Commodity prices and the business cycle in Latin America: Living and dying by commodities?*

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Abstract

We analyze the dynamic interactions between commodity prices and output growth of the seven greatest exporters Latin American countries: Argentina, Brazil, Colombia, Chile, Mexico, Peru and Venezuela. Using Markov-switching impulse response functions, we find that the responses of their respective output growths to commodity price shocks are time dependent, size dependent and sign dependent. Overall, the major evidence of asymmetries in output growth responses occurs when commodity price shocks lead to regime shifts. Accordingly, the design of optimal counter-cyclical stabilization policies in this region should take into account that the reactions of the economic activity vary considerably across business cycle regimes.

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1. Introduction

The 2008-2009 global downturn has shown Latin America's continuing dependence on primary commodities. Between 2002 and 2008, Latin America Countries (LAC) benefited greatly of the more persistent and intense increase in primary commodities since the eighties. This period corresponds with quarterly GDP growth rates situated steadily about 2% for the major LAC. As documented in Figure 1, even in the middle of the 2008 world-wide recession, some LAC still presented relatively high growth rates while commodity prices remained at record heights. However, the collapse in commodity prices in mid-2008 left LAC cruelly exposed to the world decline in economic activity. During the international recession, LAC exhibited quarterly growth rates that were far away from their historical records.

In most of the existing literature, the analysis of the reactions of LAC output growth to commodity price shocks is developed within linear frameworks (Österholm and Zettelmeyer, 2007, or Izquierdo, Romero and Talvi, 2008). However, one salient feature of both LAC output growth and commodity prices is their strong non linear cyclical behavior in terms of their own dynamics and in terms of the relation between these two variables. Some recent papers have pointed out this nonlinear behavior. For instance, Jerzmanowski (2006) or Misas and Ramirez (2007) show that output growth in LAC (among others) countries, varies considerably across business cycle regimes when they are modeled as a non linear characteristic of the data, and Arango and Melo (2006) detect nonlinear business cycle dynamics in the industrial production indexes of Brazil, Colombia and Mexico. Regarding commodity prices, Cashin, McDermott, and Scott (2002) and Reitz and Westerhoff (2007) find empirical evidence of the cyclical non linear dynamics of commodity price developments. Finally, Hamilton (2003) suggest that the relation between output growth and commodity prices (in particular, oil prices) is nonlinear, and Cerra and Saxena (2008) stress that dummy variables representing financial crisis are crucial to model with annual panel data the asymmetric effects of a set of explanatory variables (which includes commodity prices) on output growth.

According to this strand of the literature, we propose a reduced-form Markov-switching model to examine the nonlinear reactions of output growth to commodity prices shocks in the seven largest LAC, Argentina, Brazil, Chile, Colombia, Mexico, Peru, and Venezuela. There are several important contributions worth mentioning. First, to conduct the business cycle analysis, we use the bridging method of Camacho and Perez Quiros (2010) to estimate LAC series of quarterly GDP growth rate at monthly frequency from the information of monthly indicators. Since the method deal with mixed frequencies (quarterly and monthly series) and with ragged ends (indicators that start later or end sooner than the rest), it additionally permits enlarging the original series of GDP in those countries with short time series.

Second, we use several types of commodity price indexes which reinforce the results obtained in the empirical analysis. We started the analysis with the general composite indices of *Moodys*, and of *The Economist*, and its disaggregation in Food, Non-food and Metals. Using these aggregate indexes to analyze the effects of price shocks on economic activity presents the advantage of being easily available. However, composite indices might not be a good measure since the key commodities for particular countries may change significantly across time. To overcome this potential drawback, we alternatively used in the analysis the country-specific measures of commodity export and import prices computed by Cunha, Prada and Sinnott (2010).

Third, we use modern techniques for business cycle analysis to show that output growth and commodity prices presents clear non linearities that the analysts need to take into account when trying to relate the implications of changes in prices to short term growth or long term level of output. For the long term implications of commodity prices on output, we employ non linear cointegration techniques and we fail to detect a long term relation between output and commodity prices. This could be interpreted as evidence in favour of considering the shocks to commodity prices as having only transitory effects on the Latin American economies.

Fourth, to examine the potential nonlinearities in the transmission of shocks from commodity prices to output we consider a novel extension of the Markov-switching impulse responses which avoids some drawbacks of previous proposals. Contrary to Ehrmann, Ellison, and Valla (2003) who propose regime-dependent responses, and to Karame (2010) who supposes that the regime at the time of the shock is known with probability one, our impulse responses are calculated at any point in history. Hence, the case of regime-dependent responses and the case where the shocks occur in a particular regime become special cases of our method. In this context, the paper can also be viewed as a methodological extension of the literature on nonlinearities in univariate time series of output growth and commodity price growth to the multiple equation case where the evolution of the business cycles is determined endogenously.

Fifth, using the Markov-switching impulse responses we obtain that although commodity price shocks consistently show procyclical behavior regardless to the model used in the analysis, the results highly support the hypothesis that the responses are nonlinear. We find that output reactions to commodity price shocks are sign-dependent since the reactions to positive shocks do not mirror those from negative shocks. We also find that the responses are size-dependent since they are scaled by factors higher than proportional for larger shocks. Finally, find that the responses are time-dependent since the propagation of price shocks hitting the economies in recessions are notably different from the propagation of those shocks hitting the economies in expansions. In particular, the magnitude of the nonlinearities is of special interest for those shocks that are able to produce regime switches.

Our results lead to dramatically important policy implications. Policy makers should respond asymmetrically to positive versus negative commodity price shocks of the same size, they should not adopt proportional reactions to shocks of different sizes, and they should react differently to similar commodity price shocks when they affect LAC in different phases of their business cycles.

The paper is structured as follows. Section 2 presents the preliminary analysis of data. Section 3 examines the cointegration relationships between output and commodity prices, assesses the nonlinearities in the time series and in the reactions of output growth to price shocks. Finally, this section studies the propagation of commodity price shocks to GDPs within a Markov-switching framework. Section 4 concludes and points out some lines of further research.

2. Preliminary data analysis

Although real GDP is usually adopted as the single best measure of national aggregate economic developments, the empirical analyses of the time series of LAC output growths exhibit several problems. For most countries, the series are too short to make standard statistical inference since they start at the beginning of the 90s. In addition, they are only available at quarterly frequencies which make it difficult the comparison with monthly time series such as commodity prices.

Two solutions have been proposed in the literature on business cycle analyses. The first solution, adopted, for example, by Arango and Melo (2006), consists on using monthly series of economic activity such as industrial production. It has the advantage of usually being longer than GDP and available monthly, but the disadvantage of representing only a small fraction of the aggregate economic activity of some LAC. The second solution, used for example in Aiolfi, Catao and Timmermann (2007), is to build monthly coincident indicators of economic activity. These authors employ traditional approximate factor models to extract indexes that underlie a wide set of indicators. However, they require balanced panels and their method ignores the information contained in quarterly indicators such as real GDP. In addition, although these indexes are computed as linear combinations of meaningful economic indicators, the fact that they are not related with a particular variable of interest, make it difficult to find and economic interpretation of their movements.

To overcome these limitations, we adopt an alternative strategy that consists on converting the quarterly GDP growth rates into larger series of quarterly growth rates at monthly frequencies by using monthly economic indicators. For this purpose, we use the extension of the single-index dynamic factor model of Stock and Watson (1991) proposed by Camacho and Perez Quiros (2010) which deal with mixed frequencies (quarterly and monthly series) and with ragged ends (indicators that start later or end sooner than the rest).

Let us assume that GDP and the monthly indicators were available at monthly frequencies and observed without missing data. Let y_t and g_t be the quarterly and monthly growth rates of GDP, respectively. According to Mariano and Murasawa (2003), the quarterly growth rates of a flow series can be expressed as the following averaged sum of lagged monthly growth rates

$$y_t = \frac{1}{3} g_t + \frac{2}{3} g_{t-1} + g_{t-2} + \frac{2}{3} g_{t-3} + \frac{1}{3} g_{t-4}. \quad (1)$$

Let z_t be the k-vector of monthly growth rates of the indicators used in the model. Now, let us assume that the monthly growth rates of GDP and the set of indicators admit a factor model. In this case, each variable can be written as the sum of two stochastic components: a common component, f_t, which represents the overall business cycle conditions, and an idiosyncratic component, which refers to the particular dynamics of the series. To define the dynamic properties of the model, the underlying business cycle conditions are assumed to evolve with AR(p_1) dynamics

$$f_t = \rho_1 f_{t-1} + ... + \rho_{p1} f_{t-p1} + e_t, \quad (2)$$

where $e_t \sim iN(0, \sigma_e^2)$. The evolution of GDP monthly growth rate is assumed to depend linearly on f_t and on its idiosyncratic dynamic component, u_t^g, which evolves as an *AR(p_2)* process:

$$g_t = \beta_g f_t + u_t^g, \quad (3)$$

$$u_t^g = d_1^g u_{t-1}^g + ... + d_{p2}^g u_{t-p2}^g + \varepsilon_t^g, \quad (4)$$

where $\varepsilon_t^g \sim iN(0, \sigma_g^2)$. In addition, it is assumed that the k monthly indicators can be expressed in terms the common factor and the idiosyncratic components that are autoregressive processes of p_3 orders:

$$z_t^i = \beta_i f_t + u_t^i, \quad (5)$$

$$u_t^i = d_1^i u_{t-q}^i + ... + d_{p3}^i u_{t-p3}^i + \varepsilon_t^i, \quad (6)$$

where $\varepsilon_t^i \sim iN(0, \sigma_i^2)$, and i = 1,2,…,k. Finally, all the shocks e_t, ε_t^g, and ε_t^i, are assumed to be mutually uncorrelated in cross section and time series dimensions. Stated in this way, we show in the Appendix how to estimate this model by maximum likelihood using the Kalman filter.

So far, we have assumed that all the variables included in the model are always available at monthly frequencies for all time periods.[1] Although this assumption seems

[1] Note that quarterly data are only observed in the third month of the respective quarter. In addition, some indicators start too late while others are available with some lags.

quite unrealistic, Mariano and Murasawa (2003) show that the system of equations stated in the Appendix as if all the time series where always observed remains valid with missing data after a subtle transformation. These authors propose to replace the missing observations with random draws from a distribution that cannot depend on the parameter space of the Kalman filter. Skipping details, this method permits all the matrices to be conformable, leaving the likelihood unchanged up to a scale since the rows containing missing data in the Kalman matrices will be skipped from the updating recursion.[2]

In the selection of LAC monthly indicators we make use of the seminal proposal of Stock and Watson (1991). They included four key indicators: one of the supply side (Industrial Production), one of the demand side (Retail Sales) and one from the income side (Personal Disposable Income), which are combined with an employment series (Employment in non-agricultural sectors) to create an indicator of economic activity. Table 1 presents the series used for each country and the sample period in which each series is available. The quarterly series are seasonally adjusted and used in quarterly growth rates to avoid unit root problems. Monthly series are used in annual growth rates to diminish the effects of seasonal patterns and noisy signals. However, as shown in the appendix, we carefully take into account the fact that an annual growth rate is a moving average of monthly growth rates. Due to data availability constraints, we use Unemployment instead of Employment in all countries but Argentina and Venezuela. In these countries, we use quarterly growth rates of Employment which are treated in the Kalman filter in the same way as the quarterly growth rates of GDP.

Table 2 presents the maximum likelihood estimates of the loading factors (β_g and β_i) for each country (standard errors within parentheses) which capture the correlation between the unobserved common factor and the indicators. As expected, the signs of the loading factors are positive for all indicators but Unemployment suggesting that the common factor can be viewed as an index of broad economic activity. For most of the cases, the loading factors are statistically significant and the larger loading factors are those corresponding to GDP and Industrial Production. This fact points out its high explanatory power as direct indicator of the economic activity. Employment in

[2] Interested readers can check the details in Camacho and Perez Quiros (2010)

Argentina and Sales in Mexico also belong to the set of indicators with larger correlations with their respective common factors. Finally, the last column of Table 2 shows the percentage of the variance of GDP that is explained by the common factor. The high percentage of the variance of GDP explained by the factor in all countries reinforces the interpretation of the factor as the monthly estimate of economic activity.

Figure 1 plots the monthly estimates of GDP quarterly growth rates along with their actual values which are displayed as plot marks in the third month of each quarter. The figure helps the reader to understand the advantages of our proposal in the analysis of business cycles: GDPs quarterly growth rates, which are issued quarterly, are converted to monthly observations and the time series are extended to the larger extension of the longest available series. In accordance with the methodology employed in this paper, the Kalman filter anchors monthly estimates to actual GDP growth when this is observed. Hence, for those months where GDP is known, the actual value and the estimates coincide.

As a last remark in this section, it is worth describing the time series of commodity prices used in the paper. We started the analysis with the general composite indices of *Moodys*, and of *The Economist*, and its disaggregation in Food, Non-food and Metals. Using these aggregate indexes to analyze the effects of price shocks on economic activity presents the advantage of being easily available.

However, in analyzing the commodity price conditions faced by individual LAC countries at different periods of time, general composite indices might not be a good measure since the key commodities for particular countries may change significantly across time. To overcome this potential drawback, we alternatively used in the analysis the country-specific measures of commodity export and import prices computed by Cunha, Prada and Sinnott (2010). These price indexes periodically recalculate commodity weights and therefore reflect changes in the country's trade flows and exports structure. Figure 3 shows the country-specific and Figure 4 shows the general composite indexes in quarterly growth rates.[3] Although these series are potted for the

[3] Quarterly growth rates of commodity prices are required to develop a balanced comparison with quarterly growth rates of national outputs.

period 1971.01-2009-03, the effective sample employed in the empirical analyses is adapted to the sample of the series of output used for each country.

3. Output responses to commodity price shocks

The analysis of output responses to commodity price shocks is developed in three stages. First, we examine the long-run relationships between output and commodity prices. Second, we point out that the dynamics of these series are nonlinear. Third, we propose Markov-switching impulse responses to capture these nonlinearities.

3.1. Analysis of cointegration

Analyzing the (if significant) sign of the long-term effects of shocks to commodity prices (among other external factors) on the economic activity of LAC has been the source of many debates. On the one hand, a drop in commodity prices may lead to increase the real exchange rate and, consequently to increase aggregate demand and income. On the other hand, if the institutional environment of a country is not adequate, the country can fall in rent seeking affecting negatively the long-term growth. Which forces will dominate output reactions to commodity price shocks in LAC? This section uses pairwise cointegration tests of output and commodity prices to look for empirical evidence regarding this effect.

In the presence of cointegration, there exists a long-run attractor in the dynamics of output and prices which implies that possible disturbances are not purely transitory. Table 3 displays the test statistics for three cointegration tests. The first test is the well-known Engle and Granger (1987) cointegration test for which the null of no cointegration is rejected if the statistics are lower than -3.42. The second test is the Stock and Watson (1988) test of common trends for which the null of two stochastic trends (no cointegrating relationship) versus one common stochastic trend is rejected if entries are lower than -8. However, these linear methods may fail to detect cointegration due to misspecification problems when the true nature of the adjustment process is nonlinear. To overcome this potential problem, Table 3 also includes the nonparametric cointegration analysis advocated by Bierens (1997) whose results are independent of the data-generating process due to the nonparametric nature of this approach. In this case, the null of no cointegration is rejected if the statistic is greater than the critical value of 0.0169.

Doubtless, the result on cointegration analysis is that the null hypothesis of no cointegration and that the null of no common stochastic trends cannot be rejected at 5%. Almost uniformly, the test statistics uniformly lie in the non rejection areas for all the countries, all the prices and all the methods employed in this analysis. According to these results, we conclude that the relation between commodity prices and output only captures temporary effects. This could be interpreted as evidence in favour of considering the shocks to commodity prices as having only transitory effects on the Latin American economies. Hence, no error correction term will be added to the multivariate specifications of output growth and commodity prices growth that will be analyzed in the article.

3.2. Assessing the need of nonlinear models

The dynamics of the time series used in this paper, which are plotted in Figures 1, 2 and 3, exhibit some special features. Although the series fluctuate around their respective means, the broad changes of direction in the series, which show pronounced drops and subsequent recoveries, seem to mark business cycle patterns with asymmetric features in terms of the duration and amplitude of the business cycle phases. During the downturns, as in the mid nineties in Argentina and Mexico or in the last part of the sample in almost all countries, the growth rates go deeply from positive to negative. Output growth developments show that business cycle expansions are more persistent than recessions and the transition between states seems to be sharp.

Despite the marked non linear cyclical pattern of the series, most of the studies of LAC output growths and external factors have focused on linear relations, probably because moving to nonlinear frameworks is costly. Nonlinear algorithms are sometimes burdensome and there are much less statistical results available for nonlinear models. Therefore, before moving to nonlinear specifications, we need to gather statistical evidence in favor of these potential nonlinearities.

A natural approach to model the business cycle behavior of output and prices is the regime switching model proposed by Hamilton (1989). Following his seminal proposal, we assume that the switching mechanism of a time series at time t, w_t, is controlled by an unobservable state variable, s_t, that is allowed to follow a first-order Markov chain. Thus, a simple switching model may be specified as:

$$w_t = c_{s_t} + \sum_{j=1}^{p} \phi_j w_{t-j} + \varepsilon_{wt}, \tag{7}$$

where $\varepsilon_{wt} \sim iidN(0,\sigma_w^2)$. The nonlinear behaviour of the time series is governed by c_{s_t}, which is allowed to change within each of the two distinct regimes $s_t = 0$ and $s_t = 1$. The Markov-switching assumption implies that the transition probabilities are independent on the information set at t-1, χ_{t-1}, and on the business cycle states prior to t-1. Accordingly, the transition probabilities from state i to state j are

$$p(s_t = j / s_{t-1} = i, s_{t-2} = h, ..., \chi_{t-1}) = p(s_t = j / s_{t-1} = i) = p_{ij}. \tag{8}$$

The maximum likelihood estimates of parameters, which are obtained by regressing the time series on a switching mean, are reported in Table 3.[4] Overall, they show that in the regime represented by $s_t = 0$, the average growth rate is positive (estimates range from 0.46 to 1.44 for output and 5.72 to 56.65 for prices), so we can interpret this regime as the expansion period. By contrast, with the exception of the country-specific price of Venezuela, the average is negative (estimates go from -0.21 to -12.18 for output and from -1.86 to -2.83 for prices) in the regime represented by $s_t = 1$, so we can interpret this regime as the recession period. In addition, each regime is highly persistent, with estimated probabilities of one regime being followed by the same regime of about 0.9 although the persistence of slumps is higher in the case of commodity prices. Again with the exception of Venezuela, whose results greatly depend on the sharp and deep slowdown in 2002, the estimated parameters of all the models are in line with the estimated parameters for non Latin American economies.[5]

[4] According to the results of Camacho and Perez Quiros (2007), we do not necessarily need to include lags in the dynamics of the shocks because the Markov-switching specification may account for all the time series autocorrelation.

[5] Graphs of filtered and smoothed probabilities are available from the authors upon request.

Within regime switching models, testing for nonlinearities consists on testing the null hypothesis of one state against the alternative of two. These tests are not straightforward due to the presence of nuisance parameters under the null which leads the standard asymptotic not to be valid.[6] Hansen (1992) proposes a bounds test that is valid in spite of these difficulties. In particular, he shows that the likelihood ratio test statistic for the null hypothesis of one state is the supremum over all admissible values of the nuisance parameters (the transition probabilities). The p-values of this test, which are reported in Table 4 for lag lengths p of 0 and 1, show that the null hypothesis of no switching is overwhelmingly rejected for all the national outputs and commodity prices time series.

In an independent contribution, Carrasco, Hu, and Ploberger (2009) propose an optimal test to examine whether the parameter of a model change according to Markov-switching dynamics. The advantage of this test is that it only requires estimating the model under the null hypothesis where the parameters are constant. Table 4 also shows the empirical p-values which are computed from 1,000 iterations for a sample size equal to the size of the original data set. Reinforcing the result obtained from Hansen's test that the dynamics of the series of output growth and commodity prices growth are nonlinear, the null of linear parameters is again overwhelmingly rejected.[7]

3.3. Assessing the need of nonlinear responses

Having detected strong evidence of nonlinearities in the dynamics of output and prices, let us move to examine the potential nonlinearities in the transmission of shocks from commodity prices to output. As a first approach to detect these nonlinearities, we develop a twofold exercise. In the first analysis, we examine the potential asymmetries in the marginal effects of price changes to output dynamics. For this purpose, the flexible framework proposed by Hamilton (2001) constitutes an ideal starting point

[6] Note that the transition probabilities p and q are not identified under the null.
[7] Our results differs from Cashin et al. (2002) who found little evidence of a relationship between the NBER referenced cycles in the US and cycles in commodity prices. The reasons for the discrepancies could presumably be related to the fact that we analyze LAC business cycles and that we use different approaches to check for the potential nonlinearities in commodity prices.

since it permits a broad change of nonlinear alternatives. Let y_t be the series of output growth, let π_t be the vector of explanatory variables which may include lags of the series of interest, exogenous variables, and their respective lags. Let $\otimes$ be the element by element product. Skipping technical details, the method is based on estimating the flexible semiparametric model

$$y_t = \mu(x_t) + \varepsilon_t = a + b' x_t + \lambda m(\theta \otimes \pi_t) + v_t, \tag{9}$$

where $v_t \sim iidN(0, \sigma_v^2)$, and θ is a vector of parameters that governs the curvature of the nonlinear function *m*. In the empirical analysis, we include among π_t the series of commodity prices growth rates. The parameter λ marks the degree of nonlinearity in the transmission of shocks π_t into output growth. Accordingly, the other natural test of nonlinearity would be based on testing the null hypothesis that $\lambda = 0$. Hamilton (2001) shows that the asymptotic regarding this test is not standard due to the presence of nuisance parameters and he describes a procedure to test for the null of linearity by using resampling methods. The *p*-values of the tests based on 1,000 replications, are reported on Table 5. They reveal that with some few exceptions, the effects of price changes on output growth are nonlinear for all countries and almost all prices. Among the exceptions, this method fails to reject nonlinear reactions of output growth in Argentina and Brazil in the case of country-specific price shocks.

Apart from the large amount of linearity rejections, Figure 5 helps to illustrate some key features of the nonlinear nature of the relationship between commodity price shocks and output reactions. This figure plots the estimates of the expected output growth for different values of country-specific price changes in Argentina and Mexico.[8] In both countries output growth increases when commodity prices grow although the reactions of output to prices are very different from each other. In the case of Argentina for which linearity in the responses was not rejected by using the test advocated by Hamilton (2001), the figure shows that both positive and negative shocks seem to have proportional effects on GDP growth rate. However, in the case of Mexico there is a clear asymmetric transmission of positive versus negative price growth rates to output growth. Although (large) commodity price reductions imply growth of GDP

[8] The Brazilian figure is similar to the Argentinean figure. Figures for Chile, Colombia and Venezuela are similar to that of Mexico. To save space, they are omitted from the text but are available upon request.

contractions, the decrease in output growth is higher than proportional. By contrast, the reaction of output growth to positive price changes is clearly diminished.

Although the previous approach is very intuitive and flexible enough to account for nonlinearities of different nature, it does not specifically address the potential Markov-switching asymmetric dynamics across the business cycle phases. This may diminish the power of flexible nonlinear tests against this particular type of nonlinearity. In addition, the flexible functional form model may underestimate the nonlinearities in the responses of output growth to commodity price shocks that come from time-varying responses. Of special interest in our research, it is worth noting that this method is unable to differentiate responses to shocks that occur in the course of an expansion from those that occur in the course of a recession.

To illustrate the importance of considering time-varying responses, Figure 6 (left-hand-side chart) displays the four-year-window rolling responses of Argentinean output growth to one-standard-deviation shocks in its country-specific index of commodity prices. The responses are successively computed from bivariate linear VAR models as output reactions to price shocks hitting the economy in the months that go from 2006.04 to 2009.03.[9] According to the figure, the instantaneous responses of output growth to commodity price shocks have restricted to be zero but they exhibit hump-shaped paths in the following periods. At a few months after the shocks, the responses climb to their maximum values and exhibit a substantial decline since then.

Noticeably, the responses of output growth are not time invariant. In the last part of the sample, the responses of output are about twice as large as those computed in the first part of the sample (the breakpoint is about 2008.06) although all of then are calculated from shocks of the same size and sign. Interestingly, the different features in the responses are roughly coincident with the two phases of the business cycle exhibited by the Argentinean output growth (right-hand-side chart) in these years. The lower responses of output growth to positive commodity price shocks refer to the period 2006.04-2008.03 when quarterly output grew about 2%. The responses of output growth

[9] Since we failed to detect cointegration between output and prices, the benchmark models employed in this analysis were VAR (1) specifications where error correction terms were not considered. Identification is achieved by using the standard Cholesky decomposition of the covariance matrix of the residuals.

to commodity price shocks are substantially higher when output growths were much lower or even negative (period 2008.04-2009.03). This business cycle feature in the responses of output growth to commodity price shocks leads us to consider Markov-switching impulse responses.

3.4. Markov-switching impulse responses

The traditional tools employed in the literature to examine the propagation of shocks to commodity price growth (variable x_t) into output growth (variable y_t) are the impulse-response functions.[10] They can be computed by simulating the effects of a shock to x_t (called ε_t) on the conditional forecast of y_t. In linear models, the impulse response of y_t at horizon h to shocks in x_t of magnitude δ, can be defined as the estimated difference between the expected realizations of y_{t+h} and a baseline "no shock" scenario:

$$IRF(h,d) = E\left(y_{t+h} / \varepsilon_t = \delta\right) - E\left(y_{t+h}\right) , \tag{10}$$

where $E(\bullet)$ is the expectation operator.[11]

Figures 7 to 13 display the output reactions to commodity price shocks in Argentina, Brazil, Chile, Colombia, Mexico and Venezuela, respectively. In particular, the shocks δ hitting the systems are set to d times the standard deviation of commodity prices, with d being ±1, ±3 and ±6. The responses to positive shocks are on the left-hand side graphs while the responses to negative shocks are on the right-hand side graphs. To account for the possibility of correlation of the errors across different equations in the VAR(1) specifications, the impulse response functions have been orthogonalized with the commodity price growths ordered last.[12] For comparison purposes, the figures show the linear responses in the first two graphs of each figure. They show that commodity price shocks evoke responses on output growth of the same sign. Therefore, positive shocks in commodity prices lead to expansionary responses in output whereas negative commodity price shocks are followed by periods of output slowdowns. In some

[10] To save space, this section is concentrated on country-specific commodity prices. Results for composite commodity price indexes are qualitatively similar and they are available from the authors upon request.
[11] All shocks in intermediate periods between t and $t+h$ are set equal to zero for convenience.
[12] Therefore, commodity price shocks are restricted to have no contemporaneous impact on output growth.

countries such as Brazil, Chile, Colombia and Venezuela, a one-standard-deviation commodity shock (equal to about 5 percent in one quarter) leads to about ⅓ percentage point change in Latin American growth after two quarters, and about 1 percentage point change after four quarters. This is significant bearing in mind that commodity prices rose by on average over 20 percent a year over 2004-2007.

However, based on our previous findings, two shortcomings of the linear responses can be assessed from the figures. First, the responses of output growth to commodity price shocks are not shock-dependent. The reaction of output growth is symmetric since +1-standard-deviation shocks have exactly the opposite effect of shocks of -1-standard-deviation shocks. In addition, 3-standard-deviation shocks have exactly three times the effect of 1-standard-deviation shocks. Second, the responses of output growth are not history-dependent. Shocks occurred in recessions are expected to change output growth in the same manner as if the shocks occurred in expansions.

On the contrary, within the Markov-switching framework described in previous sections we can assess the business cycle asymmetries in the impact of commodity price shocks on output growth. For this purpose, let us define the following Markov-switching impulse responses which are allowed to be regime dependent, and to account for nonlinearities in the output reactions to positive versus negative shocks and to large versus small shocks. Let us assume that the output growth series and the commodity price growth series are driven by an unobserved process, s_t, which evolves according to the Markov-switching statistical properties stated in (8). Let Y_t be the bivariate specification of output growth and commodity prices growth, $Y_t = (y_t, x_t)'$. Let C_{s_t} be the vector of regime-dependent constants and let A be the matrix of autoregressive parameters. Finally, let U_t be the vector of reduced form shocks and let B be the Cholesky decomposition of its covariance matrix. Assuming a lag length of one, the autoregressive representation can be stated as

$$Y_t = C_{s_t} + AY_{t-1} + U_t, \qquad (11)$$

where $U_t \sim iidN(0,\Omega)$. In contrast to linear VAR(1) specifications, the vector C of constants is now conditional to the state.[13]

[13] The analysis of state-dependent autocorrelation and covariance matrices is left for further research.

The Markov-switching responses of output growth for an arbitrary reduced form shock to commodity price shocks of size δ and history w_{t-1} can be computed as:

$$MSIRF(h, \delta, w_{t-1}) = E(y_{t+h} / \varepsilon_t = \delta, w_{t-1}) - E(y_{t+h} / w_{t-1}). \qquad (12)$$

This conceptualize an experiment where we investigate the time profile of the effect on output growth of a shock of size δ hitting the commodity price at time t as compared with a baseline where no shocks hit the system. It is worth pointing out that in contrast to (11), expression (12) is history dependent, i.e., it depends on the "history" w_{t-1} or initial values of the variables in the model which also determine the probability of occurrence of the business cycle states. Post-multiplying the responses to reduced-form shocks by B, one can obtain the orthogonalized responses to structural shocks.

To fully understand the nature of the business cycle nonlinearities which are accounted for by the Markov-switching responses presented in (11) and (12), a point worth carefully describing is the way to compute the expectations. Calling $\xi_{t/t}$ the (2×1) vector whose jth element is $p(s_t = j/\chi_t)$, its optimal h-period-ahead forecast conditional on information available at date t is $\xi_{t+h/t} = P^h \xi_{t/t}$, where P is the matrix of transition probabilities whose (i,j) element is p_{ij}. Now, let the (2×1) vector $\Gamma_{i.t+h}$ be the h-period-ahead forecasts of the ith variable whose jth element is the forecast conditional to the state j.[14] Then the unconditional h-period-ahead forecasts of the ith variable can be computed as $E(Y_{i,t+h} / s_t) = \xi'_{t+h/t} \Gamma_{i,t+h}$ which are easy to compute once the vector $\xi_{t/t}$ is inferred from the model. In the baseline case where no shocks hit the system, $\xi_{t/t}$ coincides with the filtered probabilities at time t.

In the case where a shock hits the system at time t, the vector of probabilities after the shock, $\xi^*_{t+h/t}$, can be inferred from the model as well. Let Y_t^* be the value of the variables after the shock, let $\eta(Y_t^*)$ be the vector whose ith element is the conditional density function of variable i, let $\tilde{1}$ be the (2×1) vector of ones, and let $\otimes$ represent the element-by-element multiplication. The inference of the states at time t, $\xi^*_{t/t}$, can be computed as

[14] They can be computed sequentially from ith element of the vector $Y_{t+h} = C_j + AY_{t+h-1}$.

$$\xi_{t/t}^{*}=\frac{\xi_{t/t-1}\otimes\eta\left(Y_t^{*}\right)}{\tilde{1}'\left(\xi_{t/t-1}\otimes\eta\left(Y_t^{*}\right)\right)}. \qquad (13)$$

The path followed by the inferred probabilities and the forecast of the variables after the shock can be computed as in the case of no shock.

The reasons why the Markov-switching transmissions of shocks depend on the sign and the size of the shocks and on the history of the variables can be assessed from this expression. Large and positive shocks will increase the probability of expansion and will reduce the probability of recession and the value of the variables when $s_t = 0$ are overweighed when computing expectations. In addition, the state probabilities will react to the size of the shocks in a nonlinear manner. Finally, the history or value of the time series up to the time of the shock will be crucial to compute the time paths of the responses.

Two recent proposals of the literature can be seen as special cases of our impulse response analysis.[15] First, Ehrmann, Ellison, and Valla (2003) study the response of the system conditionally to the regime (e.g. $s_{t+h}=j$ for all h) in which the shock occurs but they require the assumption that there is no more change in regime in the wake of the shock. This occurs when we assume that $\xi_{t+h/t}$ is a vector with one in the position j and zeroes elsewhere for all h. Second, Karame (2010) propose Markov-switching impulse responses that capture the possible different impact of a shock depending on the regime in which it occurs. However, this author supposes that the regime at the time of the shock is known with probability one (e.g. $s_0 = j$). By contrast, our impulse responses are calculated at any history w_{t-1}, being the case where the shocks occur in a particular regime a special case.

The Markov-switching reactions of output growth to shocks in commodity price growth rates in the seven LAC are examined in Figures 7 to 13. To asses the degree of business cycle asymmetries in the responses, the effects of shocks are computed on the conditional "history" of being close to each one of the two different states of the business cycle when the shocks hit the system. In particular, shocks are assumed to hit

[15] Other approaches are Artis, Krolzig and Toro (2004) who examine responses to changes regime, and Markku, Lutkepohl and Maciejowska (2010) who consider switching covariances but they assume that the responses are invariant across states.

each country at the periods that correspond to its highest filtered probability of recession (second row of graphs) and to its lowest filtered probability of recession (third row of graphs). Finally, the last row of graphs in each figure examines the business cycle consequences of commodity price shocks. For each figure, the graphs show the evolution of the national probabilities of recession both after the shock (plot of the second element of $\xi^{*}_{t+h/t}$) and under the assumption of no shock (plot of the second element of $\xi_{t+h/t}$). To be sure that the shocks have relevant business cycle consequences, the probability responses are computed after a large expansionary shock of +6 standard deviations when the probability of recession is highest (left-hand-side graphs) and after a large contractionary shock of -6 standard deviations when the probability of recession is lowest (right-hand-side graphs). For expositional purposes, let us classify the results into those which are common to the vast majority of LAC and those which are country specific.

General features on the Markov-switching responses. As in the case of linear responses, commodity price shocks are procyclical since they are followed by output reactions of the same sign. Noticeably, nonlinear responses become similar to the linear responses in the cases of low shocks, positive shocks in expansions, and negative shocks in recessions. However, the Markov-switching responses of output growth to commodity price shocks are strongly supportive of the hypothesis that responses are size dependent, sign dependent and history dependent. As an illustrative example regarding these nonlinear features, let us concentrate in the case of Brazil (Figure 8).

Output reactions are size dependent since they are usually scaled by higher than proportional factors for larger shocks. The reaction of output growth to positive commodity price shocks hitting the Brazilian economy at its highest filtered probabilities of recession (second row at left-hand side) constitutes an illustrative example. It can be observed that responses to three-standard-deviation shocks are about three times the responses to one-standard-deviation shocks. However, six-standard-deviation shocks produce disproportionately larger expansionary responses of output growth. We will show that this effect is due to the change in regime induced by the large price shock.

These graphs also provide evidence of asymmetry in the effects of positive versus negative commodity price shocks on output growth. Attending to the left-hand-side-graphs, it is noticeable that output growth reactions to positive price shocks (second graph) do not mirror those reactions to negative price shocks. By contrast, it seems that output responds more strongly to positive large shocks than to negative shocks, especially in low growth states.

Finally, Figure 8 constitutes an encouraging piece of evidence on whether commodity price shocks have different effects on output growth depending on whether the economies are in expansion or recession. Overall, the expansionary impacts on output arising from positive commodity price shocks are larger when the shocks occur in the course of recessions. In addition, the largest output reductions generated by negative price shocks occur when they arrive within expansions.

The two bottom graphs will help us to fully understand the mechanism behind the asymmetries that are accounted for by the time-varying. At the moment of the shock, the left-hand-side graph shows that Brazil was in recession with filtered probability of almost 1. If no shock hit economy, the inferred probability of recession displayed on the top would follow the typical stationary path towards its ergodic value. However, after a large expansionary commodity price shock of six-standard-deviation size, the probability of recession decreases to about 0.06. This shift in regime is associated with a larger than proportional expansionary reaction on output growth (see the responses displayed in the second left-hand graph). Noticeably, if the same shock affected the economy at its lowest probability of recession (third left-hand graph), the responses of output growth would be proportional to the size of the shock as in the case of linear responses. In fact, a general feature of LAC output growth responses is that the major evidence of asymmetries occurs when commodity price shocks lead to regime shifts.

Country-specific features on the Markov-switching responses. Despite the common features on the responses of output to price shocks outlined above, Figures 7 to 13 also suggest some country specific features on output growth responses to commodity price shocks. First, the output growth reactions are larger in magnitude in

Brazil, Chile, Colombia and Venezuela than in Argentina, Mexico, and to less extent Peru.

Second, the asymmetric interaction across regimes between output growth and commodity price shocks in Argentina (Figure 7) is very particular. This country is singular in the sense that the reaction of its output growth to commodity price increases in expansions or to commodity price decreases in recessions is negligible.[16] Shocks to commodity prices are only propagated to output growth when the (very large) negative shocks imply changes in regime from recessions to expansions and to less extent from expansions to recessions.

Third, Colombia is the country that exhibits the lowest asymmetries in the responses of output growth to commodity price shocks. Figure 10 shows that the Markov-switching responses are not very different from the impulse responses computed from linear models. In spite of this comment, the nonlinearities in the responses of Colombian output growth to commodity pieces shocks are observed in the relatively higher reaction to positive shocks in recessions and in the ability of price shocks to produce regime shifts.

Fourth, the business cycles identified by the Markov switching in Peru (Figure 12) and Venezuela (Figure 13) are dominated by their respective output growth dynamics. As accounted for by the bottom graphs for these countries, even for large commodity price shocks the evolution of the probabilities of recession with and without the commodity price shocks are very similar. This diminishes the asymmetries across business cycles and the asymmetries that come from the responses of output to different signs. However, there are still asymmetries in the effects of large versus small shocks in comparison to those computed from linear models.

Fifth, the bottom graphs in the case of Brazil (Figure 8) reveal that the switches in recession probabilities which are due to commodity prize shocks are much larger when price increase in recessions than when prices decrease in expansions. Accordingly, the relative expansionary reactions of output growth to positive

[16] In these cases, output reactions are obviously dominated by their own shocks.

commodity price shocks in recessions are much larger than the contractionary reactions of output growth to negative commodity price shocks in expansion when they are compared with the linear results. Hence, the bad news that comes if commodity prices decrease in the curse of an expansion have relatively lower impact in the Brazilian economy than the good news arrived when in the curse of a recession commodity prices increase unexpectedly.

5. Conclusion

Although assessing the effect of external factors on LAC output growth has been the source of an extensive debate, the specific role of commodity prices affecting the business cycles of these countries has not frequently been investigated, and in these cases the baseline frameworks have been linear models. In addition, some recent proposals detect asymmetries between recession and expansions in output and commodity prices. Noticeably, they have concentrated on univariate analyses of nonlinear time series such as output (Jerzmanowski, 2006; Misas, and Ramírez, 2007), industrial production (Arango and Melo, 2006) and commodity prices (Cashin, McDermott, and Scott, 2002). However, we think that examining the effects of commodity price shocks to output growth, which is crucial in the design of counter-cyclical stabilization policies in this region, is essentially nonlinear *and* multivariate.

In this paper, we add to the previous contributions further evidence regarding the nonlinear behavior of output growth and commodity prices growth in the seven great exporters LAC (Argentina, Brazil, Colombia, Chile, Mexico, Peru and Venezuela). Interestingly, even though we use tests that are robust to the presence of nonlinearities, we fail to detect a long term relation between output and commodity prices. This could be considered as empirical evidence in favor of the arguments regarding the transmission mechanism of shocks which consider that increases in commodity prices are short-term demand shocks instead of being the main driving force of the long-term level of GDP.

In addition, we seek to develop a multivariate Markov-switching model that accounts for time-dependent transmission of commodity price shocks to output growth.

Using this model, we provide an encouraging piece of evidence on the nonlinear nature of the common evolution of output and commodity prices. We assess that although commodity price shocks are procyclical, their effects on output growth depend on the state of the economy, the size of the shock and the sign of the shock. Noticeably, the major evidence of asymmetries occurs when commodity price shocks imply regime shifts. In this sense, large positive commodity price shocks hitting the economies in recessions lead to larger than proportional expansionary effects on output growth. However, it is also true that negative price shocks have dramatic consequences on expected output growth when they arrive in the course of expansions.

The model used in this paper provides a solid foundation for starting a line of research trying to explain the specificities in the asymmetric behaviour of each country and up to what point adequate fiscal or monetary reactions to commodity price help to accommodate and smooth the effect of the commodity price shock. We consider that these extensions are important enough to leave them for further research.

Appendix

According to expressions (1)-(4), GDP quarterly growth rates, y_t , are

$$y_t = \beta_x\left(\frac{1}{3}g_t + \frac{2}{3}g_{t-1} + g_{t-2} + \frac{2}{3}g_{t-3} + \frac{1}{3}g_{t-4}\right) + \left(\frac{1}{3}u_t^g + \frac{2}{3}u_{t-1}^g + u_{t-2}^g + \frac{2}{3}u_{t-3}^g + \frac{1}{3}u_{t-4}^g\right), \quad \text{(A1)}$$

and the annual growth rates the *i*th monthly indicator, Z_{it} , are

$$Z_{it} = \beta_i \sum_{j=0}^{11} f_{t-j} + u_t^i \,, \quad \text{(A2)}$$

with i = 1, 2, …, k. The model can easily be written in state space representation which can then be estimated by using the Kalman filter. Without lost of generalization, let us assume that the model contains GDP and only one indicator which are collected in the vector $\psi_t = \left(y_t \,, Z_{it}\right)'$. Let us also assume that $p1$ = $p2$ = $p3$ = 1. In this case, the observation equation, $\psi_t = H\alpha_t$, is

$$\begin{pmatrix} y_t \\ Z_{it} \end{pmatrix} = \begin{pmatrix} \frac{2\beta_g}{3} & \frac{\beta_g}{3} & \beta_g & \frac{\beta_g}{3} & \frac{2\beta_g}{3} & 0 & \cdots & 0 & \frac{2}{3} & \frac{1}{3} & 1 & \frac{1}{3} & \frac{2}{3} & 0 \\ \beta_i & & & \cdots & & & \cdots & \beta_i & 0 & & \cdots & & 0 & 1 \end{pmatrix} \begin{pmatrix} f_t \\ f_{t-1} \\ \vdots \\ f_{t-11} \\ u_t^g \\ \vdots \\ u_{t-5}^g \\ u_t^i \end{pmatrix}. \quad \text{(A3)}$$

The transition equation, $\alpha_t = T\alpha_{t-1} + \eta_t$, is

$$\begin{pmatrix} f_t \\ f_{t-1} \\ \vdots \\ f_{t-11} \\ u_t^g \\ \vdots \\ u_{t-5}^g \\ u_t^i \end{pmatrix} = \begin{pmatrix} \rho_1 & \cdots & 0 & 0 & 0 & & \cdots & 0 \\ 1 & \cdots & 0 & 0 & 0 & & & 0 \\ & & & & & & & \\ 0 & \cdots & 1 & 0 & & & & \vdots \\ & \cdots & & & d_1^g & \cdots & 0 & \\ & \cdots & & & \ddots & & & \\ 0 & \cdots & & & 0 & 1 & 0 & 0 \\ 0 & \cdots & & & \cdots & & 0 & d_1^i \end{pmatrix} \begin{pmatrix} f_{t-1} \\ f_{t-2} \\ \vdots \\ f_{t-12} \\ u_{t-1}^g \\ \vdots \\ u_{t-6}^g \\ u_{t-1}^i \end{pmatrix} + \begin{pmatrix} e_t \\ e_{t-1} \\ \vdots \\ e_{t-11} \\ \varepsilon_t^g \\ \vdots \\ \varepsilon_{t-5}^g \\ \varepsilon_t^i \end{pmatrix}. \quad \text{(A4)}$$

where $\eta_t \sim iN(0,Q)$ and $Q = diag\left(\sigma_e^2, 0, ..., 0, \sigma_g^2, 0, ..., 0, \sigma_i^2\right)$.

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Table 1. Indicators used to construct the indexes

Country	GDP	Employ.	Unemploy.	IP	Sales
Argentina	93.II-09.I	03.I-09.I	-	86.01-09.05	97.01-09.05
Brazil	90.II-09.I	-	90.01-09.04	92.01-09.05	01.01-09.04
Chile	90.II-09.I	-	86.01-09.04	82.03-09.05	06.01-09.05
Colombia	91.II-09.I	-	01.01-09.04	81.01-09.04	90.01-09.04
Mexico	90.II-09.I	-	87.01-09.03	71.01-09.03	02.01-09.04
Peru	80.II-09.I	-	01.05-08.12	87.01-09.04	-
Venezuela	97.II-09.I	94.III-09.I	-	98.01-09.04	03.01-09.01

Notes. The source of the data are World Bank and Datastream.

Table 2. Loading factors

Country	GDP	Employ.	Unemploy.	IP	Sales	% variance
Argentina	0.55 (0.08)	0.34 (0.38)	-	0.32 (0.01)	0.04 (0.03)	49.18
Brazil	0.57 (0.06)	-	-0.02 (0.01)	0.31 (0.03)	0.10 (0.03)	67.12
Chile	0.11 (0.01)	-	-0.01 (0.01)	0.05 (0.01)	0.01 (0.01)	79.53
Colombia	1.12 (0.40)	-	-0.05 (0.04)	0.01 (0.01)	0.05 (0.01)	95.00
Mexico	1.12 (0.06)	-	-0.02 (0.02)	0.42 (0.01)	0.42 (0.04)	92.44
Peru	0.81 (0.05)	-	-0.02 (0.12)	0.36 (0.02)	-	77.42
Venezuela	0.05 (0.03)	0.02 (0.03)	-	0.04 (0.01)	0.04 (0.01)	23.29

Notes: Loading factors capture the correlation between the unobserved common factor and the variables Standard errors are in parentheses. Last row refers to the percentage of variance of GDP growth explained by the common factor.

Table 3. Cointegration tests

Countries	Prices					
	Economist	Moodys	Food	Nonfood	Metal	Specific
Engle and Granger (1987)						
Argentina	-1.60	-1.52	-1.69	-1.53	-1.56	-1.49
Brazil	-3.08	-3.59	-2.87	-2.97	-3.22	-2.93
Chile	-1.48	-1.44	-1.46	-1.74	-1.73	-1.67
Colombia	-1.39	-1.43	-1.58	-1.50	-1.39	-0.84
Mexico	-2.44	-2.57	-2.23	-2.85	-2.37	-2.06
Peru	-2.21	-2.20	-2.22	-2.33	-2.20	-2.21
Venezuela	-1.57	-1.56	-1.74	-1.56	-1.57	-1.43
Stock and Watson (1988)						
Argentina	-3.41	-0.32	-4.35	-8.30	-4.98	-1.73
Brazil	-1.95	-0.18	-2.76	-5.93	-2.60	-0.01
Chile	-2.54	0.08	-4.83	-6.70	-4.29	-1.61
Colombia	-2.65	-0.02	-6.07	-6.35	-4.46	-4.48
Mexico	-6.72	-1.85	-8.62	-7.89	-6.19	-6.47
Peru	-2.52	-0.06	-5.41	-5.38	-4.29	-1.94
Venezuela	-1.50	-0.44	-2.98	-3.96	-2.16	-1.92
Bierens (1997)						
Argentina	0.0008	0.0008	0.0001	0.0005	0.0035	0.030
Brazil	0.001	0.0001	0.0001	0.0001	0.0001	0.000
Chile	0.0016	0.0023	0.0011	0.0006	0.0024	0.003
Colombia	0.0002	0.0004	0.0001	0.0003	0.0007	0.001
Mexico	0.001	0.0011	0.0005	0.0011	0.0001	0.001
Peru	0.0001	0.0001	0.0001	0.0001	0.0001	0.001
Venezuela	0.0001	0.0001	0.0003	0.0002	0.0001	0.000

Notes: Critical values (5%) for Engle-Granger, Stock-Watson and Bieren tests are -3.42, -8.0, and 0.0169, respectively. The country-specific commodity price indexes (last column) have been obtained from Cunha, Prada and Sinnott (2010).

Table 4. Markov-switching parameter estimates

	c_0	c_1	σ_w^2	p_{00}	p_{11}	Pseudo R^2
Argentina	1.30 (0.09)	-2.17 (0.04)	1.43 (0.04)	0.97 (1.29)	0.83 (0.49)	0.58
Specific	9.05 (0.64)	-1.83 (0.06)	29.25 (0.01)	0.85 (0.32)	0.95 (0.58)	0.53
Brazil	1.44 (0.11)	-1.08 (0.09)	1.37 (0.05)	0.92 (0.57)	0.82 (0.39)	0.59
Specific	11.35 (0.96)	-1.09 (0.14)	27.86 (0.01)	0.79 (0.28)	0.95 (6.04)	0.53
Chile	0.44 (0.11)	-0.26 (0.19)	0.46 (0.03)	0.97 (2.60)	0.82 (0.84)	0.74
Specific	23.16 (1.58)	-1.51 (0.13)	88.24 (0.01)	0.74 (0.25)	0.96 (0.64)	0.51
Colombia	1.19 (0.06)	-0.21 (0.05)	0.47 (0.03)	0.94 (0.81)	0.84 (0.40)	0.54
Specific	20.00 (1.91)	-1.76 (0.12)	81.15 (0.01)	0.81 (0.33)	0.96 (0.84)	0.52
Mexico	1.15 (0.11)	-4.75 (0.01)	3.31 (0.01)	0.98 (2.07)	0.85 (0.62)	0.55
Specific	14.09 (1.35)	-2.10 (0.08)	86.86 (0.01)	0.84 (0.35)	0.95 (0.65)	0.42
Peru	1.27 (0.11)	-0.21 (0.05)	0.47 (0.03)	0.94 (0.81)	0.84 (0.40)	0.54
Specific	15.74 (1.54)	-0.90 (0.23)	47.20 (0.01)	0.78 (0.29)	0.95 (0.91)	0.40
Venezuela	0.70 (0.14)	-12.18 (0.05)	3.37 (0.01)	0.99 (6.63)	0.85 (1.26)	0.63
Specific	56.65 (8.86)	0.98 (0.39)	208.73 (0.01)	0.71 (0.52)	0.90 (3.72)	0.32
Economist	6.39 (0.43)	-2.27 (0.03)	21.44 (0.01)	0.93 (0.61)	0.95 (0.72)	0.56
Moodys	5.71 (0.49)	-1.86 (0.08)	16.60 (0.01)	0.91 (0.44)	0.92 (0.57)	0.53
Food	8.11 (0.55)	-2.79 (0.03)	27.22 (0.01)	0.91 (0.47)	0.94 (0.57)	0.60
Nonfood	6.91 (0.67)	-2.29 (0.06)	40.73 (0.01)	0.93 (0.68)	0.95 (0.82)	0.41
Metal	12.01 (1.36)	-2.83 (0.04)	59.30 (0.01)	0.87 (0.42)	0.95 (0.86)	0.52

Notes: The figure reports the parameter estimates from the model $w_t = c_{s_t} + \varepsilon_{wt}$, where $\varepsilon_{wt} \sim iidN(0,\sigma_w^2)$, and $p(s_t = j/s_{t-1} = i) = p_{ij}$. Entries labelled as *specific* refer to the country-specific index of prices obtained from Cunha, Prada and Sinnott (2010).

Table 5. Markov-switching tests

Time series	Hansen (1992)		Carrasco et al. (2004)
	p=0	p=1	p=1
Argentina	0.000	0.024	0.060
Specific	0.000	0.000	0.000
Brazil	0.000	0.000	0.000
Specific	0.000	0.000	0.000
Chile	0.000	0.000	0.000
Specific	0.000	0.000	0.000
Colombia	0.000	0.000	0.000
Specific	0.000	0.000	0.000
Mexico	0.000	0.016	0.040
Specific	0.000	0.000	0.000
Peru	0.000	0.018	0.170
Specific	0.000	0.000	0.000
Venezuela	0.000	0.000	0.010
Specific	0.000	0.000	0.000
Economist	0.000	0.000	0.590
Moodys	0.000	0.000	0.000
Food	0.000	0.000	0.000
Nonfood	0.000	0.000	0.822
Metal	0.000	0.000	0.000

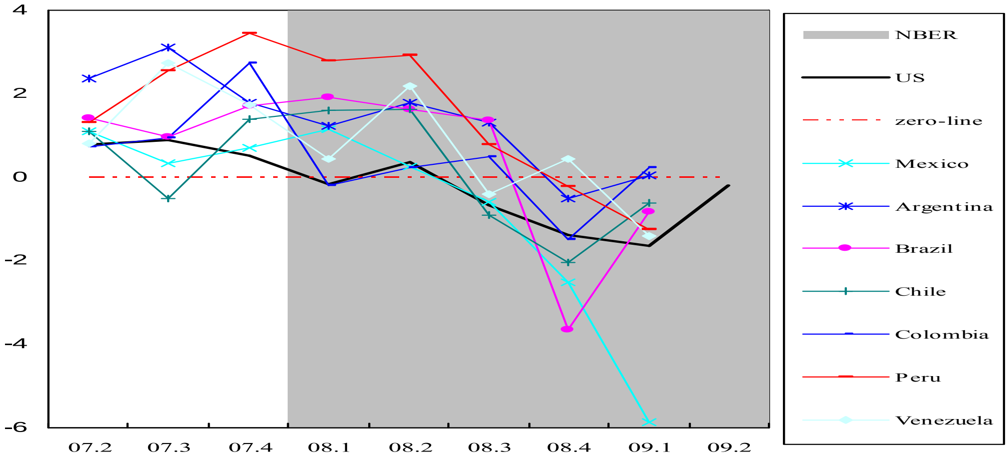
Notes: Entries are p-values of the null of linearity against Markov-switching. Entries labelled as specific refer to the country-specific index of prices obtained from Cunha, Prada and Sinnott (2010).

Table 6. Test of nonlinear responses

Countries	Prices					
	Economist	Moodys	Food	Nonfood	Metal	Specific
Argentina	0.072	0.005	0.103	0.065	0.279	0.264
Brazil	0.082	0.015	0.001	0.017	0.001	0.217
Chile	0.001	0.001	0.028	0.095	0.009	0.036
Colombia	0.021	0.069	0.349	0.104	0.528	0.000
Mexico	0.010	0.001	0.185	0.001	0.001	0.057
Peru	0.018	0.071	0.061	0.160	0.191	0.028
Venezuela	0.057	0.014	0.001	0.069	0.034	0.001

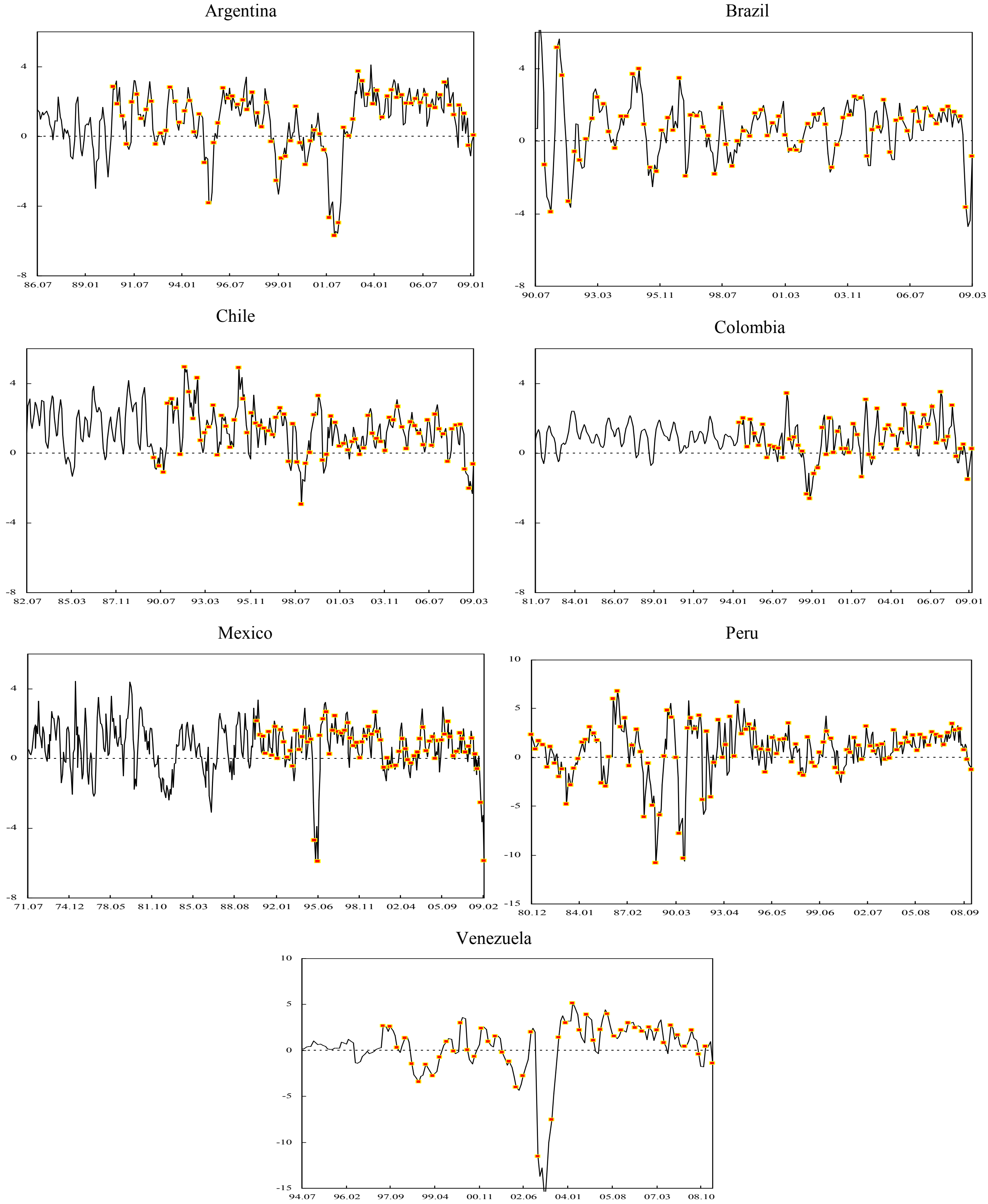
Notes: Following Hamilton (2001), entries are p-values of the null that the reaction of outputs to prices is linear. The alternative assumes that this relation is nonlinear with a flexible functional form. The country-specific commodity price indexes (last column) have been obtained from Cunha, Prada and Sinnott (2010).

Figure 1. Recent economic developments in LAC



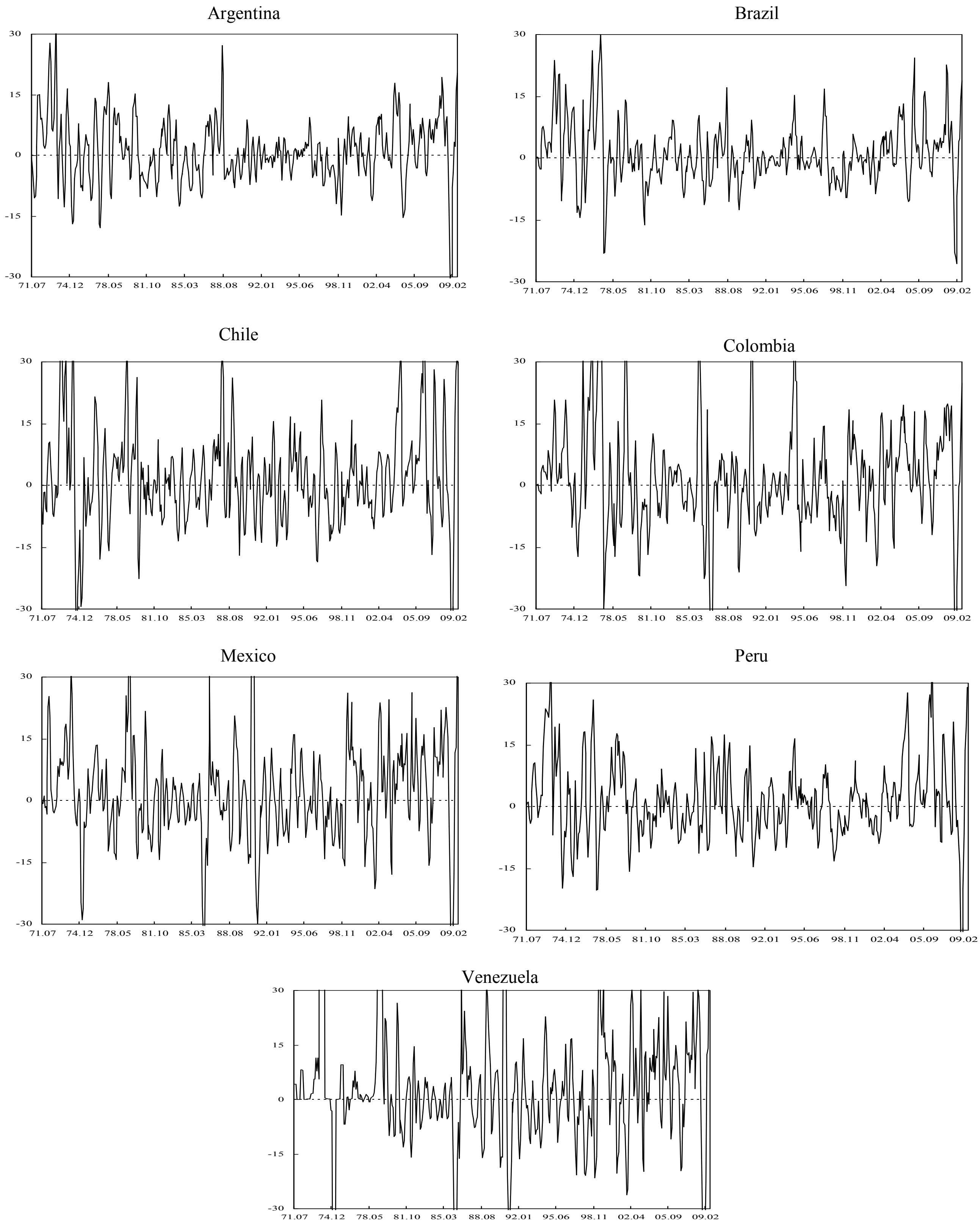
Notes. The figure plots quarterly GDP growth rates. Shaded area refers to the 2008 NBER recession (the through has not been dated yet).

Figure 2: Quarterly GDP growth rates: Data and interpolation.



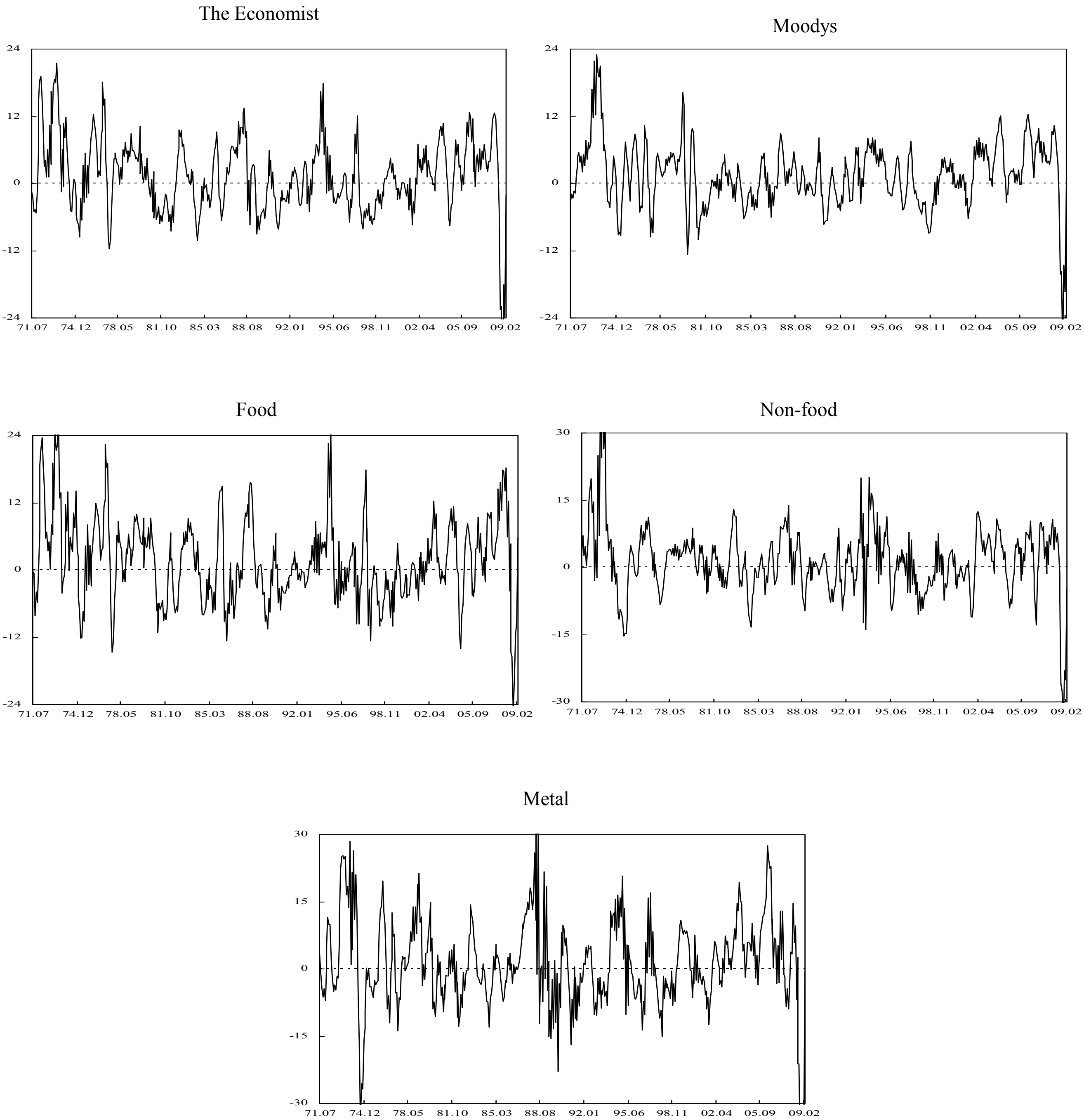
Notes. The charts plot quarterly growth rates of GDP which have been interpolated by using monthly indicators with dynamic factor models. Plot marks refer to actual quarterly growth rates.

Figure 3. Quarterly growth rates of country-specific commodity price indexes.



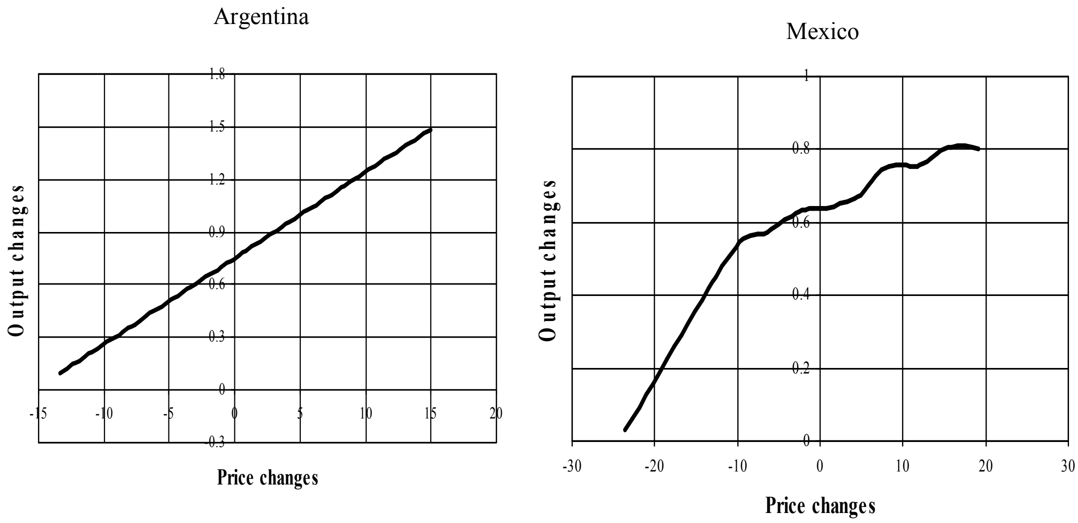
Notes. The country-specific price indexes have been obtained from Cunha, Prada and Sinnott (2010).

Figure 4. Quarterly growth rates of composite commodity price indexes.



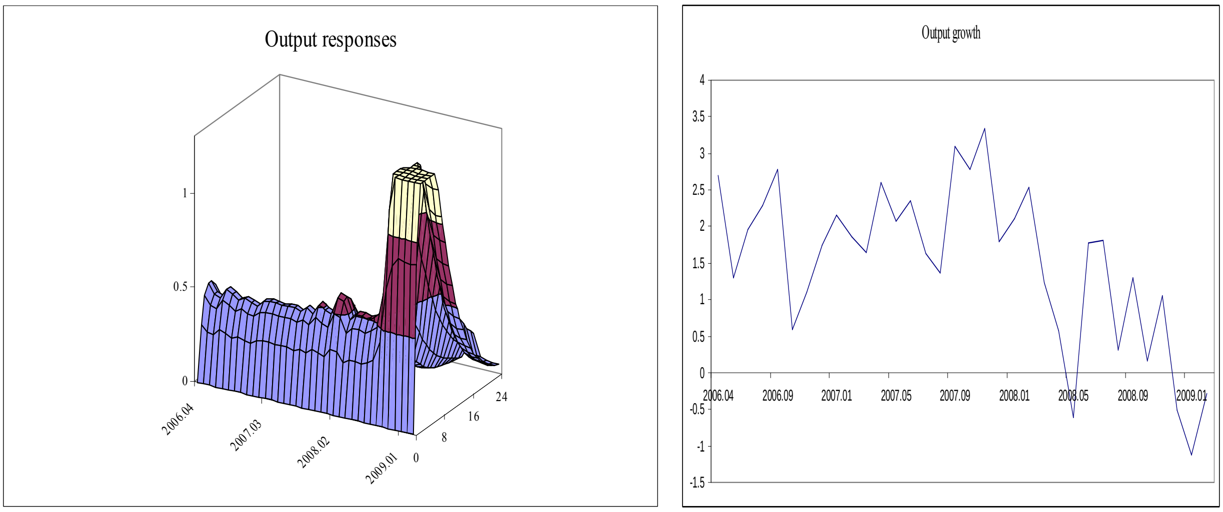
Notes. The composite price indexes have been obtained from Moodys, and The Economists (including its disaggregation in Food, Non-food and Metals).

Figure 5. Flexible functional forms



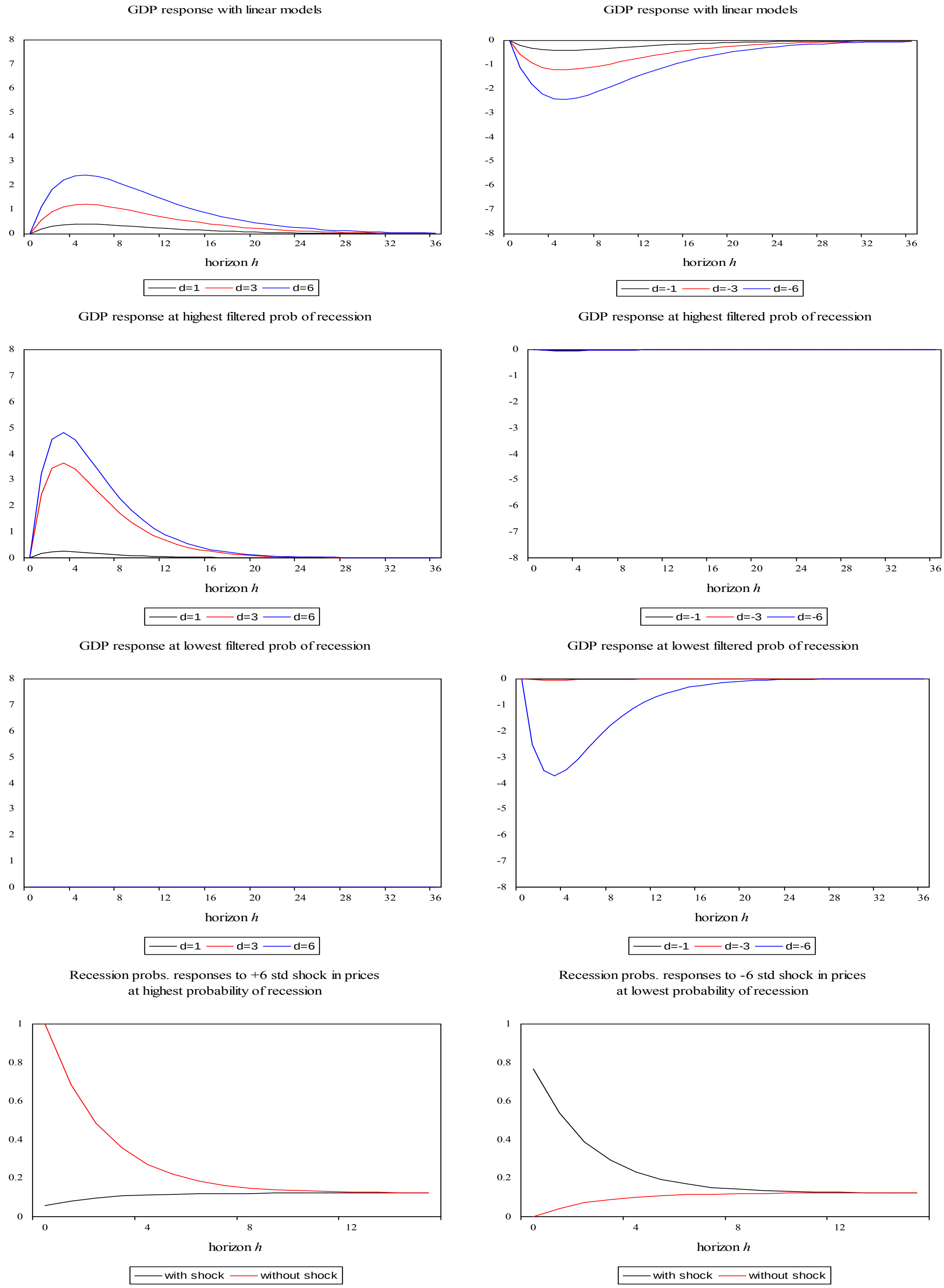
Notes. The charts plot the predicted values of GDP growth against different values of country-specific prices growths from the flexible functional form model advocated by Hamilton (2001).

Figure 6. Evolution of linear responses and output growth in Argentina



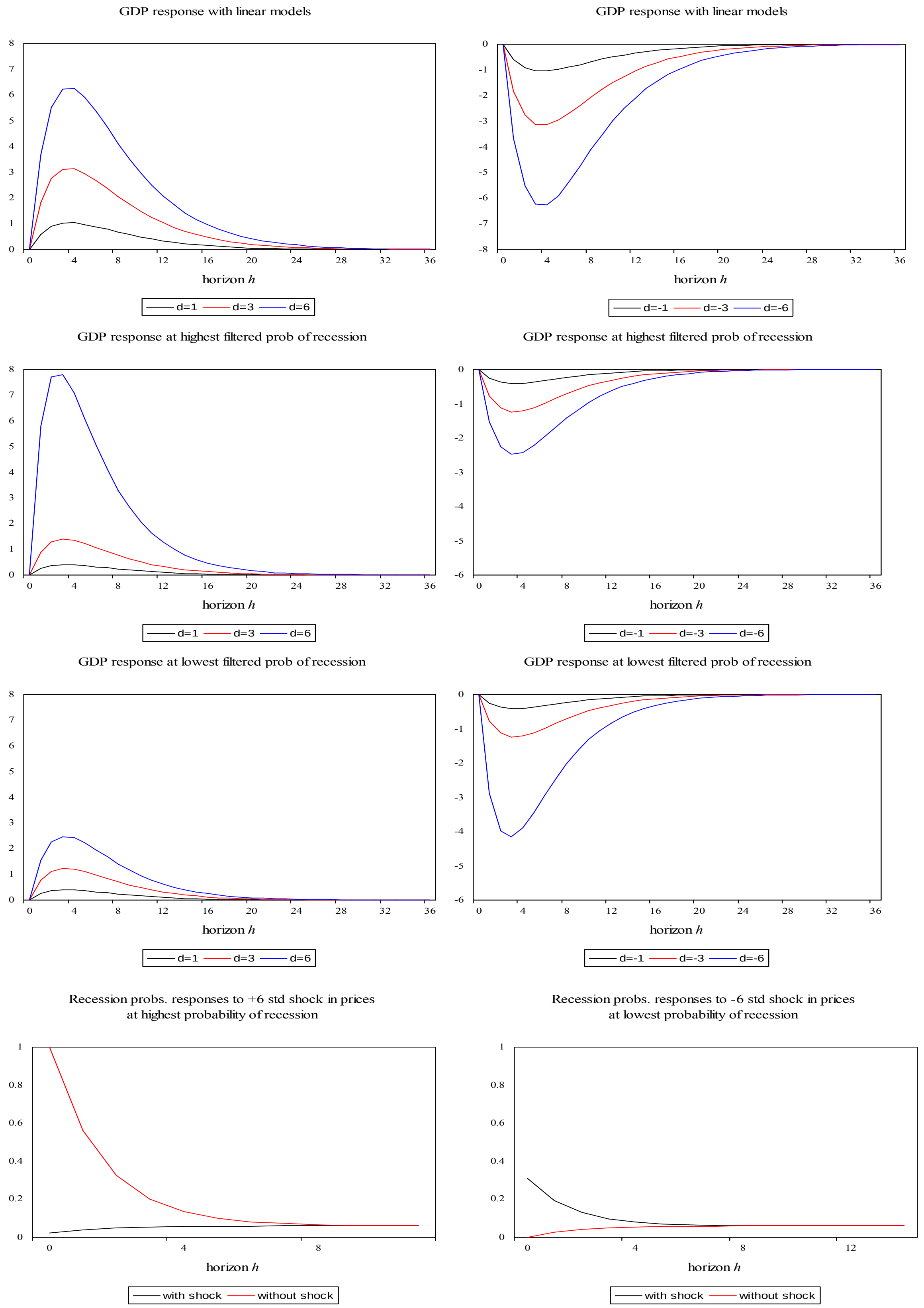
Notes. The left-hand-side chart plots the 24-month (X axis) linear responses of Argentinean GDP growth to one standard deviation shock in its country-specific commodity price shock . They are calculated from 2006.04 to 2009.03 (Y axis) using a rolling window of four years. The right-hand-side graph plots the quarterly growth rate of GDP at monthly frequency.

Figure 7. IRF Argentina



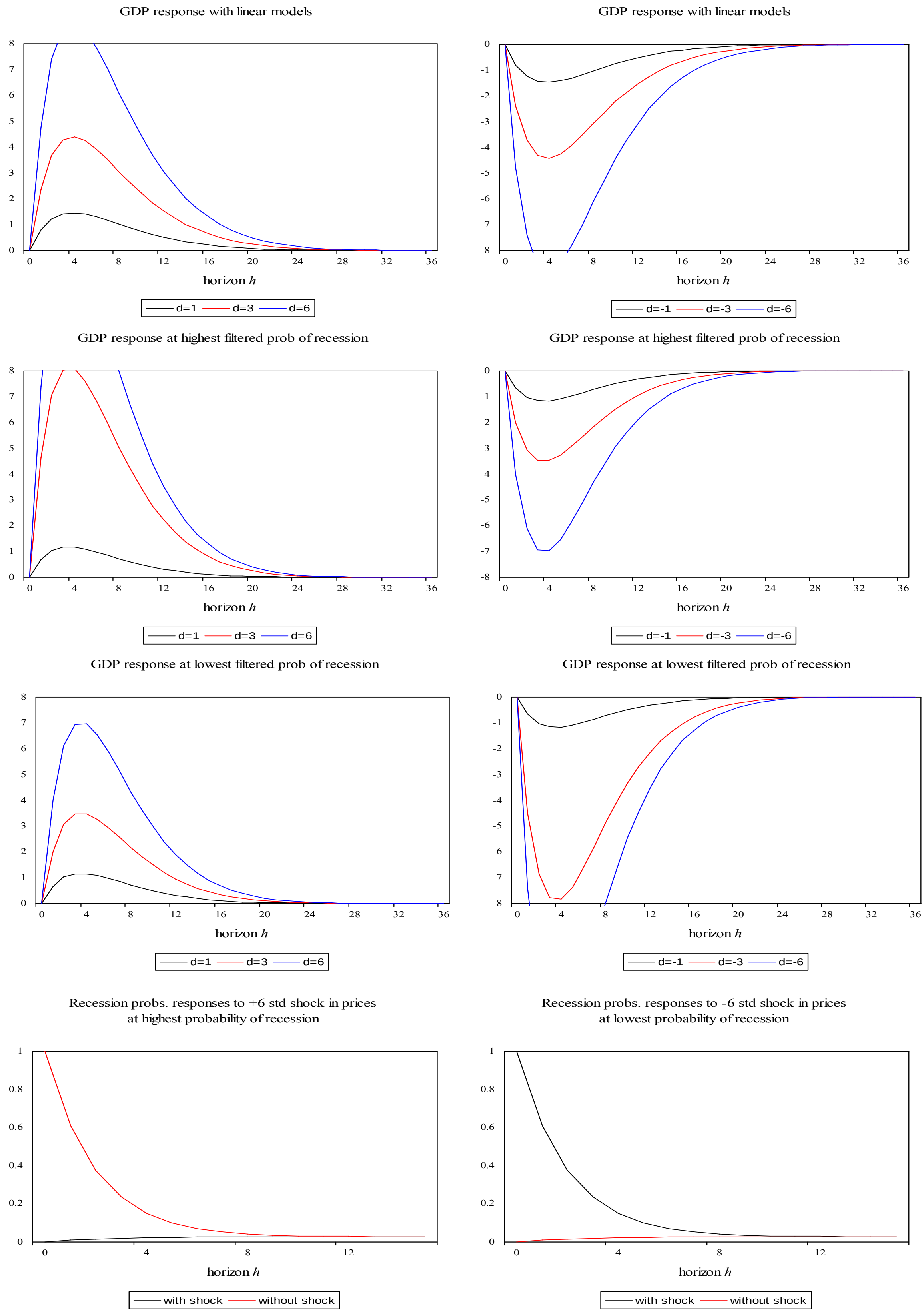
Notes. Reactions to positive (negative) commodity price shocks of different sizes *d* are on the left (right) hand graphs. The first row graphs are linear responses of output growth. The next graphs show the Markov-switching responses of output (second and third row graphs) and recession probabilities (last row of graphs) to price shocks that occur at the highest and lowest probabilities of recession.

Figure 8. IRF Brazil



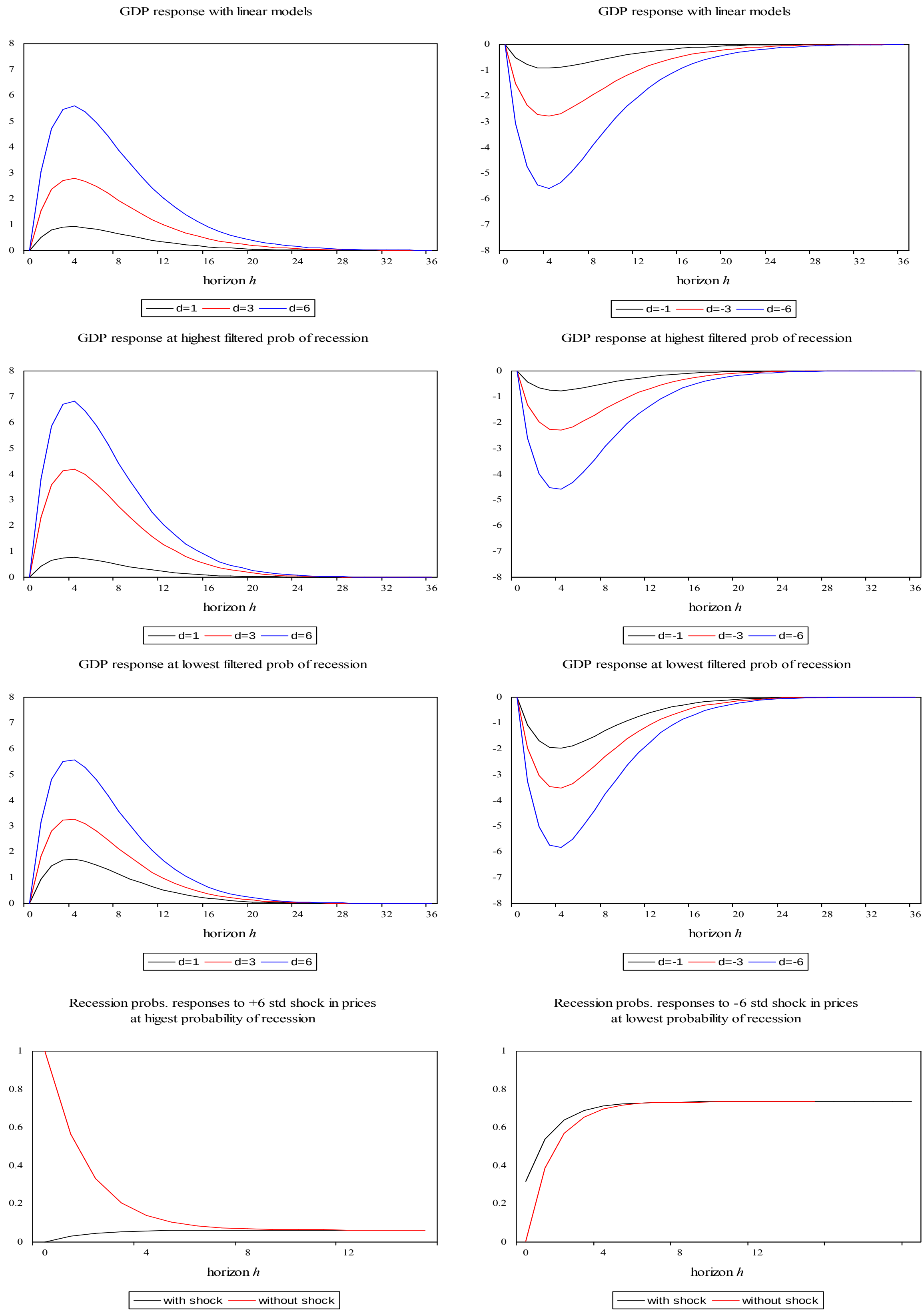
Notes. See notes of Figure 7

Figure 9. IRF Chile



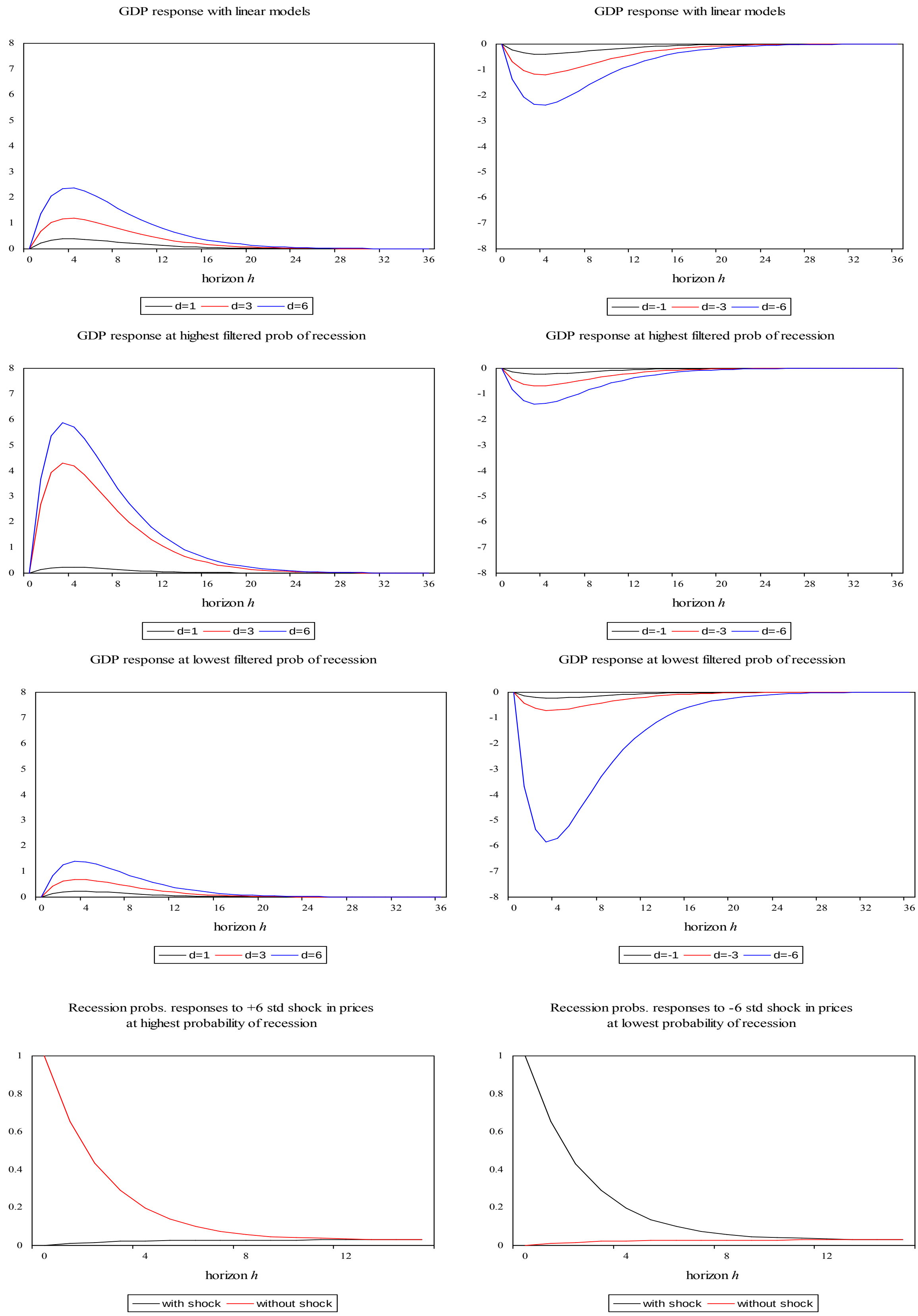
Notes. See notes of Figure 7

Figure 10. IRF Colombia



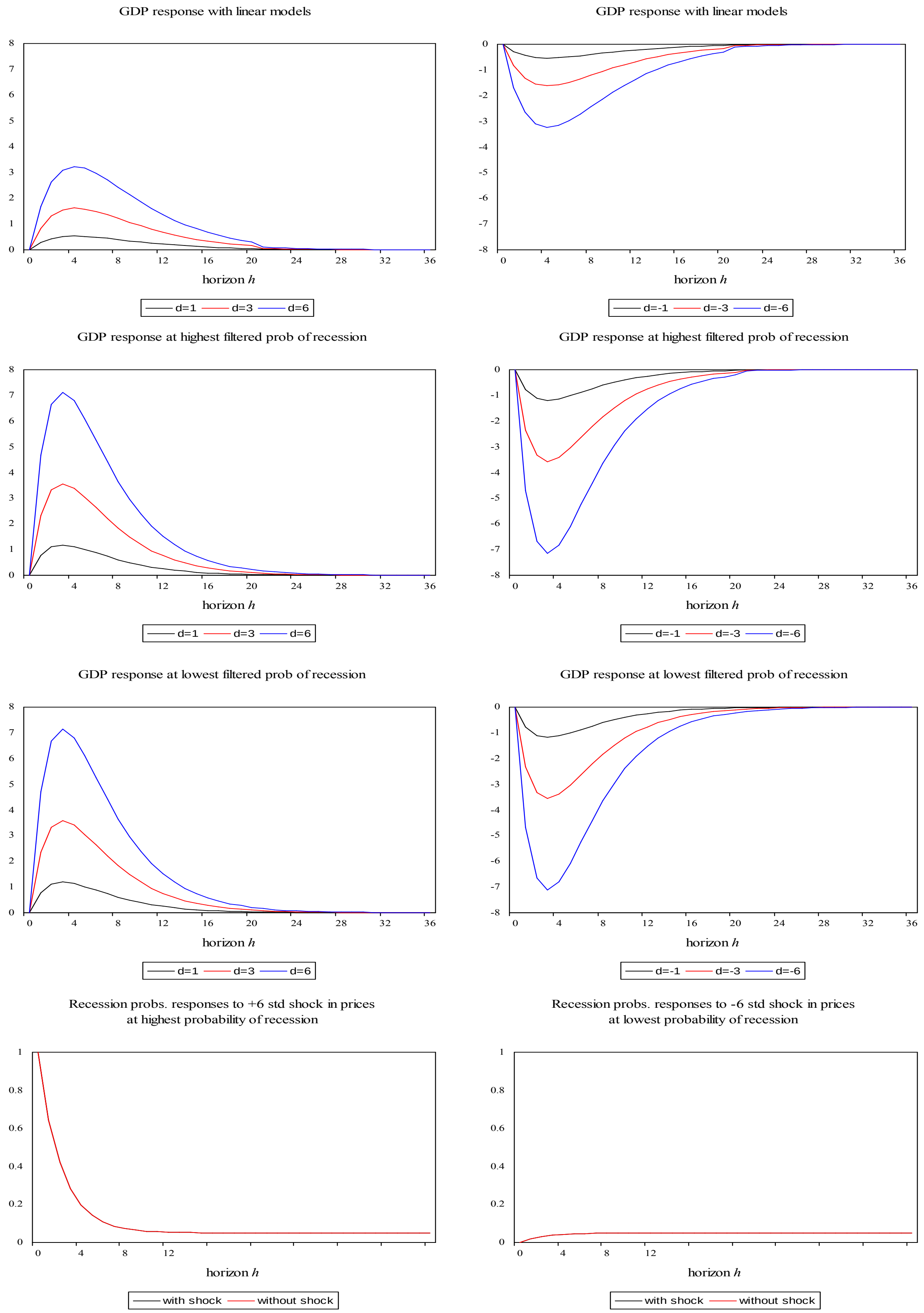
Notes. See notes of Figure 7

Figure 11. IRF Mexico



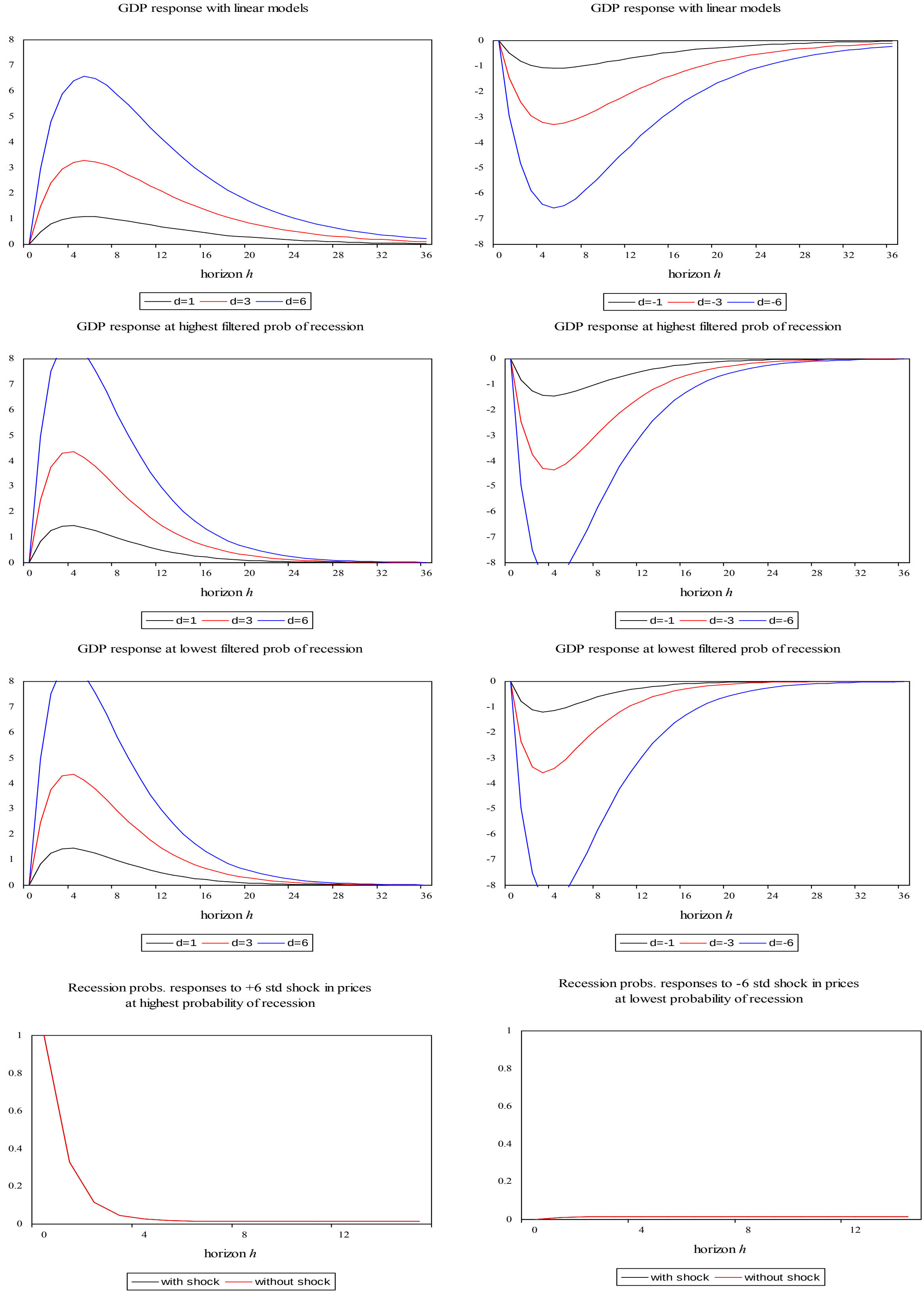
Notes. See notes of Figure 7

Figure 12. IRF Peru



Notes. See notes of Figure 7

Figure 13. IRF Venezuela



Notes. See notes of Figure 7