Output Growth and Structural Reform in Latin America: Have Business Cycles Changed?

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Abstract

This paper documents important changes in real GDP growth of six large Latin American countries. The main results can be summarized as follows. First, there is evidence of a structural break in real GDP towards stronger mean growth and a substantial reduction in volatility. Second, the timing of the breaks suggests that the important changes in economic policies of the 1980s and 1990s have been effective in permanently improving economic growth in the region. These changes in the growth processes imply recessions that are shorter in duration and milder in amplitude. The sustained increase in commodity prices observed in recent years explains an important share of growth in the region since 2003. But after accounting for the effect of commodity prices, there is even stronger evidence of a structural break in real GDP growth.

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1 Introduction

In recent years, Latin American (LatAm) countries have experienced an unprecedented period of booming asset prices and investment, appreciating real exchange rates, and strong output growth (Izquierdo et al., 2008; Sosa et al., 2013). The improvement in output growth is evident in the statistics for quarterly real GDP summarized in Table 1 for six large LatAm countries (Argentina, Brazil, Chile, Colombia, Mexico, and Peru). For example, during the 1980s the region's average quarterly real GDP growth rate was 0.43% (almost 2% in annual terms).[1] The region's average growth rate rose to 0.70% in the 1990s, and was 1.20% (almost 5% in annual terms) in the 2000s. In addition, we observe a substantial reduction in the volatility of the growth rate of GDP. During the 1980s the region's average standard deviation of quarterly real GDP growth was 2.85%. There is a reduction to 1.70% in the 1990s, and in the 2000s the standard deviation of the quarterly growth rate was only 1.27%. In sum, over the last three decades, LatAm economies have shown a trend towards stronger mean growth and reduced volatility in real GDP.

[TABLE 1 ABOUT HERE]

The empirical literature on LatAm business cycles suggests several explanations for (some of) the changes observed in output growth. For example, Easterly et al. (1997) argue that growth in the 1990s was stronger than in previous decades due to the changes in economic policies (structural reforms) implemented in LatAm mainly in the 1980s (Lora, 1997, 2012). In contrast, recent research has focused on the role of external conditions. For example, Calderón and Fuentes (2014) argue that the decline in the

[1] LatAm values are a simple average of the corresponding values (mean growth rates, standard deviations) of the six countries considered.

amplitude of recessions observed in LatAm during what they call the 'globalization era' (after 1990) can be partially attributed to structural changes in advanced economies (the Great Moderation). In addition, Österholm and Zettelmeyer (2007), Izquierdo et al. (2008), and Camacho and Perez-Quiros (2013) show that favorable external conditions such as abundant international liquidity and a rise in commodity prices can explain a significant share of recent LatAm growth. As summarized in Table 1 (bottom panel), the average quarterly growth rate of relevant commodity price indexes was negative (-0.31%) in the 1980s and approximately 0% in the 1990s. But after 2003, the average quarterly growth rate of commodity prices rose to 3.75% (approximately 15% in annual terms). That is, the relevant commodity price indexes exhibit a trend similar to the one observed in the mean growth rate and standard deviation of real GDP in LatAm.[2]

Based on these results, in this paper I ask whether there has been a structural break in LatAm towards stronger growth and more stability. To investigate the nature of the potential structural break in the real GDP growth processes, autoregressive (AR) models with Markov-switching parameters (Kim and Nelson, 1999a; Kim et al., 2004) that allow for permanent regime shifts are fitted to the quarterly series of the six LatAm countries. Next I ask: If there has been a structural break, has the timing been similar across countries? Finally, I ask: How much of the recent improvement in LatAm's real GDP growth can be attributed to the boom in commodity prices observed in the last decade? To answer this last question, I incorporate the growth rate of commodity prices to the Markov-switching AR models allowing for linear and nonlinear effects on real GDP and compute the average contribution of changes in commodity prices to growth. The approach is similar to the one used in Hamilton (2003, 2011) to model

[2]Country-specific commodity price indexes are defined as the world price of a country's commodity exports and obtained from Chen and Lee (2013) for the period is 1980Q1–2010Q4.

the potentially nonlinear relationship between U.S. real GDP growth and changes in oil prices.

Estimation results suggest important changes in the real GDP processes of the six LatAm countries. First, there is strong evidence of a structural break in real GDP with break dates clustered between the early 1990s and late 1990s. Although there are differences between countries, the break is towards stronger mean growth and a substantial reduction in volatility. Second, the timing of the breaks (around the time of the structural reforms) suggests that the important changes in economic policies of the 1980s and 1990s have been effective in permanently improving economic growth in LatAm. But while the literature has documented mainly an improvement in the mean growth rate of real GDP (e.g., Easterly et al., 1997), the results presented here show that the structural reforms also led to important reductions in the volatility of real GDP growth (49% for Argentina, 51% for Brazil, and 67% for Peru). Overall, these changes in the real GDP growth processes have important implications for the characteristics of the business cycle phases in LatAm. As explained in Blanchard and Simon (2001) and Harding and Pagan (2002), these results imply that in the post-break sample recessions are shorter in duration and milder in amplitude. This result is consistent with recent findings of Gonçalves and Salles (2008), Aiolfi et al. (2011), and Calderón and Fuentes (2014) who show that the amplitude of recessions has declined in LatAm after 1990.

Finally, I find strong evidence of a positive and linear relationship between the growth rate of real GDP and the growth rate of commodity prices in LatAm. But contrary to Camacho and Perez-Quiros (2013), the hypothesis that commodity price increases have different economic effects from commodity price decreases (i.e., the relationship is nonlinear) is strongly rejected for all countries considered. On average, the effect of commodity prices on real GDP growth in the 1980s and 1990s was small and sometimes negative. In contrast, the effect after 2003 was positive and much larger in magnitude (up to 2% of annual real GDP growth for Peru). As a result, the sustained increase in commodity prices observed in recent years explains an important share of LatAm growth since 2003. But after accounting for the effect of commodity prices, there is even stronger evidence of a structural break in real GDP growth towards an increase in mean growth and a reduction in volatility.

This paper is organized as follows. Section 2 documents important changes in the real GDP growth processes of the six LatAm countries considered, as well as changes in the evolution of country-specific commodity price indexes. Section 3 presents model specifications used to investigate the nature of the potential structural break in real GDP growth. Section 4 reports estimation results for each country, including a discussion of the timing of the breaks and an estimation of the contribution of commodity prices to real GDP growth. Section 5 concludes.

2 Recent Changes in Output Growth

In this section I document important changes in the output growth processes of six large LatAm countries. Data employed is quarterly real GDP (seasonally adjusted) for Argentina, Brazil, Chile, Colombia, Mexico, and Peru obtained from Cesa-Bianchi et al. (2012) and Rondeau (2012). Quarterly growth rates (Δy_t) are computed as $100 \times \Delta \ln Y_t$, where Y_t is quarterly real GDP for a given country. The sample period is 1983Q1–2010Q4 for all countries except Brazil and Colombia. In the case of Brazil the sample period is 1990Q1–2010Q4, while in the case of Colombia the sample period is 1994Q1–2010Q4. Figure 1 shows the rolling average of quarterly real GDP growth using an 8-year window (solid line) for the six countries. The reported value for quarter t is the average growth rate over quarters $t-31$ to t. Figure 1 also shows the estimated linear trends (dashed line) fitted to the rolling averages and the regression R^2. Average growth rates have increased over the sample period, i.e. the trend coefficient is positive and statistically significant at the 5% level, for all countries except Chile and Mexico. While in the case of Chile the coefficient is negative and statistically significant at the 5% level, Mexico does not show significant changes in average growth.

[FIGURE 1 ABOUT HERE]

Figure 2 shows the rolling standard deviation of quarterly real GDP growth using an 8-year window (solid line) for the six countries. The values are computed in the same way as the rolling averages, i.e. the reported value for quarter t is the standard deviation of the growth rate over quarters $t-31$ to t. Figure 2 also shows the estimated linear trends (dashed line) fitted to the rolling standard deviations and the regression R^2. Over the sample period, the volatility of real GDP growth has substantially declined in LatAm. The regressions yield a coefficient on the trend term that is negative and statistically significant at the 5% level for all countries.

[FIGURE 2 ABOUT HERE]

The next question is whether these changes in the unconditional moments (mean and variance) of real GDP growth arise from changes in the conditional mean (i.e., changes in the autoregressive coefficients), changes in the conditional variance (i.e., changes in the innovation variance), or both. To answer this question, I test for parameter instability using an autoregressive (AR) model for real GDP growth given by

$$\Delta y_t = c + \sum_{j=1}^{k} \phi_j \Delta y_{t-j} + \varepsilon_t, \quad \varepsilon_t \sim \text{iid N}(0, \sigma_\varepsilon^2), \tag{1}$$

where Δy_t is quarterly real GDP growth and $k = 1$. For each country, the models are estimated by ordinary least squares (OLS), using all the available data. Table 2 (top panel) reports parameter estimates and the qLL test of parameter instability of Elliott and Muller (2006). The test statistic for parameter instability in the conditional mean (i.e., c and ϕ) is qLL_1 and the null hypothesis of joint stability is rejected for small values of the statistic. The results show that we can reject the hypothesis of stability in the conditional mean at the 10% level for Argentina, Mexico, and Peru. To test for instability in the conditional variance, the qLL test is computed for the regression

$$\sqrt{\pi/2} \times |\hat{\varepsilon}_t| = s + \eta_t, \tag{2}$$

where $|\hat{\varepsilon}_t|$ is the absolute value of the OLS residuals in (1) and s is a constant (a similar approach to McConnell and Perez-Quiros, 2000). The test statistic for parameter instability in s is qLL_2 and the results reported in Table 2 (bottom panel) show that we can reject the hypothesis of stability in the conditional variance at the 10% level for all countries except Brazil and Colombia. In sum, based on univariate AR models, there is strong evidence of instability in the conditional moments of real GDP growth.

[TABLE 2 ABOUT HERE]

3 Model Specification

Based on the results in the previous section, a model for real GDP growth should allow for changes in the conditional mean (the coefficients in the AR model) as well as changes in the conditional variance (the innovation variance) of the time series process. Time variation in the parameters can be incorporated in different ways. The most common approach consists in estimating the AR model allowing for one or more permanent structural breaks in the model parameters. For example, Kim and Nelson (1999a), McConnell and Perez-Quiros (2000), Stock and Watson (2002), and Kim et al. (2004) use this approach to analyze the reduction in volatility observed in the U.S. economy in the 1980s known as the Great Moderation.[3] In addition, there is evidence of a positive link between real GDP growth and commodity prices for LatAm countries (Camacho and Perez-Quiros, 2013).

As a result, I consider an AR(k) model with Markov-switching parameters and commodity prices given by

$$\phi_{S_t}(L)\big(\Delta y_t - \mu_{S_t} - \beta(L)\Delta p_t - \gamma(L)\Delta p_t^+\big) = \varepsilon_t, \quad \varepsilon_t \sim \text{iid N}(0, \sigma_{S_t}^2), \tag{3}$$

where Δy_t is quarterly real GDP growth, Δp_t is the quarterly growth rate of a relevant commodity price index, $\Delta p_t^+ = \max\{0, \Delta p_t\}$, $\phi_{S_t}(L) = 1 - \phi_{S_t} L$ with $|\phi_{S_t}| < 1$, $\beta(L) = \beta_1 L + \beta_2 L^2$, $\gamma(L) = \gamma_1 L + \gamma_2 L^2$, and S_t is an unobserved two-state first-order Markov process with transition probabilities given by

$$\text{Prob}(S_t = 1 \,|\, S_{t-1} = 1) \;=\; p, \tag{4}$$

$$\text{Prob}(S_t = 2 \,|\, S_{t-1} = 2) \;=\; 1, \tag{5}$$

with $0 < p < 1$. This model allows for a one-time permanent structural break in the autoregressive parameters μ, ϕ, and σ (i.e., the second regime is absorbing). Therefore, μ_1, ϕ_1, and σ_1 are the parameters of the AR(1) model in the pre-break sample (regime 1) and μ_2, ϕ_2, and σ_2 are the parameters of the model in the post-break sample (regime

[3]An alternative approach considered in Blanchard and Simon (2001) consists in estimating AR models with time-varying parameters, i.e. models that allow the parameters to change every quarter.

2). A key component of this model is the term $\beta(L)\Delta p_t$ which captures the contribution of commodity prices to real GDP growth (measured as the optimal one-quarter-ahead forecast). In addition, the term $\gamma(L)\Delta p_t^+$ allows commodity price increases to have different economic effects from commodity price decreases.[4] These terms imply $\mu^*_{S_t t} = \mu_{S_t} + \beta(L)\Delta p_t + \gamma(L)\Delta p_t^+$ and, as a result, an alternative interpretation of (3) is $\phi_{S_t}(L)\big(\Delta y_t - \mu^*_{S_t t}\big) = \varepsilon_t$, i.e. a model with a time-varying and regime-specific mean growth rate. For example, Hamilton (2003, 2011) uses a similar approach to model the potentially nonlinear relationship between U.S. real GDP growth and changes in oil prices. Finally, the unknown break date (τ) is treated as a parameter to be estimated as the expected duration of regime 1, i.e. $E(\tau) = 1/(1-p)$. The model can be estimated by maximum likelihood (ML) following Kim and Nelson (1999b) and Kang et al. (2009).

I consider four competing models based on the following restrictions:

1. Model I: A standard AR(1) without structural breaks or commodity prices.

   Restrictions: $\mu_1 = \mu_2$, $\phi_1 = \phi_2$, $\sigma_1 = \sigma_2$, and $\beta_1 = \beta_2 = \gamma_1 = \gamma_2 = 0$.

2. Model II: A MS-AR(1) with one break but no commodity prices.

   Restrictions: $\beta_1 = \beta_2 = \gamma_1 = \gamma_2 = 0$.

3. Model III: A MS-ARX(1) with one break and linear commodity prices.

   Restrictions: $\gamma_1 = \gamma_2 = 0$.

4. Model IV: A MS-ARX(1) with one break and nonlinear commodity prices.

   No restrictions.

[4] Consistent estimation of (3) with contemporaneous commodity prices requires the assumption that ε_t is uncorrelated with Δp_t and Δp_t^+. I avoid this potential endogeneity problem by including only predetermined variables $(\Delta p_{t-1}, \Delta p_{t-2}, \Delta p^+_{t-1}, \Delta p^+_{t-2})$. See, for example, Hamilton (2011).

Model selection is based on the Akaike Information Criterion (AIC) calculated as $-\ln L + k$ where $\ln L$ denotes the log likelihood and k is the number of parameters in the model. In addition, I investigate the nature of the potential structural break in real GDP growth by testing the following two null hypotheses: (i) No break in the conditional mean (H_0: $\mu_1 = \mu_2$ and $\phi_1 = \phi_2$); (ii) No break in the conditional variance (H_0: $\sigma_1 = \sigma_2$). Finally, I investigate the relationship between real GDP growth and the growth rate of commodity prices by testing the following two null hypotheses: (i) No relationship (H_0: $\beta_1 = \beta_2 = 0$ and $\gamma_1 = \gamma_2 = 0$); (ii) A linear relationship (H_0: $\gamma_1 = \gamma_2 = 0$). These hypotheses are tested using standard likelihood ratio (LR) tests.

4 Empirical Results

In this section, I present empirical results for the four competing models. As before, the sample period is 1983Q1–2010Q4 for all countries except Brazil and Colombia. In the case of Brazil the sample period is 1990Q1–2010Q4, while in the case of Colombia the sample period is 1994Q1–2010Q4. Country-specific commodity price indexes are from Chen and Lee (2013).[5] Sections 4.1 and 4.2 report model selection and estimation results for each country. The results are discussed in more detail in sections 4.3 and 4.4. Finally, section 4.5 reports some robustness results.

4.1 Model Selection

Table 3 reports the value of the $\ln L$, number of parameters in the model, and AIC for each of the four models considered. For all countries except Argentina, AIC selects a model with a structural break and commodity prices. But contrary to Camacho and

[5]Other LatAm countries could not be considered due to data availability issues. See Cesa-Bianchi et al. (2012), Rondeau (2012), and Chen and Lee (2013) for a detailed description of their data sources.

Perez-Quiros (2013), models that allow for commodity price increases to have different economic effects from commodity price decreases (i.e., a nonlinear relationship) are consistently rejected. Only in the case of Argentina AIC selects a model with a break but no commodity prices—a surprising result. Overall, based on the value of the $\ln L$, the improvements in fit relative to the linear AR(1) models with no commodity prices (Model I) can be substantial.

Other model specifications were considered (results not reported). For example, AR models with commodity prices but no breaks, ARX(1), were systematically rejected in favor of models with at least one structural break. Models allowing for two permanent structural breaks, with and without commodity prices, were also rejected in favor of models with one break. Only Argentina shows evidence of a second structural break taking place around 2002Q2. Finally, I considered models with different ARMA processes including a $\phi_{S_t}(L)$ polynomial of order 2 and a $\beta(L)$ polynomial of order 4. I found that the specification reported above was sufficient to capture the dynamics of real GDP growth in all countries.

[TABLE 3 ABOUT HERE]

4.2 Estimates

ML estimates of the MS-AR(1) models (Model II), i.e. models without commodity prices, are reported in Table 4. Ljung-Box Q-statistics computed for the (centered) generalized residuals (Kim et al., 2004; Kang et al., 2009) support that the selected specification was sufficient to capture the dynamics of real GDP growth in all countries as there is no evidence of remaining serial correlation. The first noticeable result is that average quarterly growth rates are larger in regime 2 for all countries except Chile (consistent with the results reported in Figure 1). The changes in mean growth range from modest (Mexico) to very large (Argentina, Brazil, and Peru). For example, in the case of Peru, pre-break quarterly mean growth is -0.39% while post-break mean growth is 1.33%. The case of Chile is different as quarterly mean growth in the post-break period exhibits a 34% reduction relative to the pre-break average. Similarly, the autoregressive coefficients also exhibit important changes. In this case, however, no clear pattern emerges as some countries exhibit an increase in persistence while others exhibit a reduction. Likelihood ratio tests can be used to determine whether these changes in the conditional mean are statistically significant. The null hypothesis of no break in the conditional mean is $\mu_1 = \mu_2$ and $\phi_1 = \phi_2$, and the test statistic is LR_1. Without commodity prices, we reject the null hypothesis for Argentina, Colombia, and Mexico. We also observe an important reduction in the conditional variance for all countries except Mexico (consistent with the results reported in Figure 2). The post-break estimates of the standard deviation are generally smaller than the pre-break estimates and the reductions range from almost 20% for Colombia to over 67% for Peru. The null hypothesis of no break in the conditional variance is $\sigma_1 = \sigma_2$ and the test statistic is LR_2. Without commodity prices, we reject the null hypothesis for all countries except Colombia and Mexico.[6] Point estimates of the break dates are obtained from the expected duration (in quarters) of regime 1 and computed as $\hat{\tau} = 1/(1-\hat{p})$. Based on the MS-AR(1) models, the estimated break dates are: 1991Q3 for Argentina, 1992Q1 for Brazil, 1999Q2 for Chile, 2000Q2 for Colombia, 1993Q2 for Mexico, and 1993Q1 for Peru.

[TABLE 4 ABOUT HERE]

[6] Mexico exhibits a very deep recession in the middle of the sample which makes the identification of the (potential) structural break date difficult (see Figure 3 below).

Table 5 reports ML estimates of the linear MS-ARX(1) models (Model III), i.e. models with a structural break and linear commodity prices. This is the preferred model for all countries except Argentina. As before, Ljung-Box Q-statistics show no evidence of remaining serial correlation. The coefficients on the two lags of commodity prices (β_1 and β_2) are generally positive and significant. A likelihood ratio test (LR_3) can be used to test the null hypothesis that the commodity price coefficients are zero ($\beta_1 = \beta_2 = 0$). Consistent with the model selection results, we reject the null hypothesis of no commodity effects for all countries except Argentina. On the other hand, there is no evidence of commodity price increases having different economic effects from commodity price decreases (nonlinearity).[7] In addition, with the inclusion of commodity prices, the shift in mean growth is typically smaller but more accurately estimated. As a result, the null hypothesis of no break in the conditional mean (LR_1) is now rejected more often. In this case, the hypothesis is rejected for all countries except Brazil and Mexico. Similarly, the null hypothesis of no break in the conditional variance (LR_2) is now rejected for all countries except Colombia. Therefore, with commodity prices in the model, there is stronger evidence of a structural break in real GDP growth towards an increase in mean growth and a reduction in volatility.

[TABLE 5 ABOUT HERE]

Figure 3 plots quarterly real GDP growth rates for each country and the smoothed probabilities of structural break from Model III. The probabilities are computed using Kim's smoothing algorithm (see Kim and Nelson, 1999b). The break dates appear to be clustered in the early 1990s (Argentina, Brazil, and Peru) and late 1990s (Chile, Colombia, and Mexico). Break date densities are computed by differencing

[7] LR tests of the null hypothesis that $\gamma_1 = \gamma_2 = 0$ are based on Model IV (results not reported).

the smoothed probabilities and plotted in Figure 4. All countries except Brazil and Mexico exhibit very concentrated densities for the break date and the timing of the shift appears to be well identified using models with and without commodities. In the case of Brazil and Mexico there is more uncertainty about the timing of the break. Overall, these results are consistent with the findings of Calderón and Fuentes (2014) who find that during what they call the 'globalization era' (after 1990) recessions in LatAm are shorter in duration and milder in amplitude. As discussed in Blanchard and Simon (2001) and Harding and Pagan (2002), an increase in the mean growth rate combined with a reduction in volatility implies business cycles with fewer and shorter recessions.

[FIGURE 3 ABOUT HERE]

[FIGURE 4 ABOUT HERE]

4.3 Discussion: Structural Breaks and Structural Reform

Between the mid 1980s and late 1990s LatAm countries embarked in a process of structural reform of their economies inspired by the "Washington Consensus". These changes are documented in great detail in Lora (1997, 2012), Morley et al. (1999), and Escaith and Paunovic (2004). While initial reactions suggested disappointment with post-reform growth, Easterly et al. (1997) argue that growth in the region was in fact stronger during the 1990s than the previous decade. In addition, de Carvalho Filho and Chamon (2012) argue that reforms led to large improvements in real household income and a substantial reduction in income inequality. The results presented above provide more evidence in this direction. That is, the important changes in economic policies of the 1980s and 1990s have been effective in permanently improving economic growth in LatAm.

For example, Figure 5 plots the smoothed probabilities of structural break from the MS-AR(1) and MS-ARX(1) models and the (normalized) indexes of structural reform of Lora (2012) and Escaith and Paunovic (2004) for Argentina, Brazil, Mexico, and Peru.[8] Existence of a structural break in real GDP growth around the time of the structural reforms provides strong evidence of the effectiveness of these policy changes in Argentina, Brazil, and Peru. But while Easterly et al. (1997) document only an improvement in the average growth rate of output, the results presented above show that the structural reforms of the 1980s and 1990s also led to very important reductions in the volatility of real GDP growth: 49% for Argentina, 51% for Brazil, and 67% for Peru. For the remaining countries (Chile, Colombia, and Mexico), the data does not cover the period before the structural reforms took place in those countries and, as a result, such calculations are not possible.[9]

[FIGURE 5 ABOUT HERE]

In addition to the structural breaks identified in the early 1990s, Chile and Colombia show evidence of a break in 1999 and 2000, respectively. These breaks, however, do not appear to be linked to the structural reforms inspired by the "Washington Consensus" but to the adoption of inflation targeting regimes in these countries. For example, García-Solanes and Torrejón-Flores (2012) argue that the starting date of the inflation targeting regimes corresponds to the moment when the central banks began publishing

[8] If I_t is the value of a structural reform index at time t, the normalized index I_t^* is calculated as $I_t^* = \frac{I_t - \min(I)}{\max(I) - \min(I)}$, with $I^* \in [0, 1]$.

[9] In Chile the main policy changes were implemented in the 1970s and in Mexico in the early 1980s. In both cases the data is only available for the period 1983Q1–2010Q4. In the case of Colombia the data is only for the period 1994Q1–2010Q4.

inflation reports with multi-year targets. These dates are May 2000 in Chile and January 1999 in Colombia, while the estimated break dates are 1999Q2 and 2000Q2, respectively. The reductions in the volatility of GDP growth associated with these breaks are important: 46% for Chile and 20% for Colombia.[10] This result is consistent with the findings of Gonçalves and Salles (2008) and García-Solanes and Torrejón-Flores (2012) who show that the adoption of inflation targeting regimes led to lower variability in GDP growth.

4.4 Discussion: The Effect of Commodity Prices

Recent research by Izquierdo et al. (2008) and Camacho and Perez-Quiros (2013) has shown the existence of a positive link between LatAm output growth and commodity prices. Consistent with this result, I find strong evidence of a positive and linear relationship between the growth rate of real GDP and the growth rate of commodity prices for five of the six LatAm countries considered. As a result, in this section I ask: How much of the recent improvement in LatAm growth can be attributed to the boom in commodity prices observed during the last decade?

To answer this question, Table 6 reports the average contribution of changes in commodity prices to real GDP growth calculated as $(\hat{\beta}_1 + \hat{\beta}_2)\bar{x}$, with $\bar{x}$ the average growth rate of commodity prices in the sample. The results are based on the estimates of Model III (Table 5) and reported for three relevant sub-samples. As we can observe, changes in commodity prices had a small and sometimes negative effect on real GDP growth in the 1980s and 1990s. On the other hand, the contribution during the last sub-

[10]In addition, when the model allows for two structural breaks, Peru exhibits a 45% reduction in the volatility of real GDP growth in 2003Q2. The second break is located about a year after the adoption of inflation targeting (June 2002 according to García-Solanes and Torrejón-Flores, 2012). Results not reported.

sample (the period 2003–2010) was positive and larger in magnitude for all countries. For example, during this period the average contributions to real GDP growth range from around 0.5% annual growth (Argentina, Chile, and Colombia) to almost 2% (Peru). As a result, the sustained increase in commodity prices observed in recent years explains an important share of LatAm growth since 2003.

[TABLE 6 ABOUT HERE]

4.5 Robustness

This section examines the robustness of the results presented above to an alternative index of commodity prices. In this case, instead of using the country-specific indexes of Chen and Lee (2013), the models are estimated using the index of all commodity prices constructed by The Economist.[11] One advantage of using an index of all commodity prices is that potential endogeneity concerns should be mitigated.[12] Table 7 reports ML estimates of the linear MS-ARX(1) models (Model III) using this index of all commodity prices. Overall, the results are quite similar to those obtained using country-specific price indexes (Table 5) and, for all countries except Argentina, the results of the tests are unchanged. In the case of Argentina, we now find a stronger and significant effect of commodities on real GDP growth.

[TABLE 7 ABOUT HERE]

[11] Quarterly data was obtained from Global Financial Data (Ticker: CMECALLW).

[12] In addition to the robustness results reported in this section, generalized method of moments regressions with the lagged commodity prices Δp_{t-1} and Δp_{t-2} as instruments for Δp_t were estimated for the pre- and post-break samples. In the case of Argentina, Chile, Colombia, and Peru, the data appears to support the model specification as all J-tests fail to reject at the 5% level. In the case of Brazil and Mexico there is some evidence of misspecification in the post-break samples but not before the break.

5 Conclusion

This paper documents strong evidence of a structural break in real GDP of six LatAm countries towards stronger mean growth and a substantial reduction in volatility. The timing of the breaks suggests that the important changes in economic policies of the 1980s and 1990s have been effective in permanently improving economic growth in the region. But while the literature has documented mainly an improvement in the mean growth rate of real GDP (e.g., Easterly et al., 1997), this paper documents substantial reductions in the volatility of real GDP growth (49% for Argentina, 51% for Brazil, and 67% for Peru). These changes in the growth processes imply recessions that are shorter in duration and milder in amplitude, consistent with recent findings of Aiolfi et al. (2011) and Calderón and Fuentes (2014). In addition, there is strong evidence of a positive and linear relationship between the growth rate of real GDP and the growth rate of commodity prices. As a result, the sustained increase in commodity prices observed in recent years explains an important share of LatAm growth since 2003. But after accounting for the effect of commodity prices, there is even stronger evidence of a structural break in real GDP growth towards an increase in mean growth and a reduction in volatility.

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Table 1: Quarterly Summary Statistics

	ARG	BRA	CHI	COL	MEX	PER	LatAm
Real GDP growth: Mean (%)							
1983Q1–1992Q4	0.37	-0.10	1.60	–	0.53	-0.25	0.43
1993Q1–2002Q4	0.13	0.71	1.16	0.45	0.67	1.06	0.70
2003Q1–2010Q4	1.87	0.97	1.04	1.14	0.56	1.63	1.20
Full Sample	0.71	0.72	1.28	0.79	0.59	0.75	0.81
Real GDP growth: Standard Deviation (%)							
1983Q1–1992Q4	3.30	2.64	1.86	–	1.50	4.97	2.85
1993Q1–2002Q4	2.28	1.47	1.60	1.28	1.78	1.80	1.70
2003Q1–2010Q4	1.30	1.54	1.14	0.95	1.58	1.09	1.27
Full Sample	2.58	1.68	1.59	1.18	1.62	3.28	1.99
Commodity price growth: Mean (%)							
1983Q1–1992Q4	-0.06	-0.07	0.67	-1.10	-1.18	-0.10	-0.31
1993Q1–2002Q4	0.06	-0.09	-0.66	0.29	0.80	-0.25	0.02
2003Q1–2010Q4	2.87	3.72	4.52	3.57	3.55	4.27	3.75
Full Sample	0.82	1.01	1.29	0.73	0.88	1.09	0.97

Notes: Real GDP growth statistics are computed for the sample period 1983Q1–2010Q4 for all countries except Brazil and Colombia. For Brazil, the sample period is 1990Q1–2010Q4. For Colombia, the sample period is 1994Q1–2010Q4. Commodity price mean growth rates are computed for the period 1983Q1–2010Q4. LatAm values are a simple average of the corresponding values (mean growth rates, standard deviations) of the six countries considered.

Table 2: AR(1) OLS Estimates and qLL Tests for Stability

	ARG	BRA	CHI	COL	MEX	PER
Specification: $\Delta y_t = c + \phi \Delta y_{t-1} + \varepsilon_t$, $\varepsilon_t \sim$ iid $\mathrm{N}(0, \sigma_\varepsilon^2)$						
c	0.54 (0.25)	0.56 (0.20)	1.17 (0.20)	0.67 (0.18)	0.53 (0.16)	0.54 (0.29)
ϕ	0.24 (0.09)	0.15 (0.11)	0.08 (0.10)	0.12 (0.13)	0.13 (0.10)	0.38 (0.09)
σ_ε	2.53	1.63	1.60	1.18	1.61	3.00
qLL_1	-14.19*	-8.27	-9.06	-11.97	-14.58**	-13.43*
Specification: $\sqrt{\pi/2} \times \lvert\hat{\varepsilon}_t\rvert = s + \eta_t$						
s	2.35 (0.20)	1.42 (0.16)	1.47 (0.13)	1.13 (0.11)	1.40 (0.14)	2.54 (0.26)
σ_η	2.09	1.43	1.34	0.92	1.43	2.73
qLL_2	-9.06**	-5.81	-10.58**	-2.64	-7.36*	-9.15**

Notes: Δy_t is quarterly real GDP growth. Standard errors are in parentheses to the right of the OLS estimates. The sample period is 1983Q1–2010Q4 for all countries except Brazil and Colombia. For Brazil, the sample period is 1990Q1–2010Q4. For Colombia, the sample period is 1994Q1–2010Q4. qLL test is described in Elliott and Muller (2006). qLL_1 is the test statistic of parameter stability in c and ϕ. 10% and 5% critical values for the qLL_1 test are -12.80 and -14.32, respectively. qLL_2 is the test statistic of parameter stability in s. 10% and 5% critical values for the qLL_2 test are -7.14 and -8.36, respectively. * (**) denotes rejection of the null hypothesis of parameter stability at 10% (5%) level.

Table 3: Model Selection

	Model	Breaks	Δp_t	Δp_t^+	$\ln L$	k	AIC
ARG	AR(1)	0	No	No	-158.45	3	161.45
	MS-AR(1)	1	No	No	-143.74	7	**150.74**
	MS-ARX(1)	1	Yes	No	-142.68	9	151.68
	MS-ARX(1)	1	Yes	Yes	-142.31	11	153.31
BRA	AR(1)	0	No	No	-82.02	3	85.02
	MS-AR(1)	1	No	No	-77.11	7	84.11
	MS-ARX(1)	1	Yes	No	-71.55	9	**80.55**
	MS-ARX(1)	1	Yes	Yes	-71.29	11	82.29
CHI	AR(1)	0	No	No	-107.22	3	110.22
	MS-AR(1)	1	No	No	-99.49	7	106.49
	MS-ARX(1)	1	Yes	No	-96.66	9	**105.66**
	MS-ARX(1)	1	Yes	Yes	-96.05	11	107.05
COL	AR(1)	0	No	No	-42.68	3	45.68
	MS-AR(1)	1	No	No	-39.22	7	46.22
	MS-ARX(1)	1	Yes	No	-36.62	9	**45.62**
	MS-ARX(1)	1	Yes	Yes	-35.50	11	46.50
MEX	AR(1)	0	No	No	-108.28	3	111.28
	MS-AR(1)	1	No	No	-105.25	7	112.25
	MS-ARX(1)	1	Yes	No	-96.39	9	**105.39**
	MS-ARX(1)	1	Yes	Yes	-94.61	11	105.61
PER	AR(1)	0	No	No	-179.78	3	182.78
	MS-AR(1)	1	No	No	-148.02	7	155.02
	MS-ARX(1)	1	Yes	No	-141.69	9	**150.69**
	MS-ARX(1)	1	Yes	Yes	-140.35	11	151.35

Notes: The sample period is 1983Q1–2010Q4 for all countries except Brazil and Colombia. For Brazil, the sample period is 1990Q1–2010Q4. For Colombia, the sample period is 1994Q1–2010Q4. $\ln L$ denotes the log likelihood. AIC denotes the Akaike Information Criterion and is calculated as $-\ln L + k$ where k is the number of parameters in the model.

Table 4: MS-AR(1) ML Estimates (Model II)

	ARG	BRA	CHI	COL	MEX	PER
Specification: $\phi_{S_t}(L)(\Delta y_t - \mu_{S_t}) = \varepsilon_t$, $\varepsilon_t \sim$ iid $\mathrm{N}(0, \sigma^2_{S_t})$						
μ_1	0.10 (0.59)	-0.77 (1.10)	1.51 (0.23)	0.41 (0.39)	0.55 (0.18)	-0.39 (1.13)
ϕ_1	-0.05 (0.18)	-0.17 (0.56)	-0.01 (0.13)	0.36 (0.22)	-0.29 (0.15)	0.37 (0.15)
σ_1	3.39 (0.42)	2.92 (0.87)	1.84 (0.16)	1.23 (0.19)	1.41 (0.16)	4.53 (0.53)
μ_2	0.96 (0.42)	0.83 (0.19)	0.99 (0.20)	1.00 (0.14)	0.62 (0.28)	1.33 (0.23)
ϕ_2	0.52 (0.10)	0.14 (0.11)	0.26 (0.13)	-0.20 (0.15)	0.31 (0.12)	0.24 (0.12)
σ_2	1.74 (0.15)	1.44 (0.12)	1.00 (0.11)	0.99 (0.11)	1.60 (0.14)	1.48 (0.13)
p	0.97 (0.03)	0.86 (0.13)	0.98 (0.02)	0.96 (0.04)	0.98 (0.02)	0.98 (0.02)
$\ln L$	-143.74	-77.11	-99.49	-39.22	-105.25	-148.02
$Q^*(1)$	0.03	0.05	0.01	0.05	0.18	0.16
$Q^*(6)$	6.97	7.56	2.04	4.50	6.94	3.71
Date	1991Q3	1992Q1	1999Q2	2000Q2	1993Q2	1993Q1
LR_1	8.41**	1.21	4.47	6.75**	5.87*	3.10
LR_2	21.07**	8.36**	12.78**	1.43	0.74	46.27**

Notes: Δy_t is quarterly real GDP growth. Standard errors are in parentheses to the right of the ML estimates. The sample period is 1983Q1–2010Q4 for all countries except Brazil and Colombia. For Brazil, the sample period is 1990Q1–2010Q4. For Colombia, the sample period is 1994Q1–2010Q4. $\ln L$ denotes the log likelihood and the LR test statistic are constructed as $-2(\ln L_r - \ln L_u)$ where $\ln L_r$ is the log likelihood of the restricted model and $\ln L_u$ is the log likelihood of the unrestricted model. LR is distributed $\chi^2(q)$ where q is the number of restrictions imposed. LR_1 tests $\mu_1 = \mu_2$ and $\phi_1 = \phi_2$. LR_2 tests $\sigma_1 = \sigma_2$. $Q^*(k)$ is the Ljung-Box statistic for the (centered) generalized residuals with k the number of lags. * (**) denotes rejection of the null hypothesis at 10% (5%) level.

Table 5: MS-ARX(1) ML Estimates (Model III)

	ARG	BRA	CHI	COL	MEX	PER
Specification: $\phi_{S_t}(L)\big(\Delta y_t - \mu_{S_t} - \beta(L)\Delta p_t\big) = \varepsilon_t$, $\varepsilon_t \sim$ iid $\mathrm{N}(0, \sigma^2_{S_t})$						
μ_1	0.12 (0.58)	0.57 (0.53)	1.50 (0.22)	0.38 (0.38)	0.62 (0.23)	-0.29 (1.19)
ϕ_1	-0.06 (0.18)	0.14 (0.24)	-0.02 (0.13)	0.33 (0.23)	-0.07 (0.13)	0.41 (0.15)
σ_1	3.42 (0.43)	2.17 (0.34)	1.83 (0.16)	1.23 (0.19)	1.89 (0.18)	4.57 (0.53)
μ_2	0.90 (0.40)	0.62 (0.14)	0.90 (0.17)	0.94 (0.13)	0.45 (0.19)	1.13 (0.19)
ϕ_2	0.51 (0.10)	-0.20 (0.15)	0.15 (0.17)	-0.21 (0.15)	0.24 (0.13)	0.09 (0.12)
σ_2	1.71 (0.15)	1.21 (0.12)	0.94 (0.10)	0.93 (0.10)	0.98 (0.10)	1.33 (0.12)
β_1	0.03 (0.02)	0.11 (0.03)	0.01 (0.01)	0.03 (0.01)	0.05 (0.01)	0.06 (0.02)
β_2	0.02 (0.02)	-0.01 (0.02)	0.02 (0.01)	-0.01 (0.01)	0.02 (0.01)	0.05 (0.02)
p	0.97 (0.03)	0.96 (0.04)	0.99 (0.02)	0.96 (0.04)	0.98 (0.02)	0.98 (0.02)
$\ln L$	-142.68	-71.55	-96.66	-36.62	-96.39	-141.69
$Q^*(1)$	0.02	0.02	0.00	0.01	0.03	0.12
$Q^*(6)$	7.62	7.46	3.35	2.01	5.88	2.82
Date	1991Q3	1996Q2	1999Q3	2000Q1	1998Q1	1993Q2
LR_1	8.10**	0.63	4.85*	6.18**	2.92	4.86*
LR_2	22.57**	9.30**	15.58**	2.12	5.37**	56.22**
LR_3	2.12	11.11**	5.66*	5.20*	17.73**	12.66**

Notes: Δy_t is quarterly real GDP growth and Δp_t is the quarterly growth rate in commodity prices. Standard errors are in parentheses to the right of the ML estimates. The sample period is 1983Q1–2010Q4 for all countries except Brazil and Colombia. For Brazil, the sample period is 1990Q1–2010Q4. For Colombia, the sample period is 1994Q1–2010Q4. $\ln L$ denotes the log likelihood and the LR test statistic are constructed as $-2(\ln L_r - \ln L_u)$ where $\ln L_r$ is the log likelihood of the restricted model and $\ln L_u$ is the log likelihood of the unrestricted model. LR is distributed $\chi^2(q)$ where q is the number of restrictions imposed. LR_1 tests $\mu_1 = \mu_2$ and $\phi_1 = \phi_2$. LR_2 tests $\sigma_1 = \sigma_2$. LR_3 tests $\beta_1 = \beta_2 = 0$. $Q^*(k)$ is the Ljung-Box statistic for the (centered) generalized residuals with k the number of lags. * (**) denotes rejection of the null hypothesis at 10% (5%) level.

Table 6: Commodity Prices and Real GDP Growth

	ARG	BRA	CHI	COL	MEX	PER
1983Q1–1992Q4	-0.00	-0.01	0.02	-0.02	-0.08	-0.01
1993Q1–2002Q4	0.00	-0.01	-0.02	0.01	0.05	-0.03
2003Q1–2010Q4	0.13	0.38	0.15	0.07	0.23	0.47

Notes: The average contribution of commodity prices to real GDP growth is computed as $(\hat{\beta}_1 + \hat{\beta}_2)\bar{x}_i$ for $i = 1, 2, 3$ with $\bar{x}_i$ the average growth rate of commodity prices in the sub-sample i. The results are based on the estimates of Model III (Table 5).

Table 7: MS-ARX(1) ML Estimates (Model III) - Economist Index

	ARG	BRA	CHI	COL	MEX	PER
Specification: $\phi_{S_t}(L)\big(\Delta y_t - \mu_{S_t} - \beta(L)\Delta p_t\big) = \varepsilon_t,\ \varepsilon_t \sim$ iid $\mathrm{N}(0, \sigma^2_{S_t})$						
μ_1	0.06 (0.59)	0.54 (0.46)	1.50 (0.22)	0.41 (0.36)	0.46 (0.30)	-0.33 (1.20)
ϕ_1	-0.05 (0.18)	0.13 (0.24)	-0.04 (0.13)	0.35 (0.22)	0.10 (0.14)	0.41 (0.15)
σ_1	3.42 (0.43)	1.94 (0.29)	1.79 (0.16)	1.12 (0.17)	1.89 (0.19)	4.57 (0.53)
μ_2	0.83 (0.39)	0.66 (0.11)	0.92 (0.19)	0.89 (0.12)	0.54 (0.23)	1.18 (0.17)
ϕ_2	0.51 (0.10)	-0.33 (0.13)	0.25 (0.15)	-0.26 (0.14)	0.21 (0.14)	0.02 (0.12)
σ_2	1.63 (0.14)	1.04 (0.10)	0.94 (0.10)	0.91 (0.10)	1.11 (0.12)	1.33 (0.12)
β_1	0.06 (0.03)	0.12 (0.02)	-0.01 (0.02)	0.05 (0.02)	0.06 (0.02)	0.06 (0.02)
β_2	0.05 (0.03)	0.00 (0.02)	0.05 (0.02)	-0.00 (0.02)	0.04 (0.02)	0.06 (0.02)
p	0.97 (0.03)	0.96 (0.04)	0.99 (0.02)	0.96 (0.04)	0.98 (0.02)	0.98 (0.02)
$\ln L$	-139.38	-60.93	-94.77	-33.50	-100.81	-141.26
$Q^*(1)$	0.03	0.00	0.00	0.01	0.13	0.30
$Q^*(6)$	6.47	7.03	6.92	0.71	5.95	2.92
Date	1991Q3	1996Q2	1999Q2	2000Q1	1997Q3	1993Q1
LR_1	7.76**	2.74	5.81*	5.87*	0.34	6.61**
LR_2	23.00**	13.10**	15.09**	1.20	5.45**	53.51**
LR_3	8.71**	32.36**	9.43**	11.45**	8.88**	13.53**

Notes: Δy_t is quarterly real GDP growth and Δp_t is the quarterly growth rate in commodity prices. Standard errors are in parentheses to the right of the ML estimates. The sample period is 1983Q1–2010Q4 for all countries except Brazil and Colombia. For Brazil, the sample period is 1990Q1–2010Q4. For Colombia, the sample period is 1994Q1–2010Q4. $\ln L$ denotes the log likelihood and the LR test statistic are constructed as $-2(\ln L_r - \ln L_u)$ where $\ln L_r$ is the log likelihood of the restricted model and $\ln L_u$ is the log likelihood of the unrestricted model. LR is distributed $\chi^2(q)$ where q is the number of restrictions imposed. LR_1 tests $\mu_1 = \mu_2$ and $\phi_1 = \phi_2$. LR_2 tests $\sigma_1 = \sigma_2$. LR_3 tests $\beta_1 = \beta_2 = 0$. $Q^*(k)$ is the Ljung-Box statistic for the (centered) generalized residuals with k the number of lags. * (**) denotes rejection of the null hypothesis at 10% (5%) level.

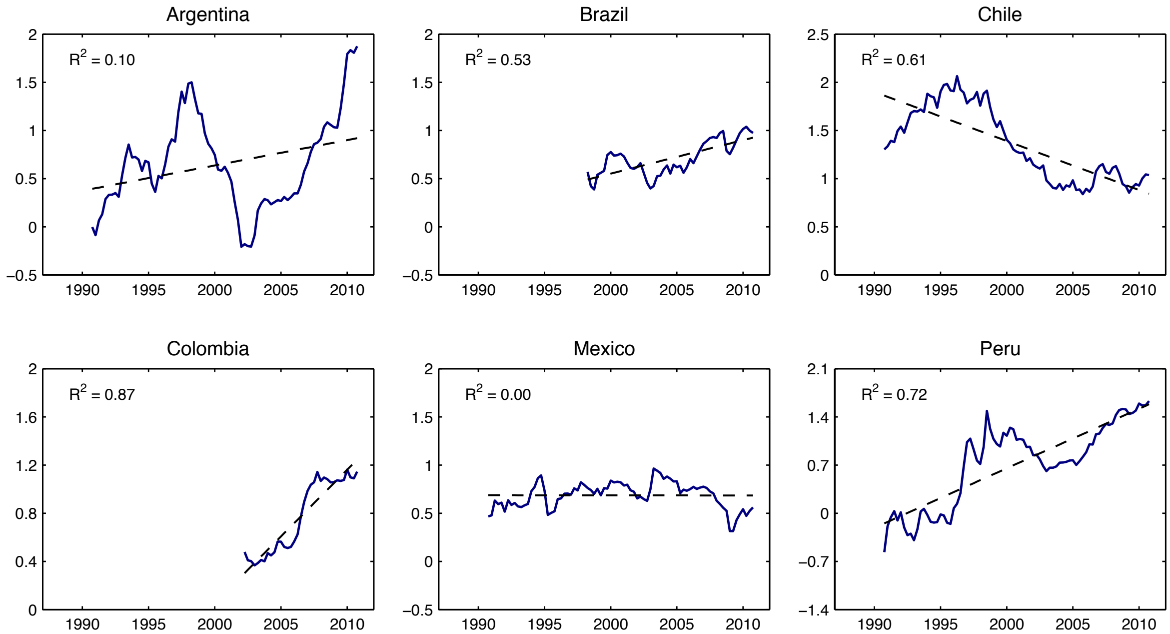

Figure 1: 8-year rolling averages of real GDP quarterly growth rates and time trends. Date on horizontal axis indicates the date of the last observation for which the 8-year average is calculated.

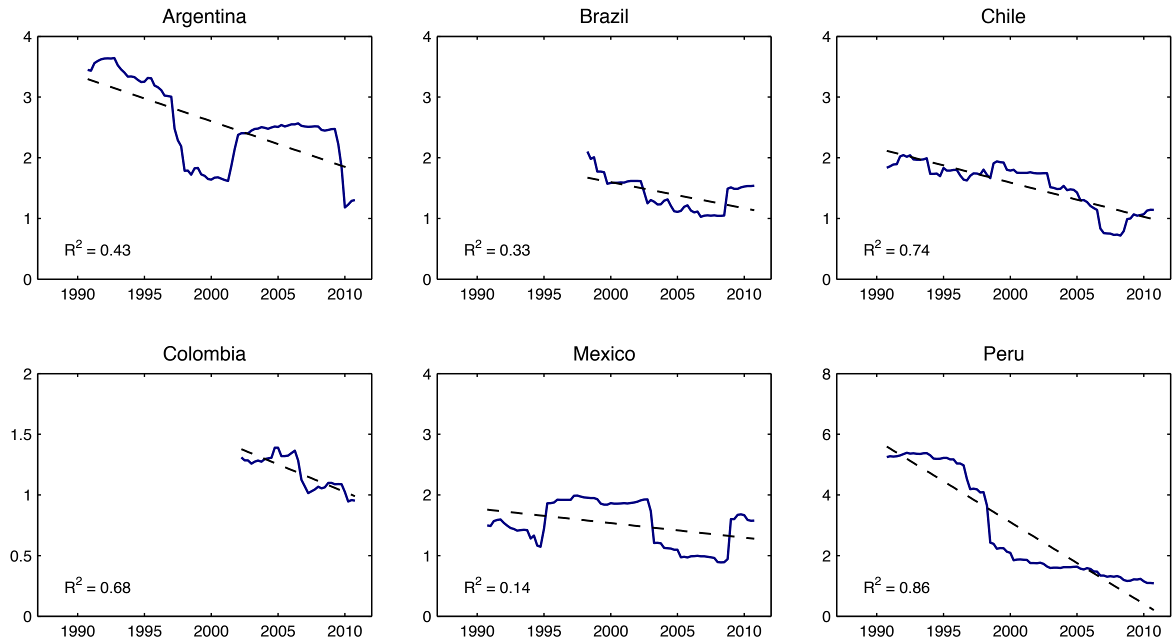

Figure 2: 8-year rolling standard deviations of real GDP quarterly growth rates and time trends. Date on horizontal axis indicates the date of the last observation for which the 8-year average is calculated.

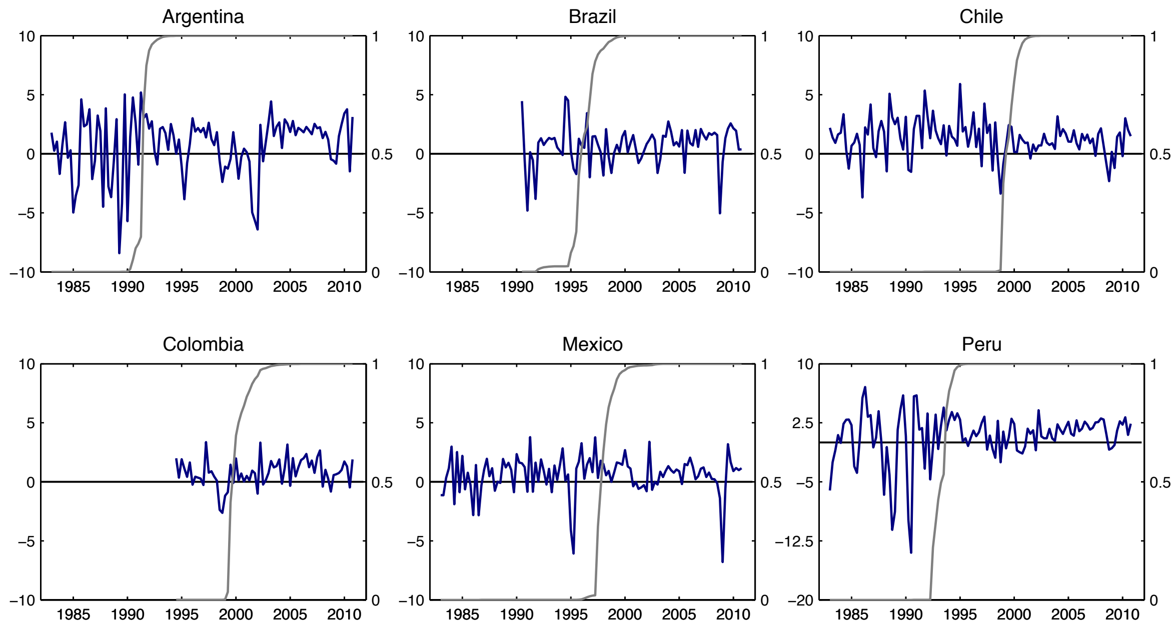

Figure 3: Quarterly real GDP growth and (smoothed) probabilities of structural break computed from the MS-ARX(1) models (Model III).

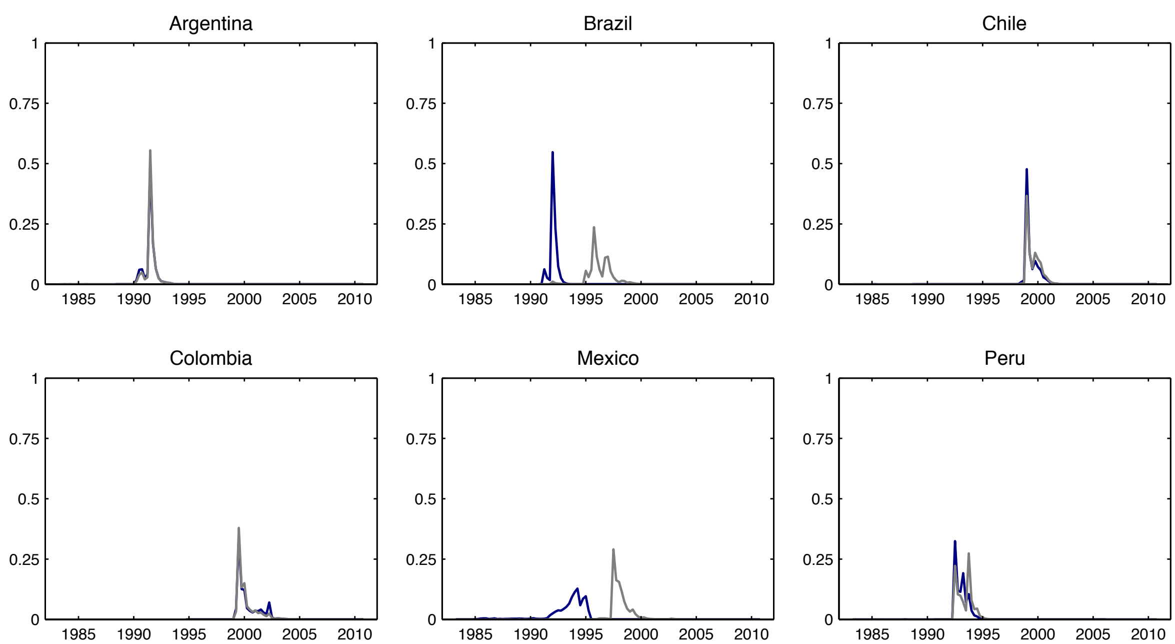

Figure 4: Break date densities (change in smoothed probabilities) computed from the MS-AR(1) models (blue) and the MS-ARX(1) models (gray).

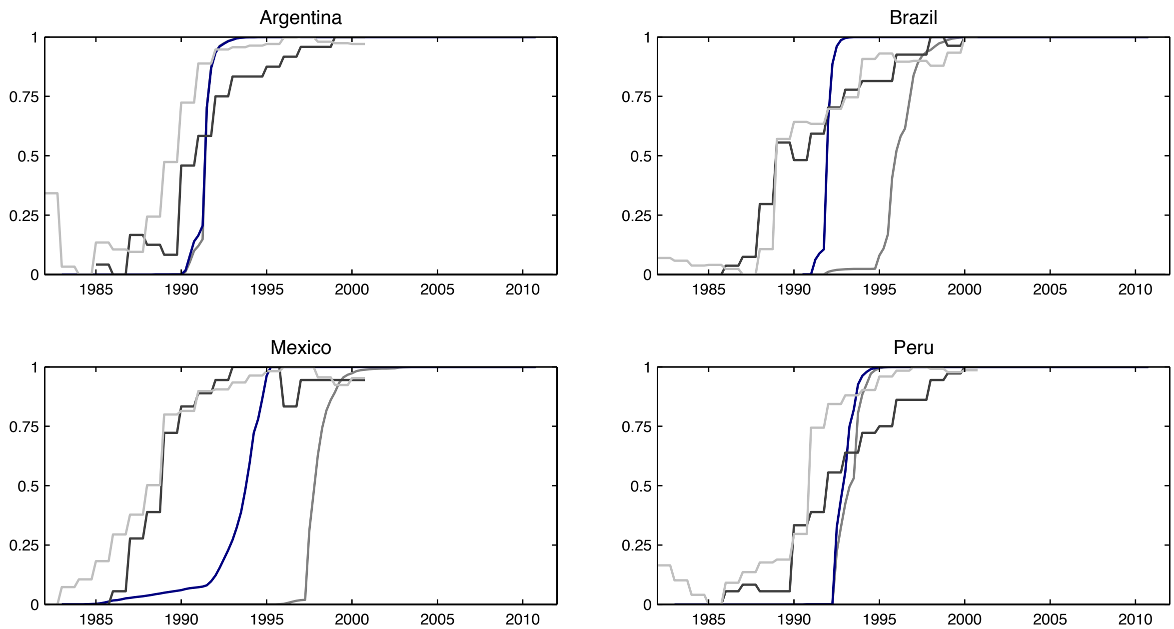

Figure 5: Probabilities of structural break computed from the MS-AR(1) models and the MS-ARX(1) models (smooth lines) and the (normalized) indexes of structural reform of Lora (2012) and Escaith and Paunovic (2004) (step lines).