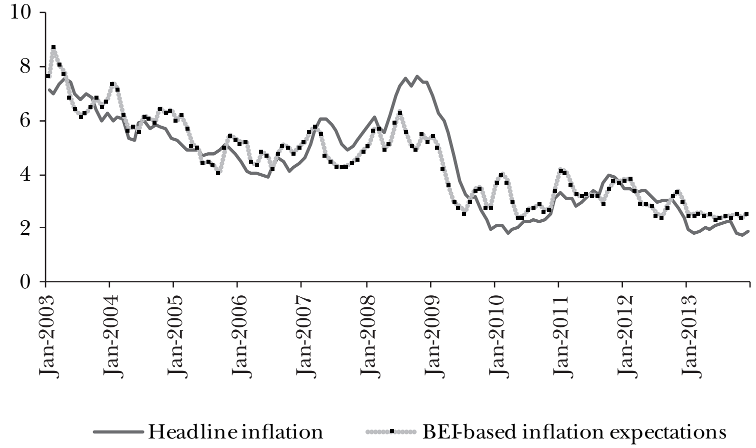


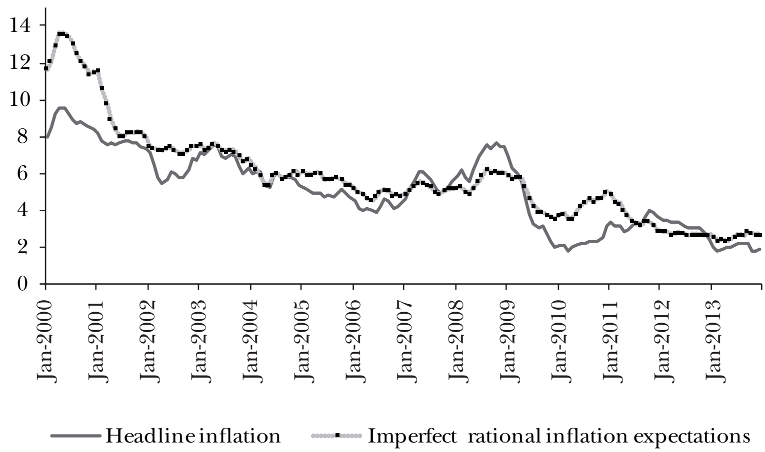
Sources: DANE, Banco de la República (Colombia), and authors' calculations.

Figure 3

PANEL B. INFLATION EXPECTATIONS TWO YEARS AHEAD

(percent)





Sources: DANE; Banco de la República (Colombia); and authors' calculations.

where both headline inflation and inflation expectations correspond to annual inflation rates. This mechanism is based on the hypothesis that agents assign a weight κ [12] to inflation observed s periods ago and a $(1-\kappa)$ weight to current inflation, when predicting how inflation is expected to behave in the future. This mechanism supports the inertia property of the inflation expectations process. Figure 3 presents the relation between headline inflation and inflation expectations. Figures are depicted accounting for the fact that agents form their inflation expectations in advance, so that current inflation outcomes are related to their corresponding inflation expectations formed 12 and 24 months ago.

When expectations are imperfeclyt rational, the inflation process can be written as:

8
$$\pi-\bar{\pi}=\frac{\kappa\left(\alpha^2\lambda+\varphi\alpha\lambda+\beta\right)}{\upsilon-(1-\kappa)\left(\alpha^2\lambda+\varphi\alpha\lambda+\beta\right)}(\pi_{-s}-\bar{\pi})+$$
$$+\frac{\beta\theta-\delta\alpha\lambda}{\upsilon-(1-\kappa)\left(\alpha^2\lambda+\varphi\alpha\lambda+\beta\right)}\left(\rho\mu^{crude\ oil}+\mu\right)+$$
$$+\frac{\left(\alpha^2\lambda+\varphi\alpha\lambda+\beta\right)}{\upsilon-(1-\kappa)\left(\alpha^2\lambda+\varphi\alpha\lambda+\beta\right)}(\omega\varepsilon^{food}+\varepsilon).$$

4.2 Commodity Prices Shocks

In assessing the pass-through of commodity price shocks to inflation is crucial to define what a price shock is. To this end, we use the Hodrick-Prescott filter to decompose the annual variation of commodity prices between permanent and transitory components, being the latter the unexpected one.

Accordingly, we have chosen three commodity price indexes from the IMF Primary Commodity Prices: crude oil, energy and food. On the demand side, we consider shocks to oil and energy

[12] The value of κ we use is 0. 44.

price fluctuations and denote them by $\mu^{crude\ oil}$ (which include a simple average of three spot prices: Dated Brent, West Texas Intermediate and the Dubai Fateh) and μ^{energy} (which include petroleum, natural gas and coal price indexes), respectively, while shocks to food price fluctuations are denoted by ε^{food}. All these shocks should satisfy the key restriction of being orthogonal to the inflation expectations mechanisms defined above, an assumption that we test. We also test the assumption that inflation expectations should be correlated to the permanent (long-run) component of crude oil, energy and food prices but no to the transitory (cyclical) component.

To verify the latter assumption, we calculate the Pearson correlation coefficients between each of the expectations mechanisms and the permanent and transitory components of commodity prices fluctuations. Table 1 shows the estimated correlation coefficients between each expectations mechanism and both the permanent and transitory components of annual variation of commodity prices, denominated in dollars. For all the mechanisms, inflation expectations are correlated with the long run component of commodity prices annual variation.

The imperfect rational expectations mechanism shows the lower correlation coefficient while expectations as predicted by BEI mechanism show the highest correlation. The cyclical components of annual variation in commodity prices show no correlation with inflation expectations, with the only exception of forward BEI mechanism. However, we will consider that, in general, transitory components of international commodity prices are not correlated with inflation expectations in Colombia.[13]

[13] We also run regressions of inflation expectations on permanent and temporary components of commodity prices indexes. The results are similar to those of Table 1; that is, in general, permanent components of commodity prices explain inflation expectations while temporary components do not. The results are not shown but are available upon request.

Table 1

CORRELATION OF INFLATION EXPECTATIONS WITH PERMANENT AND TRANSITORY COMPONENTS OF ANNUAL VARIATION OF COMMODITY PRICES DENOMINATED IN DOLLARS

	One year ahead			*Two years ahead*	
Price	*BEI*	*Forward BEI*	*Imperfect rational*	*BEI*	*Imperfect rational*
Crude oil					
Permanent	0.36[a]	0.28[a]	0.20[a]	0.47[a]	0.29[a]
p-value	0.00	0.00	0.01	0.00	0.00
Transitory	0.04	0.13	0.03	0.09	0.10
p-value	0.57	0.10	0.67	0.30	0.19
Energy					
Permanent	0.42[a]	0.33[a]	0.27[a]	0.53[a]	0.34[a]
p-value	0.00	0.00	0.00	0.00	0.00
Transitory	0.06	0.15	0.04	0.13	0.11
p-value	0.41	0.05	0.59	0.14	0.14
Food					
Permanent	−0.38[a]	−0.42[a]	−0.47[a]	0.44[a]	−0.54[a]
p-value	0.00	0.00	0.00	0.00	0.00
Transitory	0.12	0.17[b]	0.00	0.12	0.09
p-value	0.12	0.03	0.96	0.17	0.26

Note: Numbers correspond to the Pearson correlation coefficient, with the associated *p*-value below. [a] represents significance at 1%, and [b] at 5 per cent.

4.3 Monthly Target for the Annual Rate of Inflation

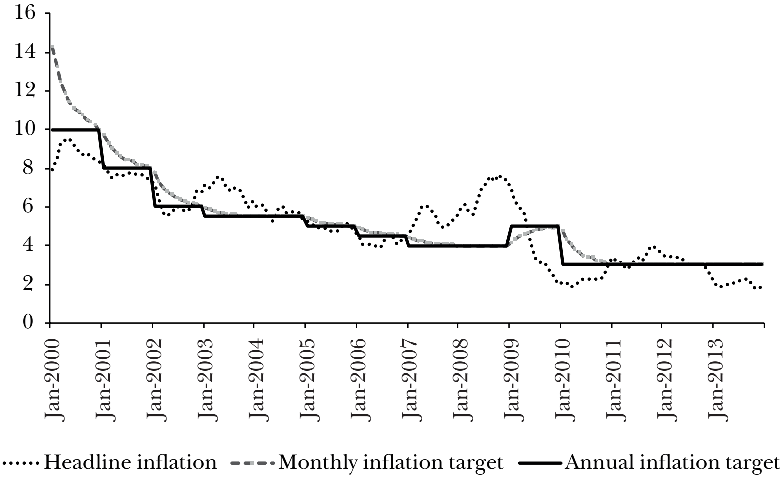
Our estimations of the inflation process require having a monthly-basis adjusted value of the target for the annual rate of inflation, a variable that is not available. To do so, we first establish a criterion to determine whether the target was reached or not in a sample of 22 years from 1991 to 2013. Thus, we calculate the ratio of the observed inflation rate to its target value and evaluate whether this ratio exceeds or falls behind a maximum level; we denote this value by g.[14] Essentially, we are estimating a monthly target based on the path of years for which target has been reached (see Arango, García and Posada, 2013).

Once we established the number of years for which target was attained, we calculate the average contribution of each month of the year to the annual rate of inflation. This is, in a year in which the target was hit, on average how much of the annual inflation rate was reached on January, how much on February, and so on, until the last month of the year. We obtain an average monthly contribution to hit the target from the sample of years matching the criterion. Then, we use these contributions and the corresponding target for every year of the sample to calculate, in a monthly basis, the inflation target. Figure 4 shows our monthly-basis adjusted target along with inflation rates observed from 2000m01 to 2013m12.

Figure 4

MONTHLY TARGET OF THE ANNUAL RATE OF INFLATION

(percent)



Sources: DANE, and authors' calculations based on Arango et al. (2013a).

[14] We fixed g equal to 0.05 and found a total of seven years in which this criterion is met.

4.4 Structural Shocks

Apart from the commodity prices shocks, the model also includes two additional shocks which are plugged into the Phillips curve and IS curve. The first is a cost-push (supply) shock and the second a demand shock. The approach we follow to obtain the set of structural supply and demand shocks is based on the estimation of a structural VAR model for the price level and output, which is derived from a basic aggregate demand-aggregate supply model, AD-AS. The set of structural shocks is obtained by using the Cover, Enders and Hueng (2006) approach (CEH thereafter), in which the usual long run restrictions are imposed to identify the shocks. In particular, CEH suggest that the aggregate demand shock has no long-run effect on real output. This approach, besides the long-run neutrality condition, allows some correlation between the demand and supply shocks. More generally, CEH do not impose any constraints to the variance-covariance matrix of structural shocks. Instead, they impose the normalization restrictions usually suggested in an AD-AS model: One-unit supply shock shifts AS by one unit and the effect of one-unit demand shock is also one-unit over AD (see Appendix 1 for details).

According to the authors, there are several arguments to justify the contemporaneous correlation between supply and demand shocks. On the one hand, monetary or fiscal policy may react according to current and past state of economic activity. On the other hand, from the new Keynesian point of view, some firms increase output, rather than prices, in response to a positive demand shock. Finally, to obtain supply and demand shocks orthogonal to commodity prices shocks, regressions of the former on the latter are estimated and the residuals from that regression are the shocks that enter into the inflation model. However, given that results remain the same with and without orthogonal shocks, we decided to maintain the original structural shocks.

5. ESTIMATION AND RESULTS

Estimations were performed using the time series of headline and core inflation, commodity prices denominated in dollars and in local currency (Colombian pesos)[15] and both monthly and annual inflation target. The results are also presented for the different expectations mechanisms and combination of commodity shocks. In the case of BEI and forward BEI expectations, the samples go from March 2000 to December 2013 with expectations one year ahead and from January 2003 to December 2013 with imperfect rational expectations. Estimations were performed for the whole inflation targeting regime period: January 2000 to December 2013 and two subsamples: from January 2000 up to December 2006, which corresponds to the period before the commodity boom, and from January 2007 to December 2013.

According to the results in Table 2, with commodity prices denominated in dollars, there is evidence of effects in the deviation of the observed headline inflation with respect to the monthly inflation target of oil and energy price shocks though the coefficients are rather small. In the two sub-periods these shocks are also significant when forward BEI expectations are considered. The coefficients have negative sign which could be suggesting that, in expression 6 $\beta\theta < \delta\alpha\lambda$ as an indication that real exchange rate deviations are not that important for the monetary authority if this result were due to a small value of β (recall this is the weight of real exchange rate in the loss function presented above). Effects on inflation derived from the structural shocks are not significant, except in the case of BEI and imperfect rational expectations. Even though this result holds for the whole period, it does not for the two subsamples since demand shocks seem to play a role for the inflation process between 2000 and 2006 but only in the case of BEI expectations. Interestingly, only expectations seem to be relevant for the whole period and the subsamples: All coefficients of the expectations processes are significant and have a positive sign.

[15] This is aimed to capture some masked effect that could be in place via the exchange rate.

Table 2

ESTIMATES USING HEADLINE INFLATION, MONTHLY TARGET AND ONE YEAR AHEAD EXPECTATIONS
(commodity prices denominated in dollars)

Expectation mechanism	*Constant*	*Expectation deviation*	*Demand-shock*	*Cost-push shock*	*Crude oil*	*Energy*	*Food*	*Number of observations*	*F*	*Adjusted R^2*
	Inflation targeting regime (January 2000 to December 2013)									
BEI	0.002	1.088[a]	0.001[b]	−0.001	−0.007		−0.007	166	18.70	0.630
	0.002	1.100[a]	0.001[b]	−0.001		−0.005	−0.010	166	18.06	0.624
Forward BEI	−0.001	1.075[a]	0.001	−0.001	−0.014[b]		−0.002	166	9.505	0.606
	−0.001	1.097[a]	0.001	−0.001		−0.014[b]	−0.005	166	8.836	0.597
Imperfect rational	−0.002	1.314[a]	0.000[b]	−0.000	−0.006		0.019	168	35.01	0.790
	−0.002	1.332[a]	0.000[c]	−0.000		−0.004	0.015	168	37.12	0.785
	January 2000 to December 2006									
BEI	−0.001	0.926[a]	0.001[a]	−0.001	−0.006[a]		−0.010	82	18.31	0.558
	−0.001	0.930[a]	0.001[a]	−0.001		−0.006[b]	−0.010	82	18.86	0.554
Forward BEI	−0.002[a]	0.778[a]	0.000	−0.000	−0.008[a]		−0.007	82	37.63	0.789
	−0.002[a]	0.784[a]	0.000	−0.000		−0.009[a]	−0.006	82	44.71	0.790

Imperfect rational	−0.002	1.110[a]	0.000	−0.000	−0.006		0.006	84	15.70	0.779
	−0.002	1.122[a]	0.000	−0.000		−0.005	0.005	84	14.81	0.774
				January 2007 to December 2013						
BEI	0.005[b]	1.183[a]	0.001	0.002	−0.008		−0.005	84	21.68	0.692
	0.005[b]	1.204[b]	0.001	0.002		−0.004	−0.013	84	19.36	0.687
Forward BEI	−0.001	1.552[a]	0.001	−0.001	−0.019[a]		−0.005	84	35.30	0.633
	−0.001	1.630[a]	0.001	−0.001		−0.021[b]	−0.008	84	35.28	0.629
Imperfect rational	−0.002	1.507[a]	0.001	0.000	−0.010		0.028	84	64.92	0.812
	−0.002	1.533[a]	0.001	0.000		−0.007	0.021	84	60.00	0.804

[a] represents significance at 1%, [b] at 5% and [c] at 10% based on Newey-West standard errors.
Source: authors' calculations.

Table 3 shows the estimates of specification 6 with commodity prices denominated in local currency. Under this specification the coefficients of commodity price shocks (oil and energy) are marginally higher than in dollars and shocks to international food prices are now significant under the imperfect rational mechanism, for the whole period.

The positive effect observed of international food prices shocks, denominated in local currency, for the whole sample period when expectations are imperfect rational, is in line with the statements of the monetary authority of Colombia in 2007 and 2008 (see Footnote 6). It seems that pressures to international food prices were effectively transmitted to inflation. On this, two points are in order. First, when the sample period is split these effects are no longer observable casting some doubt on the interpretation of the monetary authority about the inflation outcomes in those years.[16] Second, those results would also cast some doubt on the assertion that in Colombia, during 2007 and 2008, shocks to prices of food and oil were weakened because of the appreciation of local currency (see Uribe, 2010). This is because when we use commodity prices shocks denominated in (the appreciated) local currency, food becomes significant and the coefficients of oil crude and energy become higher in absolute value. According to Table 3, this is the case with crude oil and energy under forward BEI expectations. Thus far, commodity prices shocks, mainly those to oil and energy, do affect the inflation process in Colombia but –given the size of the coefficients– they do it in a moderate way.

The evidence corresponding to the cases of crude oil and energy prices is not only consistent with the recent findings in the literature of a decrease in the contribution of oil prices to headline inflation (see for example De Gregorio et al., 2007), but is also related to the fact that long-term fluctuations in energy and crude oil prices are to a great extent already incorporated on inflation expectations.[17]

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[16] A possible interpretation is that movements in commodity prices were transmitted to inflation via expectations.

[17] As shown in Table 1, second round effects coming from annual

Moreover, to the extent that inflation expectations may contain a broad range of information about inflation coming from different sources,[18] it makes sense to find out that this is the main and more robust variable accounting for the deviations of the rate of inflation from target, as supported by estimation results. Accordingly, the message so far is that central bank's accomplishment of its price stability mandate by means of anchoring inflation expectations to the target is not only crucial, but is probably the essential feature and task under an optimal monetary policy regime.

The theoretical specification of expressions 6 and 8 suggests three results. First, the coefficient associated to expectation deviations should be bounded by 0 and 1. Second, this coefficient should be equal to the coefficient of structural supply shock (cost-push shock). Finally, the coefficient of expectation deviations under imperfect rational mechanism in specification 8 should be higher than the corresponding to specification 6. Unfortunately, only the third prediction holds; this either weakens the validity of the model or casts some doubt on the construction of some variables we have used.[19]

The model was also estimated by using the underlying inflation instead of headline inflation.[20] The former was obtained by using annual variations of CPI without food and regulated prices.[21] As before, the results presented in Tables 4 and 5

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variation in international food prices might be at play via inflation expectations.

[18] Some results, not included in the text, show that the permanent component of commodity prices variations is statistically significant explaining inflation expectations.

[19] Another version of the model was estimated by using inflation obtained from annualized monthly variation of CPI, the corresponding set of structural shocks and two versions of the inflation target: Monthly and annual. However the results are almost the same.

[20] This suggestion is due to an anonymous referee that we appreciate. Core inflation is measure as total inflation excluding food and administrated goods.

[21] Also annualized monthly variations of CPI without foods and

Table 3

ESTIMATES USING HEADLINE INFLATION, MONTHLY TARGET AND ONE-YEAR AHEAD EXPECTATIONS
(commodity prices denominated in local currency)

Expectation mechanism	*Constant*	*Expectation deviation*	*Demand-shock*	*Cost-push shock*	*Crude oil*	*Energy*	*Food*	*Number of observations*	*F*	*Adjusted R^2*
				Inflation targeting regime (January 2000 to December 2013)						
BEI	0.002	1.061[a]	0.001[b]	−0.002	−0.009		0.008	166	21.05	0.619
	0.002	1.073[a]	0.001[a]	−0.002		−0.008	0.006	166	18.54	0.609
Forward BEI	−0.001	1.057[a]	0.001	−0.001	−0.016[b]		0.010	166	11.24	0.593
	−0.001	1.079[a]	0.001	−0.001		−0.017[b]	0.009	166	10.41	0.579
Imperfect rational	−0.002	1.305[a]	0.000	−0.000	−0.005		0.022[b]	168	52.93	0.800
	−0.002[c]	1.321[a]	0.000	−0.000		−0.003	0.020[b]	168	52.93	0.796
				January 2000 to December 2006						
BEI	−0.001	0.913[a]	0.001[a]	−0.000	−0.006[c]		0.006	82	14.94	0.543
	−0.001	0.916[a]	0.001[a]	−0.001		−0.006	0.006	82	15.65	0.538

Forward BEI	−0.002[a]	0.799[a]	0.000	0.000	−0.007[b]		0.001	82	41.57	0.785
	−0.002[a]	0.804[a]	−0.000	−0.000		−0.008[b]	−0.000	82	50.33	0.785
Imperfect rational	−0.002[c]	1.107[a]	0.000	−0.000	−0.006		0.017	84	13.94	0.791
	−0.002[c]	1.118[a]	0.000	−0.000		−0.006	0.016	84	13.40	0.786
					January 2007 to December 2013					
BEI	0.004[b]	1.146[a]	0.001	−0.002	−0.012		0.011	84	23.65	0.684
	0.004[b]	1.168[a]	0.001	−0.002		−0.010	0.008	84	21.67	0.671
Forward BEI	−0.001	1.455[a]	0.001	−0.002	−0.027[a]		0.019	84	12.39	0.608
	−0.001	1.540[a]	0.001	−0.002		−0.030[a]	0.019	84	10.06	0591
Imperfect rational	−0.002	1.489[a]	0.001	0.001	−0.005		0.026	84	67.20	0.818
	−0.002	1.504[a]	0.000	0.001		−0.003	0.024	84	59.39	0.815

[a] represents significance at 1%, [b] at 5% and [c] at 10% based on Newey-West standard errors.
Source: authors' calculations.

include the monthly inflation target to compute the deviations of underlying inflation and inflation expectations; below we will use the annual target along the year.

In the version of the model in which commodity prices are denominated in dollars (see Table 4), demand and supply shocks have a more prominent role than in the case of headline inflation, mainly during the second subperiod; in contrast, commodity prices shocks do not have any significant effect on underlying inflation. Table 5 shows the results in which commodity prices are denominated in local currency. In this case, shocks to crude oil and energy prices are significant only in the second part of the sample under forward BEI expectations.

Models of Tables 4 and 5 have two important characteristics. First, the coefficients of expectations deviation are between 0 and 1 for the whole sample period and for period 2007-2013. Second, the coefficient of imperfect rational expectations is higher than the corresponding to BEI and forward BEI mechanisms. Thus, these data does not reject the model in these respects; however, the restriction that the coefficients of supply shocks and deviation expectations are equal is rejected.

These results may suggest that, under this optimal monetary policy framework –everything else equal–, inflation deviations are explained to a great extent by deviations of inflation expectations from target.[22] Therefore, as long as monetary authority reacts timely and accurately, such deviations should tend to decline, driving both underlying inflation and inflation expectations to the target. Moreover, an optimal monetary policy regime effectively conducts to a lesser exposure of inflation to commodity prices cyclical fluctuations.

Another version of the model corresponding to expressions 6 and 8 was obtained by using annual inflation target instead

regulated prices were used, but the results are, in general, the same.

[22] Recall we were expecting that cost-push shocks were also significant and that its coefficient be equal to the coefficient of deviations expectations. However, this restriction is not validated.

of the monthly one that we had been using so far.[23] The results in Table 6 suggest that the inflation process in Colombia is mainly driven by expectations. Structural and commodity prices shocks are only marginally important.

One remaining question is why the coefficients associated to inflation expectations in the cases of headline inflation are greater than in the cases of core inflation. Our intuition is that some of the permanent components of shocks are allowed to pass-through to transitory components of headline inflation but are not allowed to pass-through underlying inflation.

Estimation results for commodity prices denominated in local currency and with inflation expectations two years ahead are presented in Appendix 2. Results show an important decrease on the coefficients associated with the deviation of expectations from target for BEI expectations, indicating that, as time-horizon increases, inflation expectations converge to target.

[23] This specification was also recommended by an anonymous referee.

Table 4

ESTIMATES USING CORE INFLATION, MONTHLY TARGET AND ONE YEAR AHEAD EXPECTATIONS
(commodity prices denominated in dollars)

Expectation mechanism	Constant	Expectation deviation	Demand-shock	Cost-push shock	Crude oil	Energy	Food	Number of observations	F	Adjusted R^2
	Inflation targeting regime (January 2000 to December 2013)									
BEI	−0.011[a]	0.503[a]	0.001[a]	−0.001	−0.002		−0.009	166	4.75	0.295
	−0.011[a]	0.505[a]	0.001[a]	−0.001		−0.001	−0.010	166	4.80	0.294
Forward BEI	−0.012[a]	0.626[a]	0.001	−0.000	−0.005		−0.010	166	5.09	0.451
	−0.012[a]	0.637[a]	0.001	−0.000		−0.006	−0.009	166	4.87	0.455
Imperfect rational	−0.013[a]	0.771[a]	0.000	0.000	−0.003		0.006	168	11.30	0.553
	−0.013[a]	0.775[a]	0.000	0.000		−0.003	0.006	168	11.45	0.553
	January 2000 to December 2006									
BEI	−0.015[a]	1.006[a]	0.001[c]	0.000	−0.002		−0.017	82	25.45	0.652
	−0.015[a]	1.007[a]	0.001[c]	0.000		−0.001	−0.018	82	25.01	0.651
Forward BEI	−0.002[a]	0.748[a]	0.000	−0.000	−0.004		−0.014	82	18.93	0.685
	−0.002[a]	0.751[a]	0.000	−0.000		−0.004	−0.014	82	19.08	0.684

Imperfect rational	−0.016[a]	1.078[a]	−0.001	0.001	0.001		0.000	84	19.05	0.747
	−0.016[a]	1.094[a]	−0.001	0.001		0.004	−0.002	84	17.80	0.750
				January 2007 to December 2013						
BEI	−0.007[a]	0.272[a]	0.001[a]	−0.002[a]	−0.003		0.000	84	4.53	0.418
	−0.007[a]	0.280[a]	0.001[a]	−0.002[a]		−0.003	−0.001	84	5.05	0.415
Forward BEI	−0.008[a]	0.272[a]	0.002[a]	−0.002[a]	−0.006		0.000	84	6.69	0.271
	−0.008[a]	0.286[a]	0.002[a]	−0.002[a]		−0.007	−0.001	84	6.98	0.266
Imperfect rational	−0.008[b]	0.339[a]	0.001[b]	−0.001	−0.004		0.008	84	4.24	0.458
	−0.008[b]	0.349[a]	0.001[b]	−0.001		−0.004	0.007	84	4.26	0.454

Note: [a] represents significance at 1%, [b] at 5% and [c] at 10% based on Newey-West standard errors.
Source: authors' calculations.

Table 5

ESTIMATES USING CORE INFLATION, MONTHLY TARGET AND ONE YEAR AHEAD EXPECTATIONS
(commodity prices denominated in local currency)

Expectation mechanism	*Constant*	*Expectation deviation*	*Demand-shock*	*Cost-push shock*	*Crude oil*	*Energy*	*Food*	*Number of observations*	*F*	*Adjusted R^2*
Inflation targeting regime (January 2000 to December 2013)										
BEI	−0.011[a]	0.486[a]	0.001[b]	−0.001	−0.004		0.000	166	4.57	0.280
	−0.011[a]	0.491[a]	0.001[b]	−0.001		−0.003	0.000	166	4.60	0.278
Forward BEI	−0.012[a]	0.622[a]	0.001[b]	−0.001	−0.008		−0.003	166	7.33	0.434
	−0.012[a]	0.637[a]	0.001[b]	−0.001		−0.009	−0.002	166	6.57	0.436
Imperfect rational	−0.013[a]	0.768[a]	−0.000	0.000	−0.003		0.008	168	10.16	0.555
	−0.013[a]	0.773[a]	−0.000	0.000		−0.003	0.008	168	10.28	0.555
January 2000 to December 2006										
BEI	−0.015[a]	1.008[a]	0.001	0.000	−0.002		−0.001	82	21.74	0.635
	−0.015[a]	1.009[a]	0.001	0.000		−0.000	−0.003	82	22.83	0.634
Forward BEI	−0.016[a]	−0.778[a]	0.000	0.001	−0.003		−0.007	82	17.76	0.682
	−0.016[a]	−0.780[a]	−0.000	0.001		−0.003	−0.007	82	18.11	0.681

Imperfect rational	−0.016[a]	1.078[a]	−0.001[c]	0.001	0.001		0.010	84	15.42	0.759
	−0.016[a]	1.092[a]	−0.001[c]	0.001		0.003	0.007	84	14.82	0.761
					January 2007 to December 2013					
BEI	−0.007[a]	0.269[a]	0.001[a]	−0.002[a]	−0.005		0.002	84	4.77	0.416
	−0.007[a]	0.277[a]	0.002[a]	−0.002[a]		−0.005	0.001	84	5.33	0.410
Forward BEI	−0.008[a]	0.253[a]	0.002[a]	−0.002[a]	−0.007[b]		0.005	84	8.33	0.269
	−0.008[a]	0.277[a]	0.002[a]	−0.002[a]		−0.008[c]	0.005	84	7.96	0.259
Imperfect rational	−0.008[a]	0.338[a]	0.001[b]	−0.001	−0.003		0.006	84	4.36	0.453
	−0.008[a]	0.345[a]	0.001[b]	−0.001		−0.003	0.006	84	4.42	0.451

Note: [a] represents significance at 1%, [b] at 5% and [c] at 10% based on Newey-West standard errors.
Source: authors' calculations.

Table 6

ESTIMATES USING CORE INFLATION, ANNUAL TARGET AND ONE YEAR AHEAD EXPECTATIONS
(commodity prices denominated in dollars)

Expectation mechanism	*Constant*	*Expectation deviation*	*Demand-shock*	*Cost-push shock*	*Crude oil*	*Energy*	*Food*	*Number of observations*	*F*	*Adjusted R^2*
	Inflation targeting regime (January 2000 to December 2013)									
BEI	−0.009[a]	0.302[a]	0.001[c]	−0.000	0.001		−0.007	166	6.14	0.157
	−0.009[a]	0.299[a]	0.001[c]	−0.000		0.002	−0.008	166	6.38	0.159
Forward BEI	−0.011[a]	0.435[a]	0.001	−0.000	−0.001		−0.008	166	7.36	0.323
	−0.011[a]	0.438[a]	0.001	−0.000		−0.002	−0.008	166	6.98	0.323
Imperfect rational	−0.011[a]	0.476[a]	0.000	0.000	−0.000		0.003	168	8.21	0.287
	−0.011[a]	0.475[a]	0.000	0.000		0.000	0.002	168	7.46	0.287
	January 2000 to December 2006									
BEI	−0.014[a]	0.761[a]	0.001	0.000	−0.001		−0.014	82	15.74	0.428
	−0.014[a]	0.767[a]	0.000	0.000		−0.000	−0.015	82	15.23	0.427
Forward BEI	−0.014[a]	0.594[a]	−0.000	0.001[b]	−0.002		−0.011	82	12.01	0.594
	−0.014[a]	0.595[a]	−0.000	0.001[b]		−0.002	−0.011	82	12.05	0.593

Imperfect rational	−0.015[a]	0.754[a]	−0.000	0.001[c]	0.000		−0.002	84	8.81	0.437
	−0.015[a]	0.756[a]	−0.000	0.001[b]		0.002	−0.003	84	8.42	0.439
	January 2007 to December 2013									
BEI	−0.006[a]	0.246[a]	0.001[a]	−0.001[b]	0.003		−0.006	84	5.34	0.379
	−0.006[a]	0.240[a]	0.001[a]	−0.001[b]		0.003	−0.007	84	5.80	0.382
Forward BEI	−0.007[a]	0.240[b]	0.001[a]	−0.001[b]	−0.000		−0.003	84	5.50	0.244
	−0.007[a]	0.238[b]	0.001[a]	−0.001[b]		0.000	−0.004	84	5.57	0.244
Imperfect rational	−0.007[a]	0.291[b]	0.001[b]	−0.001	0.001		0.002	84	3.62	0.366
	−0.007[a]	0.287[b]	0.001[b]	−0.001		0.002	0.001	84	3.57	0.368

Note: [a] represents significance at 1%, [b] at 5% and [c] at 10% based on Newey-West standard errors. Source: authors' calculations.

Table 7

ESTIMATES USING CORE INFLATION, ANNUAL TARGET AND ONE YEAR AHEAD EXPECTATIONS
(commodity prices denominated in local currency)

Expectation mechanism	*Constant*	*Expectation deviation*	*Demand-shock*	*Cost-push shock*	*Crude oil*	*Energy*	*Food*	*Number of observations*	*F*	*Adjusted R^2*
	Inflation targeting regime (January 2000 to December 2013)									
BEI	−0.009[a]	0.287[a]	0.001[c]	−0.001	0.001		−0.001	166	5.68	0.151
	−0.009[a]	0.284[a]	0.001[c]	−0.001		0.001	−0.002	166	5.88	0.152
Forward BEI	−0.011[a]	0.424[a]	0.001[b]	−0.000	−0.003		−0.003	166	8.68	0.309
	−0.011[a]	0.428[a]	0.001[b]	−0.000		−0.003	−0.003	166	8.21	0.308
Imperfect rational	−0.011[a]	0.474[a]	0.000	0.000	−0.000		0.005	168	8.83	0.293
	−0.011[a]	0.473[a]	0.000	0.000		0.000	0.004	168	7.84	0.293
	January 2000 to December 2006									
BEI	−0.014[a]	0.736[a]	0.000	0.000	−0.001		0.002	82	10.39	0.408
	−0.014[a]	0.733[a]	0.000	0.000		0.001	0.001	82	10.92	0.407
Forward BEI	−0.014[a]	0.604[a]	−0.000	0.001[b]	−0.001		−0.003	82	10.41	0.583
	−0.014[a]	0.604[a]	−0.000	0.001[b]		−0.001	−0.003	82	10.53	0.582

Imperfect rational	-0.015^{a}	0.745^{a}	-0.001^{c}	0.001^{b}	-0.000		0.012^{c}	84	8.14	0.469
	-0.015^{a}	0.747^{a}	-0.001^{b}	-0.000		0.001	0.010	84	8.30	0.470
January 2007 to December 2013										
BEI	-0.006^{a}	0.252^{a}	0.001^{a}	-0.001^{b}	0.002		-0.009	84	4.94	0.402
	-0.006^{a}	0.248^{a}	0.001^{a}	-0.001^{b}		0.002	-0.010	84	5.29	0.404
Forward BEI	-0.007^{a}	0.248^{b}	0.002^{a}	-0.001^{c}	-0.001		-0.006	84	6.15	0.264
	-0.007^{a}	0.248^{b}	0.002^{a}	-0.002^{b}		-0.001	-0.006	84	6.10	0.263
Imperfect rational	-0.007^{a}	0.293^{b}	0.001^{b}	-0.001	0.003		-0.004	84	3.99	0.360
	-0.007^{a}	0.286^{b}	0.001^{b}	-0.001		0.003	-0.005	84	3.86	0.363

Note: [a] represents significance at 1%, [b] at 5% and [c] at 10% based on Newey-West standard errors. Source: authors' calculations.

6. MAIN FINDINGS AND CONCLUSIONS

This paper analyzes the effects of the recent movements of commodity prices, in the domestic inflation of Colombia derived from an optimal monetary policy framework. The empirical specification is derived from a simple, yet intuitive, text-book model of a small open economy that follows an optimal monetary policy rule, close to those used by inflation targeting countries. The model is highly demanding in terms of variables; thus, we use different definitions of inflation, expectations, inflation target, and shocks. The estimations were performed for the whole period of the inflation targeting regime: January 2000 to December 2013 and two subsamples: from January 2000 up to December 2006 which corresponds to the period before the commodity boom and from January 2007 to December 2013, subsample period after the boom.

The whole picture shows an inflation process governed by expectations in Colombia. Our findings also suggest that crude oil, energy, and food price shocks have a small influence on inflation irrespective that commodity prices shocks are rated in dollars or Colombian pesos. This would support the recent findings in the literature of a substantial decrease in the pass-through of oil prices to headline inflation. Our interpretation is that much of permanent movements in commodity prices are passed through inflation via expectations if agents judge that such movements are persistent. Finally, the contribution of demand and cost-push shocks in inflation is also small.

The model and estimations suggest that under an optimal monetary policy framework –everything else equal– deviations of inflation from target will respond to deviations of inflation expectations from target. Therefore, as long as monetary authority reacts timely and accurately, such deviations will tend to decline, leading both inflation and inflation expectations to the target. In our view, monetary authority has faced rightly the shocks to commodity world prices. When the target was missed, in 2007 and 2008, the reasons should be others.

APPENDICES

Appendix 1

Identification and Estimation of Supply and Demand Shocks

The explanation of Cover, Enders and Hueng (2006), CEH, method can be undertaken by means of the standard Blanchard-Quah (1989), BQ, identification scheme; that is what we do first. Then we introduce the modification of CEH. Given the aggregate demand- aggregate supply model, AD-AS

1 $$y_t^s = y_{t/t-1} + \alpha\left(p_t - p_{t/t-1}\right) + \varepsilon_t \ ,$$

2 $$\left(y_t + p_t\right)^d = \left(y_{t/t-1} + p_{t/t-1}\right)^d + \eta_t \ ,$$

3 $$y_t^s = y_t^d \ .$$

Which can be expressed in matrix form

4 $$\begin{bmatrix} 1 & -\alpha \\ 1 & 1 \end{bmatrix}\begin{bmatrix} y_t \\ p_t \end{bmatrix} = \begin{bmatrix} 1 & -\alpha \\ 1 & 1 \end{bmatrix}\begin{bmatrix} y_{t/t-1} \\ p_{t/t-1} \end{bmatrix} + \begin{bmatrix} \varepsilon_t \\ \eta_t \end{bmatrix},$$

5 $$\begin{bmatrix} y_t \\ p_t \end{bmatrix} = \begin{bmatrix} y_{t/t-1} \\ p_{t/t-1} \end{bmatrix} + \begin{bmatrix} 1 & -\alpha \\ 1 & 1 \end{bmatrix}^{-1}\begin{bmatrix} \varepsilon_t \\ \eta_t \end{bmatrix}.$$

with variance-covariance of the vector of structural shocks

$$\begin{bmatrix} \sigma_\varepsilon^2 & \sigma_{\varepsilon,\eta} \\ \sigma_{\varepsilon,\eta} & \sigma_\eta^2 \end{bmatrix}.$$

Assuming that the expectation of each variable is a linear combination of its own lags, then Equation 5 reduces to a VAR model:

6 $$\begin{bmatrix} y_t \\ p_t \end{bmatrix} = \begin{bmatrix} A_{11}(L) & A_{12}(L) \\ A_{21}(L) & A_{22}(L) \end{bmatrix}\begin{bmatrix} y_t \\ p_t \end{bmatrix} + \begin{bmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{bmatrix}\begin{bmatrix} \varepsilon_t \\ \eta_t \end{bmatrix}.$$

The long run- response of the shocks is given by:

7 $$\Psi_\infty = \left[I - A(1)\right]^{-1}\Theta \ ,$$

where $\Theta = \begin{bmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{bmatrix}$, $A(1) = \begin{bmatrix} A_{11}(1) & A_{12}(1) \\ A_{21}(1) & A_{22}(1) \end{bmatrix}$.

According to BQ, imposing the assumption that the AD shock, η_t, has no long-run effect on output, and assuming that $\sigma_\varepsilon^2 = 1,\ \sigma_\eta^2 = 1,\ \sigma_{\varepsilon,\eta} = 0$ imply that $c_{12}\left[1 - A_{22}(1)\right] + c_{22}A_{12}(1) = 0$.

With this restriction, the signs of c_{ij} are not identified, and there are four possible solutions for those values, choosing the one that implies a positive long-run effect of demand shock on price and a positive long-run effect of supply shock on output.

On the other hand, CHE using the values of c_{ij} derived from Equation 5, and without imposing any restriction over the variance-covariance matrix of structural shocks, only one solution is gotten by assuming the neutrality condition in 7:

$$\begin{bmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{bmatrix} = \begin{bmatrix} \frac{1}{1+\alpha} & \frac{\alpha}{1+\alpha} \\ \frac{-1}{1+\alpha} & \frac{1}{1+\alpha} \end{bmatrix},$$

$$\alpha = A_{12}(1) / \left[1 - A_{22}(1)\right],$$

and the variance-covariance matrix of structural shocks can be estimated from the variance-covariance matrix of the VAR innovations, after knowing the value of α.

$$\begin{bmatrix} \mathrm{var}(e_{1t}) & \mathrm{covar}(e_{1t}, e_{2t}) \\ \mathrm{covar}(e_{1t}, e_{2t}) & \mathrm{var}(e_{2t}) \end{bmatrix} = \begin{bmatrix} \frac{1}{1+\alpha} & \frac{\alpha}{1+\alpha} \\ \frac{-1}{1+\alpha} & \frac{1}{1+\alpha} \end{bmatrix} \begin{bmatrix} \sigma_\varepsilon^2 & \sigma_{\varepsilon,\eta} \\ \sigma_{\varepsilon,\eta} & \sigma_\eta^2 \end{bmatrix} \begin{bmatrix} \frac{1}{1+\alpha} & \frac{-1}{1+\alpha} \\ \frac{\alpha}{1+\alpha} & \frac{1}{1+\alpha} \end{bmatrix}$$

In order to identify orthogonal structural shocks, the two ordering in the Cholesky decomposition are used. The first order assumes there is causality from supply shock, ε_t, to the demand shock, η_t, which may be imposed by assuming that $\eta_t = \rho\varepsilon_t + \nu_t$, where ν_t is a pure AD shock and ρ is the unexpected AD change due to an AS shock. On the other hand, the second order assumes there is causality from the demand shock to the supply shock. In this case, define $\varepsilon_t = \gamma\eta_t + \upsilon_t$, where υ_t is a pure AS shock and γ is the unexpected AS change induced by an AD shock.

Model in Equation 6 remains the same with any of the orderings, by assuming

$$\begin{bmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{bmatrix} = \begin{bmatrix} \frac{1+\alpha\rho}{1+\alpha}\sigma_{\varepsilon} & \frac{\alpha}{1+\alpha}\sigma_{v} \\ \frac{-(1-\rho)}{1+\alpha}\sigma_{\varepsilon} & \frac{1}{1+\alpha}\sigma_{v} \end{bmatrix}$$

or

$$\begin{bmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{bmatrix} = \begin{bmatrix} \frac{1}{1+\alpha}\sigma_{\delta} & \frac{\alpha+\gamma}{1+\alpha}\sigma_{\upsilon} \\ \frac{-1}{1+\alpha}\sigma_{\delta} & \frac{1-\gamma}{1+\alpha}\sigma_{\upsilon} \end{bmatrix}$$

then,

$$\begin{bmatrix} \operatorname{var}(e_{1t}) & \operatorname{covar}(e_{1t}, e_{2t}) \\ \operatorname{covar}(e_{1t}, e_{2t}) & \operatorname{var}(e_{2t}) \end{bmatrix} = \begin{bmatrix} c_{11} & c_{12} \\ c_{21} & c_{22} \end{bmatrix} \begin{bmatrix} 1 & 0 \\ 0 & 1 \end{bmatrix} \begin{bmatrix} c_{11} & c_{21} \\ c_{12} & c_{22} \end{bmatrix}$$

Appendix 2

Results with Expectations Two Years Ahead

Table A1

ESTIMATES USING CORE INFLATION, ANNUAL TARGET AND TWO YEARS AHEAD EXPECTATIONS
(commodity prices denominated in local currency)

Expectation mechanism	*Constant*	*Expectation deviation*	*Demand-shock*	*Cost-push shock*	*Crude oil*	*Energy*	*Food*	*Number of observations*	*F*	*Adjusted R^2*
	Inflation-targeting-regime (January 2000 to December 2013)									
BEI	−0.008[a]	0.311[a]	0.001[a]	−0.001[b]	−0.004		0.000	132	21.35	0.340
	−0.008[a]	0.317[a]	0.001[a]	−0.001[b]		−0.004	0.001	132	19.04	0.337
Imperfect rational	−0.009[a]	−0.065	0.000	−0.000	0.001		0.007	168	1.816	0.000
	−0.008[a]	−0.073	0.000	−0.000		0.002	0.005	168	1.825	0.004
	January 2000 to December 2006									
BEI	−0.011[a]	0.495[a]	0.001	−0.000	−0.009[a]		0.007	48	36.43	0.535
	−0.011[a]	0.500[a]	0.001	−0.000		−0.009	0.008	48	34.99	0.524
Imperfect rational	−0.012[a]	−0.008	−0.000	0.001[a]	−0.000		0.014	84	2.251	0.020
	−0.012[a]	−0.015	−0.000	0.001[a]		0.000	0.013	84	2.341	0.020

	January 2007 to December 2013									
BEI	−0.007[a]	0.288[a]	0.001[a]	−0.001[a]	−0.000		−0.008	84	6.101	0.385
	−0.007[a]	0.287[a]	0.001[a]	−0.001[a]		0.000	−0.009	84	6.392	0.385
Imperfect rational	−0.007[a]	0.104	0.002[a]	−0.002[a]	0.003		−0.007	84	2.825	0.101
	−0.007[a]	0.094	0.002[a]	−0.002[a]		0.005	−0.008	84	2.879	0.116

Note: [a] represents significance at 1%, [b] at 5% and [c] at 10% based on Newey-West standard errors.
Source: authors' calculations.

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Countercyclical Capital Buffer: The Case of Uruguay

Abstract

We study the countercyclical capital buffer introduced by Basel III and its complementarities with other regulation, particularly dynamic provisioning. We simulate different activation, adjust and deactivation rules for the buffer using historical data for Uruguay. The design and introduction of a countercyclical capital buffer following the principles in Basel III should complement current regulation and serve as an extra tool to mitigate systemic risk in the Uruguayan banking sector.

Keywords: Basel III, countercyclical requirement, macroprudential tools, Uruguay.
JEL classification: G18, G21.

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1. INTRODUCTION

Empirical evidence demonstrates the significant importance of excessive credit growth when determining the likelihood and severity of systemic financial crises.[1] The dynamic formation of systemic risk means that banking sector losses can be extremely large when a period of excess credit growth is followed by a recession. Loans granted during a period of excessive growth tend to be of lower quality than those granted during more stable periods, meaning the losses they generate can destabilize the banking sector and spark a vicious circle whereby financial system difficulties can contribute to a downturn in the real economy that then feeds back on to the banking sector.

These interactions between financial and real cycles have led to an important debate, among academics as well as financial regulators, regarding the macroprudential measures that should be adopted. In particular, there is apparent consensus on the importance of the banking sector strengthening its capital defenses during periods where risks of system-wide stress are growing markedly. Thus, the Basel Committee on Banking Supervision (2010a), in its regulatory framework for more resilient banks and banking systems (Basel III), has proposed introducing countercyclical capital buffers aimed at helping to protect banks from the effects of financial cycles. In order to ensure capital requirements, take into account the macrofinancial environment in which banks operate, the proposal of the Basel Committee on Banking Supervision aims to accumulate capital reserves when imbalances start to build up and vulnerabilities increase in order to allow them to be used in times of crisis or financial instability.

The agreement proposed by the Basel Committee on Banking Supervision (2010a) (also known as Basel III) refers to a group of reform proposals motivated by the failures identified

[1] See for example Davis and Karim (2008), Drehmann and Juselius (2013), Drehmann and Tsatsoronis (2014) and the references provided therein.

during the recent international financial crisis. The agreement is based on a review of Basel II aimed at making further progress on strengthening the banking sector. The Basel Committee has reinforced the capital requirement framework, increasing both the quality and level of regulatory capital requirements. In particular, it introduces the requirement for accumulating a countercyclical capital buffer to be used by the banking sector during the trough of the financial cycle. The idea of the buffer is for banks to accumulate extra capital (up to 2.5% of risk weighted assets) during the high point of the financial cycle, for instance when bank credit is growing very rapidly as compared to the level of economic activity, and use it as soon as risks start to materialize.

This paper analyzes the main characteristics of the countercyclical capital buffer proposed by Basel III in light of related economic literature and recent international experience. It also makes a conceptual study of introducing the countercyclical requirement in Uruguayan regulation and its complementarities with other regulatory tools, particularly dynamic provisioning. Finally, we simulate different activation, adjust and deactivation rules for the buffer using historical data for Uruguay. It can be concluded from the analysis that the design and introduction of a countercyclical capital buffer following the principles in Basel III would complement current regulation and serve as an extra tool to mitigate systemic risk in the Uruguayan banking sector.

The remainder of this paper is organized as follows. Section 2 outlines the main characteristics of the countercyclical capital buffer. Section 3 briefly reviews the literature on countercyclical capital buffers. Section 4 analyzes coexistence of the new capital buffer with the current dynamic provisioning in Uruguayan banking regulations. Section 5 describes the results of the methodology proposed by the Basel Committee on Banking Supervision (2010a) with respect to different activation, adjust and deactivation rules for the buffer using historical data for Uruguay. Lastly, section 6 gives some final remarks.

2. THE COUNTERCYCLICAL CAPITAL BUFFER

The main objective of the countercyclical capital buffer set out by the Basel Committee on Banking Supervision (2010a) is to protect banks from financial cycle effects by enhancing their capacity to absorb losses (by accumulating capital) in periods when systemic vulnerabilities are growing (during periods of excessive credit growth for instance). The capital buffer built up in such periods can be deployed in the low point of the financial cycle to absorb losses and thereby help banks overcome periods of stress. A larger capital buffer in the high point of the financial cycle can also help to reduce excess supply of credit and risk-taking.

The proposed countercyclical capital buffer, which can vary between zero and 2.5% of risk weighted assets, is a system-wide prudential tool (macroprudential). Once activated, its scope of application is the whole banking sector, regardless of the potential contribution of individual banks to the excess supply of loans. The buffer is in addition to any other capital requirements that might exist, but unlike them, its activation, adjustment and deactivation is at the discretion of the bank regulator. The Basel Committee on Banking Supervision (2010a) suggests using the long-term trend of the aggregate private sector credit-to-GDP gap as a reference to inform regulatory authorities on the phase of the financial cycle and guide activation of the countercyclical buffer.[2]

Conceptually, the countercyclical capital buffer is a complement to other existing regulatory tools. For example, consider a regulatory tool classification of two dimension: *i)* according to which their aim and basis for application is institution by institution (institution specific prudential) or systemic (system-wide prudential), *ii)* where their objective is the static or

[2] Drehmann and Tsatsaronis (2014) suggest that this indicator fulfills an important role in signaling when it is necessary to activate the countercyclical buffer. However, Repullo and Saurina (2011) argue that using it would tend to exacerbate fluctuations rather than reduce them.

dynamic dimension of financial risks. In these two dimensions, the countercyclical capital buffer is a prudential tool that addresses the rapid build-up of system-wide risks (see table 1).

Table 1

COUNTERCYCLICAL CAPITAL BUFFER AND OTHER PRUDENTIAL REGULATION

Prudential dimension	*Static*	*Dynamic*
Institution specific	Minimum capital and provisions requirement	Dynamic provisioning
System-wide	Systemic risk capital requirement	Countercyclical capital buffer

The countercyclical capital buffer therefore complements other prudential measures for the system that address the static dimension of financial risks. For example, in some jurisdictions capital requirements for systemic risk take into account the size of banking institutions, their interconnections and importance in the payments system as relevant variables for demanding extra capital requirements from banks considered systemically important. The countercyclical capital buffer is also complemented by institution specific prudential regulation tools. In particular, it complements static capital requirements (be the minimum requirement or the conservation buffer) and credit risk provisions because they explicitly consider the dynamic dimension of credit risk. Moreover, it complements dynamic provisioning, given that the latter is based on the situation of each individual bank, while the countercyclical capital buffer considers the aggregate or systemic situation.

In some jurisdictions, practical application of the countercyclical capital buffer has deviated from the recommendations of the Basel Committee on Banking Supervision (2010a). In the case of Switzerland, for instance, the countercyclical capital buffer is activated at the discretion of the authorities using as a reference a wide range of both aggregate and sectoral indicators and variables, as well as quantities such as prices. Meanwhile, both England and Switzerland base their countercyclical capital buffers on the behavior of certain sectors, particularly the mortgage sector. These types of measures therefore pursue the objective of controlling the rapid generation of systemic risk in sectors identified as particularly vulnerable, requiring additional capital buffers in banks that are most exposed to such sectors, while less exposed banks do not necessarily have to build such buffers.

The application of capital requirements based on the behavior of specific credit sectors and, therefore on the most exposed banks, is complementary to the application of capital requirements on an aggregate basis for all banks. The first approach addresses the need to recognize expected losses from the performance of individual banks in market segments where financial imbalances are being generated as a result of, for instance, substantial sectoral credit growth. In some jurisdictions, such as Spain and Uruguay, dynamic provisioning is employed in order to recognize such risks in advance and provide incentives for banks to reduce their exposure to them.[3] Thus, dynamic provisions, in their most common form, are an institution specific prudential tool. Meanwhile, a countercyclical capital buffer applied in aggregate form is a system-wide prudential tool aimed at raising the banking system as a whole's resilience to periods of stress and remain stable so it can continue providing its services to the rest of the economy.

Finally, the recommendation of the Basel Committee on Central Banking (2010a) does not take into account a key characteristic of the Uruguay's banking system: dollarization. If there were sharp differences between currency specific credit

[3] The regulatory tools of England and Switzerland do not include the possibility of implementing dynamic provisioning.

cycles, there would be a reason to consider data broken down by currency above aggregate data, given that the latter could be hiding important sources of systemic risk. Ultimately, the answer as to which series should be used as a reference is of empirical nature. Section 5 presents an analysis of credit series by currency and concludes that although their cycles have exhibited different behaviors during the recent history of Uruguay, the aggregate series appropriately captures the trend-cycle performance of the disaggregated series.

3. RELATED LITERATURE

Debate on whether the business cycle might be amplified as a consequence of the regulatory framework implemented by the Basel II agreements (see Basel Committee on Banking Supervision, 2005) began even before their approval. For instance, Kashyap and Stein (2004) argued that losses during a downturn erode bank capital, while risk sensitive capital requirements increase. If banks are not able to rapidly raise their capital they are forced to reduce their supply of loans, which contributes to a worsening of the downturn.

In light of the recent international financial crisis, reforms to the regulatory capital framework proposed by the Basel Committee on Banking Supervision (2010a) were aimed at raising both the level and quality of the regulatory capital base and, particularly, at reducing any kind of cyclical behavior in minimum capital requirements, as well as maintaining a capital buffer with the macroprudential goal of protecting the banking sector from the potentially negative impact of periods of excess credit growth.

This section provides a summary of the literature related to the financial cycle and the procyclical behavior of current capital requirements based on the Basel II agreement in order to demonstrate the potential impact of introducing countercyclical capital buffers under the Basel III framework.

Bergara and Licandro (2000) propose a microeconomic model for identifying what share of credit procyclicality responds to bank behavior and what share to the prudential regulatory framework. They conclude that credit is procyclical even if there is no prudential regulation or if the latter is loose, because bankers' myopia and risk aversion affects their return-risk perception. They therefore conclude that prudential regulations do not exacerbate the credit cycle but actually manage to smooth it.

Repullo and Suárez (2008) link bank capital requirements with the credit rationing of some firms through an bridged generations model that assumes the existence of relational banking (banks have private information about their borrowers) and the inability of some banks to access the equity market. They find that, under Basel II regulations, although banks hold larger capital buffers during expansions, the arrival of recessions is usually associated with significant credit rationing. They set forth that some adjustments in the confidence level of Basel II can substantially reduce the incidence of credit rationing throughout the business cycle without compromising solvency targets. In particular, they propose modifying the confidence levels in a way that keeps their long-term average at 99.9%, but lessens the target in situations where credit rationing turns out to be the highest.

Another alternative for correcting the procyclicality of capital requirements is that suggested by the Committee of European Bank Supervisors (CEBS, 2009), consisting of a mechanism that adjusts probabilities of default estimated by banks in order to incorporate recessionary conditions. In particular, they propose two alternatives: applying an adjustment based on the gap between current probabilities of default and those corresponding to a recession, and using confidence levels that automatically adjust as the result of changing cycles.

Repullo, Saurina and Trucharte (2010) study the most important alternatives to the credit gap indicator for mitigating the procyclical effects of Basel II requirements. The analysis is based on an estimation of the one-year ahead probabilities of default of Spanish firms during the period 1986-2007 using data from Spain's Credit Register. They therefore obtain a risk profile for each bank by calculating the corresponding Basel II capital requirements for each loan.

They compare different alternatives for adjusting capital requirements throughout the cycle, concluding that the best procedure is to use a multiplier of the economic cycle based on GDP growth. They analyze two alternatives proposed by Gordy and Howels (2006): smooth the inputs of the formula using through-the-cycle adjustment in the probabilities of default, and smooth the outputs of the formula by adjusting final capital requirements computed from probability of default estimates. The results show that the best procedure is to use a multiplier of the capital requirement. Such multiplier depends on the deviation of the growth rate of GDP growth with respect to its long-term trend.

Elekdag and Wu (2011) analyze the development of credit booms based on an event study with a panel of advanced and emerging countries covering the period 1960-2010. Among the main results of the paper stand out the association they find of credit *booms* with deteriorating bank and corporate balance sheets, as well as with symptoms of economic overheating. With respect to the referred indicator for correcting procyclicality, they suggest that the credit-to-GDP gap does not allow for contemplating the possibility that credit and output have different trends, and that a fall in GDP could give rise to decisions that might exacerbate the procyclicality instead of smoothing it.

Christensen, Meh and Moran (2011) compare the impact of bank leverage regulation with constant time-invariant requirements to that with requirements that change according to the cycle (countercyclical regulation). The outcomes suggest that the countercyclical buffer manages to keep the development of financial imbalances under control by inducing banks to alter the intensity with which they monitor their borrowers.

Gersbach and Rochet (2012) propose a formal rationale for imposing countercyclical capital ratios. They find that banks allocate too much borrowing capacity to good states and too little to bad states, creating excessive volatility in credit, GDP, asset prices and wages. Using a very simple model in which financial frictions generate excessive fluctuations in the volume of credit, they demonstrate that the latter can be smoothed by regulatory countercyclical capital ratios.

Dewatripont and Tirole (2012) also use a formal model to analyze banking regulation, understood as a combination of self-insurance mechanisms, capital buffers and provisions, in the presence of macroeconomic shocks. Their results show that the combination of mechanisms such as dynamic provisioning, countercyclical buffers, as well as other forms of capital insurance, such as contingent convertible bonds (CoCos), is optimal for neutralizing the adverse effects of macroeconomic shocks of both deterministic and random origin.

Buncic and Melecky (2013) propose a new approach to macroprudential stress testing of the banking system. Stress tests used up until now have been mainly based on financial simulations where no formal links to the macroeconomy are established. The methodology they propose incorporates explicit links between the financial system and the macroeconomy, allowing for contemplating the possible emergence of systemic risks deriving from changes in macroeconomic conditions, as well as idiosyncratic risks originating from the different risk profiles of individual banks. The results are robust when the methodology is applied to a set of Eastern European banks during the recent international financial crisis.

Repullo (2013) concludes that when models incorporate a social cost of bank failure the regulator sets higher capital requirements as compared to a situation without banking regulation. However, there is a trade-off: Banks are safer but aggregate investment is lower. The paper also analyzes the impact of a negative shock to the aggregate supply of bank capital (equivalent to a downturn of the economy). If capital requirements are kept unchanged, the reduction in the supply of capital implies a significant fall in bank lending and aggregate investment (although banks are safer). In sum, the paper compares the costs and benefits of adjusting capital requirements to changes in the business cycle, concluding that the regulator should not only focus on the credit rationing that could arise if capital requirements are not lowered during recessions, or on the greater likelihood of bank failures if they are.

Drehmann and Tsatsaronis (2014) respond to some of the criticism of the credit-to-GDP gap indicator. In particular, they offer counterarguments to the following observations: The credit gap indicator can lead to decisions that conflict with its objective, it is not the best early warning indicator for banking crises (especially for emerging economies), and it also has some measurement problems.

The first criticism argues that the relevant cycle for the instrument should be the financial cycle and not the business cycle. As mentioned above, Repullo and Saurina (2011) find a negative correlation between the credit gap and GDP growth, meaning a capital buffer determined according to such criteria could exacerbate the cycle it is attempting to smooth. However, Drehmann and Tsatsaronis find that, although it is negative, said correlation is very small and mainly determined by periods that are irrelevant in decisions for building a capital buffer.

To answer the second criticism, they use a panel of 26 countries over the period 1980 and 2012 to compare the performance of six indicators: The credit-to-GDP gap, GDP growth, residential property price growth, the debt service ratio and the non-core liability ratio. Among the variables considered, the credit-to-GDP gap ratio is statistically the best early warning indicator for two to five year forecast horizons.

As for the indicator's measurement problems, these relate to the now well-known limitations of estimating a trend with the Hodrick-Prescott (1981) filter: the most recent observations can considerably change the results. In this regard, simulated data estimations suggest using series with at least 10 years of available data to overcome this problem.

Wezel, Chan-Lau and Columbra (2012) make a brief comparison between countercyclical capital buffers and dynamic provisioning methods. They state that although dynamic provisioning considers fluctuations in the specific provisions of each loan, it does not take into account changes in the probabilities of default and losses, once default has taken place, used as an input in the capital requirement formulas of Basel II. They conclude that both tools can complement one another as long as policies for provisions focus on bolstering the banking sector against expected losses, while capital measures focus on unexpected losses. In particular, they argue that although dynamic provisioning directly protects bank results, they have little capacity to restrain excessive credit growth, suggesting they should therefore be accompanied by other macroprudential measures aimed at mitigating systemic risks.

Finally, it is important to mention that this paper is based on an initial analysis of the impact of the new countercyclical capital buffers in aggregate terms, but banks can decide to hold different amounts of capital according to their individual characteristics, such as their appetite for risk, their size or access to sources of funding other than agents' deposits. Recent literature includes a series of papers analyzing the cyclical behavior of bank capital taking into account the diversity that could exist among banking institutions.

Jokipii and Milne (2008) analyze the cyclical behavior of capital buffers that European banks decide to hold above Basel I capital requirements, as well as the possible changes in such behavior across different countries and types or sizes of institution. Using a panel for the period 1997-2004, they find that although banks hold capital amounts above the minimum requirement, said decision varies according to the type and size of bank. They conclude that capital buffers of large institutions, commercial banks and savings banks exhibit negative cyclical comovement, while those of cooperative banks and smaller banks behave procyclically.

Fonseca and González (2010) work with a panel data of banks from 70 countries for the period 1995-2002 in order to study the factors that influence the decisions to hold bank capital buffers. In particular, they analyze how different regulatory and institutional designs across countries can lead to differing behaviors in bank market power and market discipline, these being two factors that play an important role in bank decisions to hold capital buffers higher than the minimum requirement.

García-Suaza *et al.* (2012), meanwhile, study the cyclical behavior of capital in the Columbian banking sector using a panel of banks for the period 1996-2010. They conclude that although bank capital buffers vary throughout the cycle, this behavior differs according to the size of the institution. In particular, they confirm countercyclical behavior for large banks, but do not find evidence of the same behavior for small banks.

Finally, Carvallo *et al.* (2015) study the cyclical behavior of capital buffers based on a panel of the banking sectors of 13 Latin American and Caribbean countries for the period 2001-2012. The paper is interesting because it focuses on capital buffer fluctuations over the business cycle only in emerging countries. They conclude that capital buffers are more likely to fluctuate pro-cyclically in countries where capital regulation is less stringent and the costs of adjusting buffer holdings are lower, while the larger the institution, the lower the capital buffer.

4. INTRODUCING THE COUNTERCYCLICAL BUFFER INTO URUGUAYAN REGULATION

The objective of this section is to discuss how far the countercyclical capital buffer proposed in Basel III can coexist with the statistical (or dynamic) provisioning currently in force in Uruguay's banking regulation.

As mentioned in Section 2, statistical provisioning conceptually has an institution specific prudential dimension given that the formula set out in regulations governing the growth of the statistical provisioning fund depends on idiosyncratic variables for each bank, especially, the growth of credit granted by each institution and the stock of credit granted.

In particular, the formula used for generating statistical provisions is as follows:[4]

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[4] The regulations are described in *Comunicación,* number 2001/149 <www.bcu.gub.uy/Comunicados/seggco01149.pdf>; *Comunicación,* number 2012/004, *Actualización,* number 190 <www.bcu.gub.uy/Comunicados/seggco12004.pdf>; y *Comunicación* no. 2014/061, *Actualización,* no. 200 <www.bcu.gub.uy/Comunicados/seggco14061.pdf>.

$$\Delta FPE_t = \begin{cases} \left\{\left[\frac{1}{12}\sum_{i=1C}^{2B}\alpha_i\left[C_{i,t-1}-C_{1,t-13}\right]+\frac{1}{12}\sum_{i=1C}^{2B}\beta_i C_{i,t-1}\right]-\left\{\left[\sum_{i=1C}^{i=5}(\Delta E_{i,t})\right]-R_t\right\}\right\}\kappa_t \\ \qquad\qquad\qquad\qquad \text{if} \quad \sum_{i=1C}^{2B}\left[C_{i,t-1}-C_{1,t-13}\right]\geq 0 \\ Min\left[-\left\{\left[\sum_{i=1C}^{i=5}(\Delta E_{i,t})\right]-R_t\right\};0\right] \qquad \text{if} \quad \sum_{i=1C}^{2B}\left[C_{i,t-1}-C_{1,t-13}\right]<0 \end{cases}.$$

Where:

- ΔFPE_t is the increase (positive or negative) in the statistical provisioning fund during month *t*.
- $\left[\frac{1}{12}\sum_{i=1C}^{2B}\alpha_i\left[C_{i,t-1}-C_{i,t-13}\right]+\frac{1}{12}\sum_{i=1C}^{2B}\beta_i C_{i,t-1}\right]$ is the statistical loss for defaulted loans corresponding to month *t*.
- $\sum_{i=1C}^{2B}\left[C_{i,t-1}-C_{1,t-13}\right]$ is the change between month $t-1$ and $t-13$ in the stock of computable risks.
- $\sum_{i=1C}^{5}\left[\Delta E_{i,t}\right]-R_t$ is the net result of defaulted loans once statistical provisions of month *t* have been established, *i* is the category of credit risk (risk rating), $\Delta E_{i,t}$ are net charges for specific provisions and R_t represents recoveries of defaulted loans in month *t*.

The *k* parameter adjusts changes in the fund according to the relative distance of each fund with respect to its ceiling or limit, in such a way that *k* tends towards zero as the fund nears the ceiling. Meanwhile, the countercyclical capital buffer set out in Basel III has a system-wide prudential dimension, given that the regulation proposed for activating a higher capital requirement is based on the behavior of the financial system as a whole. The recommendation of the Basel Committee on Banking Supervision (2010a) is that the buffer should depend on the long-term trend of the aggregate private sector credit-to-GDP gap. Thus, from the point of view of the dimension or main focus, both instruments complement one another perfectly.

Another aspect that lends support to the coexistence of both instruments stems from the fact that statistical provisions are aimed at protecting banks against expected losses during the cycle, while the objective of the countercyclical capital buffer is to protect them against unexpected losses. Thus, the fact that both instruments are monitoring the business cycle gives them the character of dynamic instruments as they are constantly addressing the evolution of risk.

It should also be pointed out that statistical provisions in Uruguay, which are very similarly designed to those in Spain's regulation, have effectively fulfilled the role of buffer in addressing losses during recessions and smoothing the volatility of economic results. However, they have been largely ineffective in reducing the growth of credit.

In fact, empirically it can be seen how in the case of Spain the operation of dynamic provisioning did not influence to any great degree the rapid growth of credit observed during the phase preceding the recent crisis in its banking system, when credit grew at an annual average rate of 16%. However, dynamic provisioning did function adequately as a buffer against the large losses of the crisis period.[5]

Furthermore, it should be remembered that in Spain dynamic provisions are calculated in a similar way to in Uruguay, i.e.:

$$\textit{Dynamic provisioning} = \alpha\Delta\,\textit{Credit} + \beta\,\textit{Credit} - \textit{Specific provisioning}.$$

The great difference is that specific provisioning is guided by incurred loss criteria. According to the latter, a specific

[5] See these effect in BBVA (2011).

provision can only be registered if there is objective evidence of deterioration in the asset or loan. Thus, specific provisions, which are subtracted in the calculation of dynamic provisions, are very small at the time of a boom phase given that the objective evidence referred to in the regulation does not exist. This meant dynamic provisions grew sharply in Spain during the boom phase.

Given that specific provisions in Uruguay are guided by the principle of expected loss, they are much larger, meaning the statistical provisioning fund might not increase during the boom phase in the case of institutions that grow substantially. This situation has been corrected recently by introducing the following clause into current regulations: "If as a result of applying the preceding parameters at the end of month $t-1$ the statistical provisions fund does not increase –in total value and as a percentage of the maximum limit for computable risks– with respect to month $t-13$, having increased the stock of computable risks during said month, institutions may use higher statistical provision parameters in order for the fund to increase in line with that set out. The Superintendency of Financial Services can –in accordance with the observed evolution of the statistical provisions fund– issue instructions to the institutions requesting them to comply with the aforementioned objective".

Current regulations therefore ensure that the statistical provisioning fund will increase, in total value and as a percentage of the maximum limit, but it is evident that this could be insufficient for addressing expected losses during recessions, and it clearly does not perform a significant role in determining the speed of credit growth either.

In contrast, an increase in the capital requirement acts much more directly, effectively restricting credit as long as the increase in capital is demanding enough to ensure the amount of capital above the regulatory minimum is very low.

In sum, both instruments– statistical provisioning and countercyclical capital buffers– can coexist and are tools that complement one another for the following reasons:

- In general, statistical provisions have an institution specific prudential dimension and countercyclical capital a system-wide prudential dimension.

- Statistical provisioning is effective for facing expected losses, while countercyclical capital is useful for addressing unexpected losses.

- Even in the case of Spain, with a specific provisioning approach based on incurred losses, dynamic provisioning was not successful in restraining credit growth. In the case of Uruguay, with a specific provisioning approach based on expected losses, the role of dynamic provisioning in curbing credit growth becomes even more important.

- Countercyclical capital acts faster in reducing credit growth as the amount of capital above regulatory requirements becomes smaller.

Finally, one factor to take into consideration for introducing the countercyclical capital buffer concerns international standards on matters of financial institution regulation and supervision emerging from Basel III and international accounting standards issued by bodies such as the IASB (International Accounting Standard Board). The countercyclical capital buffer is a standard that has now been approved by Basel III, while dynamic provisioning, although considered within the recommended prudential tools, has still not been specifically enacted by Basel III.

In this respect Basel III establishes that the use of more forward-looking provisions should be promoted. It therefore advocates a change in accounting practices towards basing provisioning on an expected loss approach, and not one of incurred loss. To such ends it has published and made available to the IASB a set of principles aimed at modifying IAS 39. However, even if agreement is reached on the expected loss approach, the maximum horizon on which it could coordinate with the IASB for assessing expected loss would be one year, and never one business cycle.

Thus, for better adjustment to the international standards mentioned above, it is also recommendable to introduce the countercyclical capital buffer.

5. ACTIVATION, ADJUST AND DEACTIVATION: AN EXAMPLE WITH HISTORICAL DATA FOR URUGUAY

This section describes the results of applying the methodology set out by the Basel Committee on Banking Supervision (2010b) for the activation, adjust and deactivation of the countercyclical capital buffer using historical data for Uruguay. The methodology is described first, together with the data used to apply it and assess the outcomes for the period prior to the banking crisis of 2002. Indicators are also proposed that could guide deactivation of the buffer, highlighting the main advantages and disadvantages of the methodology.

5.1 Methodology: Aggregate Private Sector Credit-to-GDP Gap as a Reference

The Basel Committee on Banking Supervision (2010a, 2010b) has suggested using the gap between the aggregate private sector credit-to-GDP ratio and its long-term trend as a reference for the phase of the financial cycle. To determine the size, activation and deactivation of the countercyclical capital buffer they suggest following a three step process:

Step 1. *Calculate the aggregate private sector credit-to-GDP ratio,* taking into account a broad measure of credit that captures all the sources of private sector borrowing and is applied equally to all banks with similar exposure without considering their individual contribution to excess credit growth.

Step 2. *Calculate the credit-to-GDP gap (the gap between the ratio and its trend) using* the Hodrick-Prescott (1981) filter with a *lambda* parameter of 400,000 (Borio and Lowe, 2002) which reflects the prolongation of financial cycles as compared to traditional business cycles.[6]

Step 3. *Transform the credit-to-GDP gap into the guide buffer add-on* associating the size of the capital buffer with the magnitude of the gap calculated in step 2 according to the following approach:

$$\begin{cases} Buffer = 0\% & \text{if } Gap_t < 2\% \\ Buffer = 2.5\% * \left[\dfrac{(Gap_t - 0.02)}{0.08} \right] & \text{if } 2\% < Gap_t < 10\% \\ Buffer = 2.5\% & \text{if } Gap_t > 10\% \end{cases}$$

The size of the buffer varies linearly between 0% and 2.5% for gap values of between 2% and 10%. For values of less than 2% the buffer should not be activated, while for values of 10% or more it should be at its maximum level of 2.5%.

An alternative approach to the one presented above has recently been introduced in Switzerland. The Swiss National Bank (2014) describes the methodology used by the institution. Activation and adjustment of the countercyclical buffer is based on an historical analysis of the relevant series, for instance the credit gap calculated in step 2 along with other variables. In particular, it identifies a past period of instability (or crisis) and, based on the historical evolution of the series, builds a buffer that adjusts gradually over three years up to a maximum of 2.5% 12 months before the relevant indicator reaches its maximum level (the time of greatest imbalance, such as, for instance, the outbreak of the crisis, t^*). The size of the buffer adjusts linearly in line with the size of the gap according to the following approach:

[6] Drehmann et al. (2010) provide empirical evidence revealing that trends calculated with this parameter perform well in describing the long-term behavior of the private borrowing series.

$$
\begin{cases}
Buffer = 0\% & \text{if } Gap_t < Gap_{t^*-16} \\
Buffer = 2.5\% * \left[\dfrac{\left(Gap_t - Gap_{t^*\ 16}\right)}{\left(Gap_{t^*-4} - Gap_{t^*\ 16}\right)} \right] & \text{if } Gap_{t^*-16} < Gap_t < Gap_{t^*-4} \\
Buffer = 2.5\% & \text{if } Gap_t > Gap_{t^*-4}
\end{cases}
$$

5.2 Selecting Historical Data for Uruguay

The Superintendency of Financial Services publishes monthly aggregate data on gross loans to the private sector. The series contains data starting from 1999. In accordance with the suggestions of the Basel Committee on Banking Supervision (2010b), and given the characteristics of the Uruguayan banking sector, this series emerges as the best option to use. Nevertheless, due to the need for obtaining a longer historical series it was decided to bridge the aforementioned series with historical data taken from the Bank's internal sources and that includes loans to the private sector from public and private banks.[7] Figure 1 shows both series. The data is generally, but not precisely coherent. To carry out the final bridge it was decided to maintain the value of the public series and adjust the oldest series by differences. The right hand figure presents the final bridged series that will be used in the study.

The main disadvantage of this procedure stems from the arbitrary nature of the adjustment. The current series is used by the government and has passed through a monitoring process. For this reason, unchanged data is used. The previous series is from the internal source and has not undergone the same verification process.[8] In light of the aforementioned, it was decided to adjust this series in order to achieve an bridge consistent with the current series.

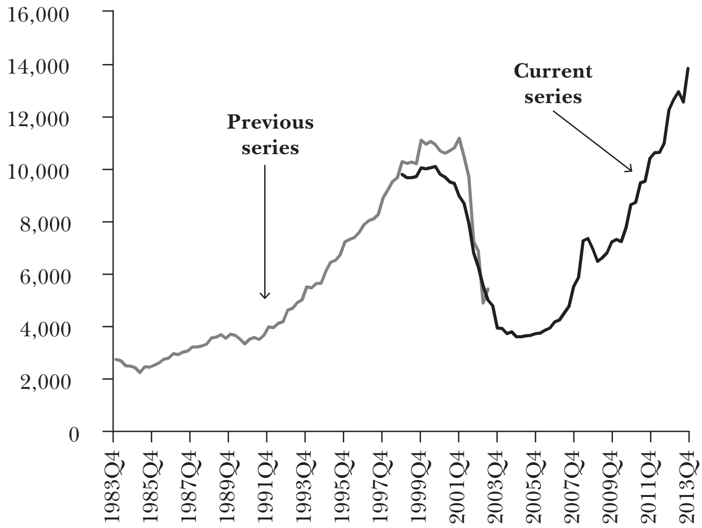
[7] Series from the Siste system, numbers 7251, 7384, 7390. Due to the fact that data was in thousand nuevos pesos the exchange rate series number 182 was used to transform it.

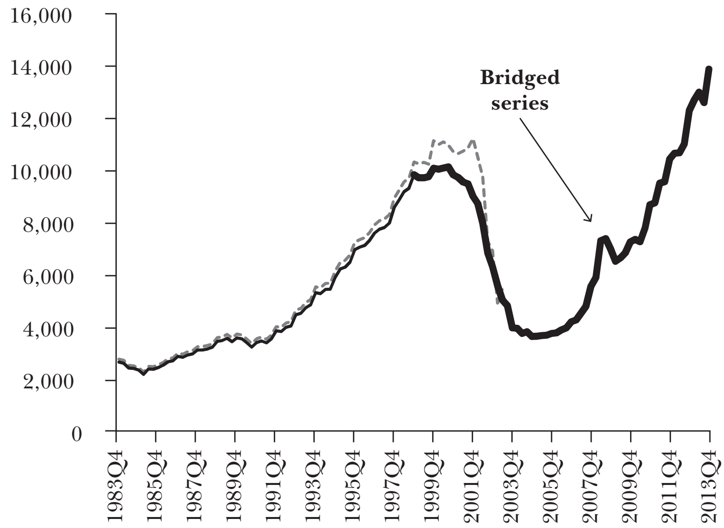
[8] In particular, data from state banks is of low quality for years before 1999.

Figure 1

AGGREGATED CREDIT TO PRIVATE SECTOR

(millions of dollars)





One specific characteristic of Uruguay's economy concerns the granting of loans in domestic as well as foreign currency. Figure 2 shows the performance of the trend-cycle component of the credit to the private sector series by currency and of the aggregate series. As can be seen, although the cycles of the series by currency have exhibited differing behaviors during Uruguay's recent history, the aggregate series properly captures the behavior of the trend-cycle component of the desegregated series. It therefore seems reasonable to use the aggregate credit series as the main reference for the authorities' decision-making, but complementing them with analysis of the disaggregated series.

The bridged GDP series provided by the Banco Central del Uruguay (BCU) are shown in figure 3.

The final series for aggregate private sector credit-to-GDP is shown in figure 4.

Figure 2

TREND-CYCLE OF CREDIT TO PRIVATE SECTOR, BY CURRENCIES

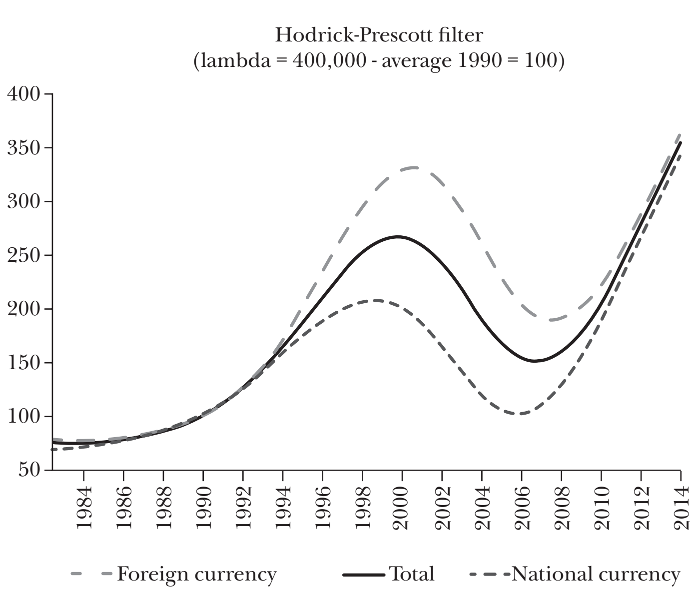
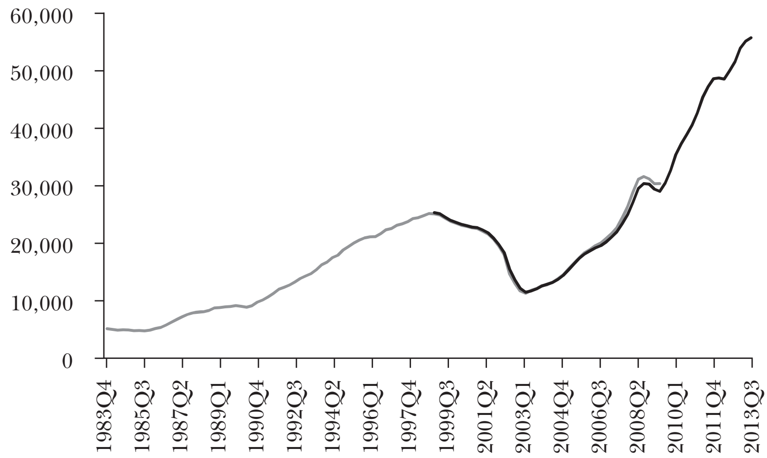


Figure 3

GROSS NATIONAL PRODUCT

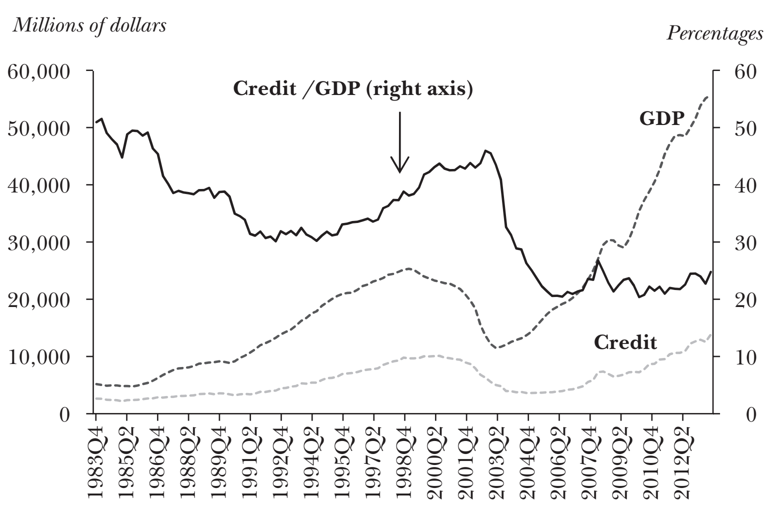
(millions of dollars)



Note: bridged series provided by BCU.

Figure 4

AGGREGATED CREDIT TO PRIVATE SECTOR AND GDP



5.3 Constructing the Credit-to-GDP Gap

To select the indicator that best adjusts to episodes of systemic risk in Uruguay's financial system, various alternatives were assessed using different adjustment values for the cycle (the classic λ value of 1,600 and the 400,000 proposed for financial series), the methodology suggested by the Swiss National Bank and an *ad hoc* measure of 35% for the credit-to-GDP ratio as a fixed reference (corresponding to the average of the historical series). The outcomes were analyzed and the capacity of these early warning indicators for anticipating the financial difficulties experienced in 2002 was assessed. Figure 5 summarizes the outcomes of applying the Hodrick-Prescott (1981) filter with $\lambda = 400{,}000$ (for the bridged series as well as the current series) and $\lambda = 1{,}600$ for the bridged series. The estimated gap corresponds to the grey curve named *Cycle*. As we can see, the outcome is very sensitive to the λ parameter as well as the length of the data. Periods in which the ratio is above trend correspond to periods of strong credit growth. In particular, calculation of the gap with $\lambda = 400{,}000$ gives reasonable results, with a less volatile cycle, and clearly identifying the episode of instability in 2002.

5.4 Determining the Countercyclical Buffer

To compare the outcomes and extract conclusions on the indicators employed, the countercyclical capital buffer was calculated according to the following approaches:

1) In line with the methodology set out by the Basel Committee on Banking Supervision (2010b), using the three series described previously:

   a. Bridged series 1983Q4-2013Q4 with $\lambda = 400{,}000$.
   b. Current series 1999Q1-2013Q4 with $\lambda = 400{,}000$.
   c. Bridged series 1983Q4-2013Q4 with $\lambda = 1{,}600$.

2) Calculating the gap over and ad hoc trend fixed at 35%.

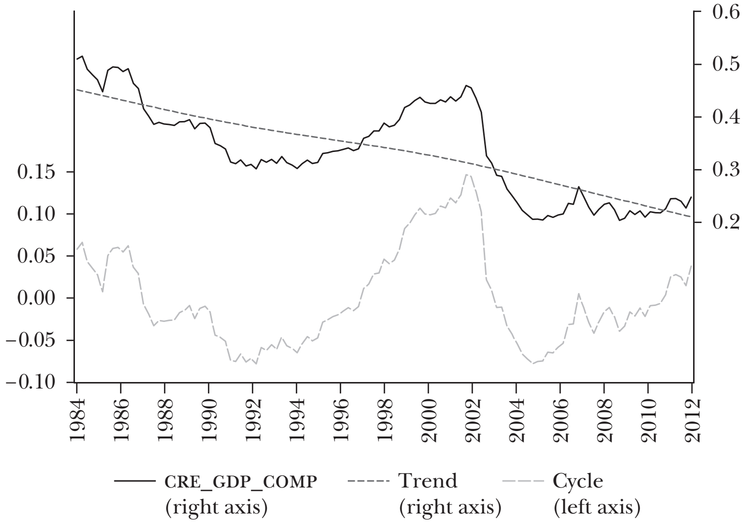
3) In line with the methodology proposed by the Swiss National Bank (2014).

Figure 5

CREDIT GAP TO GROSS NATIONAL PRODUCT

(percentages of GDP)

HODRICK-PRESCOTT FILTER (LAMBDA = 400,000)



HODRICK-PRESCOTT FILTER (LAMBDA = 400,000)

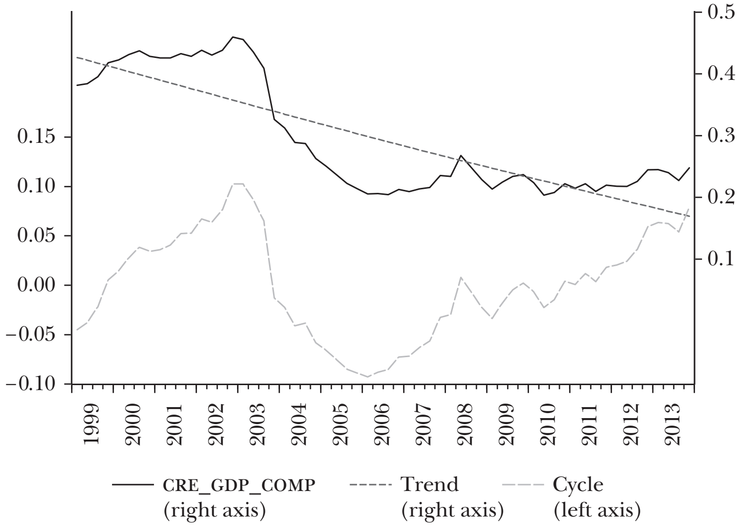


Figure 5 (cont.)

CREDIT GAP TO GROSS NATIONAL PRODUCT

(percentages of GDP)

Hodrick-Prescott filter (lambda=1,600)

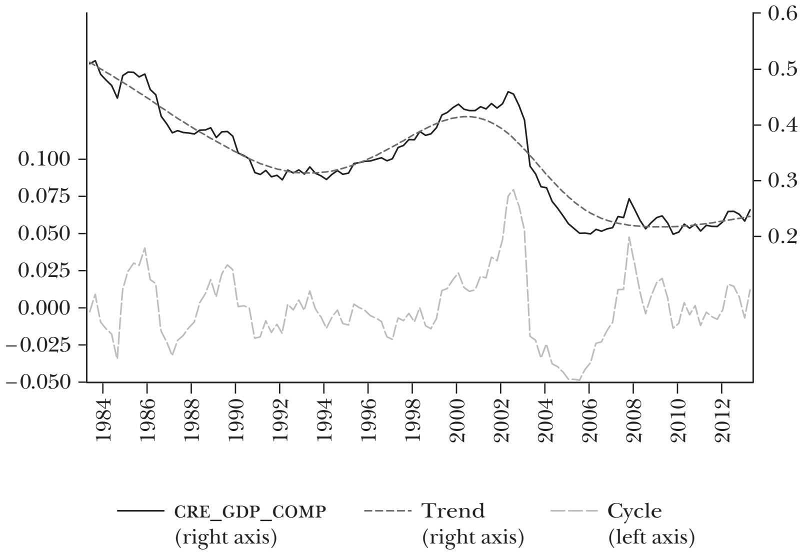
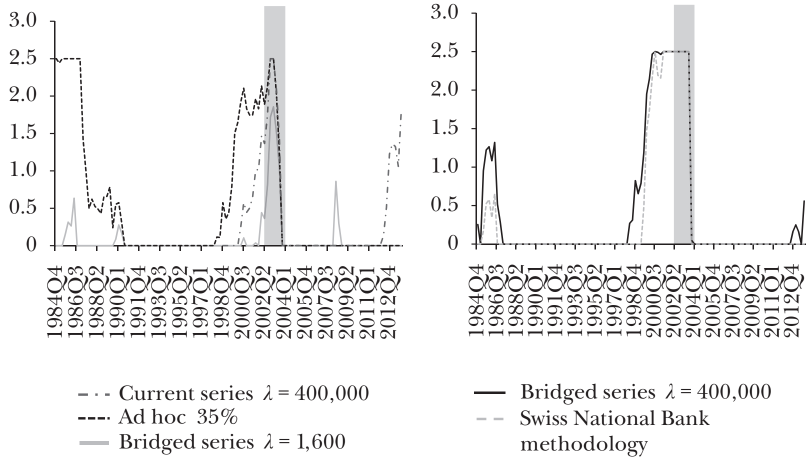


Figure 6 shows the countercyclical capital buffers resulting from the five cases described above. The analysis focuses on the capacity to anticipate the crisis of 2002 (shown with a vertical line) and the issuing of false alarms. All five indicators anticipate the period of instability to a certain degree. However, in the case of the relatively short data series (since 1999) with $\lambda = 400{,}000$, the buffer does not reach the maximum until after the outbreak of the crisis. The length of the series is therefore insufficient because it does not manage to anticipate the emergence of risk far enough in advance to allow capital reserves to be built up. The gap derived from this series would also be indicating an increase of systemic risk in the Uruguayan economy at present, which does not coincide with an assessment of the current situation in Uruguay's financial system. Moreover, the gap calculated with the bridged

Figure 6

COUNTERCYCLICAL CAPITAL BUFFER

(percent)



series and $\lambda = 1{,}600$ identifies significant deviations during the international crisis of 2008 that coincide with periods of turbulence in the international environment, but do not correspond with periods of financial instability, thereby constituting a false alarm. Its anticipation of the 2002 financial problems is also poor. On the other hand, the ad hoc approach of 35% as ratio trend, appropriately anticipates the crisis of 2002, but produces a significant false alarm in the first quarters of data, casting doubt on its efficiency. The buffer that emerges from using $\lambda = 400{,}000$ for the bridged series, meanwhile, seems to provide appropriate signs for the timing and magnitude of the 2002 crisis (see the right hand panel of figure 6). It starts generating signals four years before the crisis, reaching the maximum countercyclical capital buffer six quarters prior to the outbreak of the crisis. Finally, good results are also obtained with the approach proposed by the Swiss National Bank. The indicator begins to produce signals 10 quarters before the crisis, reaching the maximum value 12 months prior to the start of the crisis (by construction). It does not give any false alarm during the period considered either. The main disadvantage of this approach is that it is designed to capture crisis of the same type and size as that of 2002, but not necessarily periods of instability in general. Its principal advantage stems from the ease and practicality of its calculation.[9] [10]

In sum, the bridged series filtered with value $\lambda = 400{,}000$ represents the most appropriate indicator to use as a guide for determining the countercyclical buffer. It is also recommended that the methodology proposed by the Swiss National Bank be used as a complement in order to give a more complete assessment.

Drehman and Tsatsaronis (2014) argue in favor of using this indicator as a guide for setting the countercyclical capital buffer. They also answer the criticism of Repullo and Saurina (2011)[11] by finding a positive or insignificant correlation between the credit gap and GDP in relevant periods for implementing the countercyclical buffer, when the latter would have had a positive impact for smoothing financial cycles.

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9 The model estimated shows that the buffer should have been activated in the second quarter of 1998. In fact, Newsletter 1613 of 29/09/98 raised the minimum capital requirement for financial institutions from 8% to 10%. This measure is comparable to the activation of a countercyclical capital buffer because it was triggered by the excessive growth of credit during the four preceding years. It could therefore be stated that a countercyclical capital buffer guided by the referred activation and deactivation criteria would have operated in a similar way to the measures taken in 1998.

10 Drehmann and Tsatsaronis (2014) suggest using the methodology of AUC and ROC for evaluating different early warning indicators. Nevertheless, as they themselves argue, these methodologies present problems in small samples given that statistical evaluation of one country in particular is complicated due to the limited number of crises (in the case of Uruguay the data covers just one crisis period). It is therefore not applicable to this case.

11 Repullo and Saurina (2011) argue that the credit gap is countercyclical to the growth of GDP and, therefore, would tend to exacerbate rather than smooth fluctuations in GDP.

To determine the capacity of the indicator in the case of Uruguay's economy we carry out the assessment according to the approaches proposed by Drehmann and Juselios (2014): timing, stability and interpretability. As we saw previously, the credit gap constructed based on the bridged series meets these three requirements: *i)* it produces signals four years prior to the period of instability and reaches the buffer maximum at least one year in advance, *ii)* the signal is stable and increases as the period of instability draws nearer, and *iii)* it can be directly interpreted given the simplicity of constructing the indicator and its direct connection with financial cycles and the functioning of the financial system.

However, there are some limitations concerning the construction of the series, and the methodology of the filter, that should be taken into account when using this indicator as a guide. First, the *end-point* problem of the Hodrick-Prescott (1981) filter is a weakness of this methodology and, therefore, estimation of the gap for the later periods is subject to a significant standard deviation. Second, the problem of the *starting-point* of the series: as Gersl and Seidel (2012) point out, the trend calculation can depend significantly on the starting point of the series, particularly in short series. This criticism applies to the case of Uruguay. As stated previously, the outcome varies considerably when using the bridged series. Moreover, we found false alarms in the first periods, which could be due to the lack of preceding data (especially data from the crisis of 1982). Drehmann *et al.* (2014) recommend using at least 10 years of data in order to minimize this problem. Third, the Hodrick-Prescott filter is a backward-looking filter and therefore calculates recursively as new data is incorporated. This can generate changes in the results, which should be studied more deeply by the analyst. Fourth, the most effective indicator emerges from using the bridged credit-to-GDP series. The bridge is carried out arbitrarily and the series prior to 1999 has not been exposed to a verification process similar to that of the current government series.

5.5 Deactivation of the Countercyclical Capital Requirement

The next step consists of selecting the indicators to signal deactivation of the countercyclical capital buffer. The Basel Committee on Banking Supervision (2010b) sets out the following principles for identifying said indicators:

1) When there are losses in the banking system that pose a risk to financial stability.

2) When there are problems in other sectors or areas in the financial system that could potentially disrupt the normal flow of credit and threaten financial system stability.

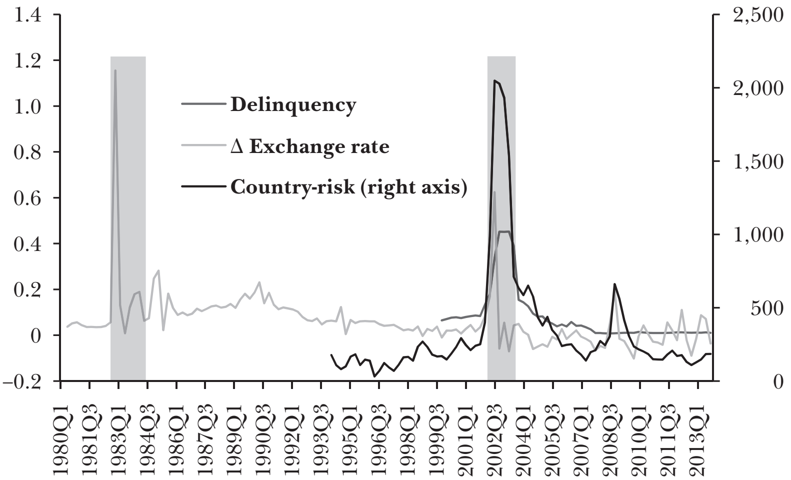
The Basel Committee on Banking Supervision suggests three indicators: profits before tax, credit spreads and TED spreads. Given the characteristics of Uruguay's banking system and the availability of data three others are also proposed: delinquency in the banking system, the exchange rate and sovereign default risk. These three indicators are closely linked to past episodes of financial instability and are relevant for identifying systemic risk. Figure 7 shows the evolution of these variables.

Historical evidence shows that the three series produce appropriate signals coinciding with the period of instability. They therefore correctly capture the start of the crisis and could be monitored together to guide the decision for deactivating the countercyclical capital buffer. Other suggested indicators that could be incorporated refer to credit conditions, although the data currently available is not sufficient to assess the capacity of such series for producing appropriate signals. Indicators associated with loan portfolio quality could also be included, such as, for instance, the percentage of loans with the lowest risk ratings (loans with ratings of 3, 4 or 5),

Although it is important that the regulator uses this information for deciding on buffer deactivation, it is also important that they are willing to do so faster or even immediately once risks have materialized. By deactivating the countercyclical capital buffer immediately the authorities allow banks to

Figure 7

INDICATORS FOR THE DEACTIVATION OF THE COUNTERCYCLICAL CAPITAL BUFFER



make use of it to cover losses incurred during times of stress, reducing the need to affect minimum capital requirements or other buffers. Thus, immediate deactivation of the countercyclical buffer immediately helps to reduce the risk of the supply of credit being severely constrained by regulatory capital requirements and, thereby, helps the banking sector to continue providing its services to support the rest of the economy.

In sum, the aggregate private sector credit-to-GDP gap overlapped since 1984 and calculated using a $\lambda = 400{,}000$ parameter fulfills the principal characteristics we look for in an indicator for guiding countercyclical capital buffer adjust and activation decisions. It also manages to answer the main criticism of the indicator by being relatively stable, exhibiting positive properties as an early warning indicator (detects the crisis and does not produce false alarms in the period studied) and easy to measure and calculate. Notwithstanding the aforementioned, it should be used with other indicators, particularly those suggested by the Swiss National Bank.

6. FINAL COMMENTS

This paper analyzed the main characteristics of the countercyclical capital buffer and studied its inclusion in Uruguayan regulation, paying special attention to how it complements other regulatory tools such as statistical provisioning. It concludes that designing and introducing a countercyclical capital buffer in accordance with the principles of Basel III would complement existing regulations for the following reasons: *i)* it introduces a dynamic dimension that depends on the phase of the business cycle to static capital requirements (or minimum requirements such as the capital buffers established by Basel III); *ii)* statistical provisions, in their current calculation formula, have an institution specific prudential dimension whereas the countercyclical capital buffer has a system-wide prudential dimension; *iii)* by definition, statistical provisions address expected losses in the financial cycle whereas the countercyclical capital buffer addresses unexpected losses; *iv)* although it is not its main objective, the countercyclical capital buffer is more effective in restraining credit growth than dynamic provisioning, as demonstrated during the recent crisis in Spain's financial system and even in Uruguay towards the end of the 2000s.

Although activation, adjust and deactivation decisions for the countercyclical capital buffer should be at the discretion of the regulatory authorities, these should also be guided by the appropriate data. In the case of Uruguay, the aggregate private sector credit-to-GDP gap overlapped since 1984 and calculated using a $\lambda = 400{,}000$ parameter fulfills the most important characteristics needed from an indicator for guiding countercyclical capital buffer decision-making. Nevertheless, it is recommended that said reference indicator be complemented by other indicators for specific sectors and the banking system as whole.

Although the countercyclical capital buffer can be deactivated gradually in accordance with the behavior of risks present in the system, the possibility of faster or even immediate release should be included. Immediate deactivation of the countercyclical capital buffer helps to reduce the risk of the supply of credit being severely constrained by regulatory capital requirements and, thereby, helps the banking sector to continue providing its services to support the rest of the economy.

Thus, design and implementation of a countercyclical capital buffer following the principles in Basel III would complement current regulation and serve as an extra tool to mitigate systemic risk in the Uruguayan banking sector. From the point of view of the institutional distribution of responsibilities for applying this requirement, the fact that it is a prudential regulation tool means it should be handled by the Superintendency of Financial Services. However, given the systemic nature of the risk it is aimed at, it would be recommendable for activation, adjust and deactivation decisions to take into account the vision of systemic risk arising from deliberations under the framework of the Financial Stability Committee.

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Comparing the Transmission of Monetary Policy Shocks in Latin America: A Hierarchical Panel VAR

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This paper assesses and compares the effects of monetary policy shocks across Latin American countries that put in practice the Inflation Targeting scheme (Brazil, Chile, Colombia, Mexico and Peru). An estimated Hierarchical Panel VAR allows us to use the data efficiently and, at the same time, exploit the heterogeneity across countries. Monetary shocks are identified through an agnostic procedure that imposes zero and sign restrictions. We find a real short run effect of monetary policy on output (with a peak around 12-15 months); a significant medium run response of prices with the absence of the so-called price puzzle and a hump-shaped response of the exchange rate, i.e. weak evidence of the so-called delayed overshooting puzzle phenomenon. Nevertheless, we find some degree of heterogeneity on the impact and propagation of monetary shocks across countries. In particular, we find stronger effects on output and prices in Brazil and Peru relative to Chile, Colombia and Mexico and a stronger reaction of the exchange rate in Brazil, Chile and Colombia relative to Mexico and Peru. Finally, we present a weighted-average impulse response after a monetary shock, which is representative for the region.

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  Fanny Warman
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  Keywords: Prudential regulation, currency mismatches, exchange rate regimes, Latin America, Caribbean.

- ***Intermediarios financieros no bancarios en América Latina: ¿banca paralela?***
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