Asymmetries and Common Cycles in Latin America: Evidence from Markov-Switching Models

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Abstract: Markov-switching models are estimated to characterise expansions and contractions for Latin American countries. In general, univariate analysis results imply that recessions are deeper in absolute magnitude, less persistent, and more volatile than expansions. From an international perspective, it is found that there is not a common Latin American cycle, but there exists some evidence about common regime shifts and cycles between Brazil-Peru and Chile-United States. However, it seems that their causes are very different and related to common shocks and similar policies. Therefore, it is concluded that individual business cycles are largely independent in Latin America.

Resumen: Se usan modelos de cambio de régimen para caracterizar expansiones y contracciones de varios países latinoamericanos. Del análisis univariado se obtiene que las recesiones son más agudas en magnitud absoluta, menos persistentes y más volátiles que las expansiones. Asimismo, se concluye que no hay un ciclo económico latinoamericano, pero sí cambios simultáneos de régimen y, por tanto, ciclos comunes entre Brazil y Perú y entre Chile y Estados Unidos. Sin embargo, al parecer sus causas son diferentes y están relacionadas con choques comunes y políticas económicas similares. Por lo tanto, se concluye que los ciclos individuales son fundamentalmente independientes.

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1. Introduction

For many decades economists have realized that economic downturns are typically brief and severe, whereas upturns are longer and more gradual (Mitchell, 1927, and Keynes, 1936). Also, recently Kähler and Marnet (1992) and Hamilton (1993) have reported that variances in expansions are different to variances in recessions. However, it has been since only the first half of the 1980s that a branch of the literature has paid attention to the analysis of the asymmetric behaviour of economies over the business cycle (see for example Neftci, 1984; DeLong and Summers, 1986; Hamilton, 1989; Sichel, 1989; and Artis, Kontolemis, and Osborn, 1997). Some of these features are apparent in the Latin American experience. In a recent paper, Mejía-Reyes (1999) analyses the experience of the same sample of 8 countries, that is used here and shows that asymmetric behaviours are present in many cases. In particular, he reports that the range of variation of the growth rates of GDP per capita is really wide. Except in one case, there is a difference of at least 18 percentage points between the minimum and the maximum growth rates of GDP for the other countries over the period 1951-1995, while the largest range equates 31 percentage points. Also, he finds three characteristics that support the claims of Mitchell and Keynes: *1)* the minimum GDP growth rate value is greater than the maximum one in absolute terms for most countries; *2)* the skewness is negative and the median is greater than the mean for all economies; *3)* there is excess of kurtosis in five cases, which may reflect the importance of the minimum growth rates.[1] Despite this evidence, most analyses of Latin American fluctuations have used linear methods and, consequently, ignored the role of asymmetries.

Existing studies of Latin American cyclical fluctuations have found that real GDP exhibits significant persistence and that current shocks have permanent effects (Mejía-Reyes and Hernández-Veleros, 1998; Ruprah, 1991; Cuddington and Urzúa, 1988). Also, they have shown that supply shocks and real factors tend to dominate economic fluctuations even in the short-run (Hoffmaister and Roldós, 1996, 1997; Kydland and Zarazaga, 1997). Recently, however, evidence of nonlinea-

[1] DeLong and Summers (1986) argue that the claims of Mitchell and Keynes imply that there should be significant skewness in a frequency distribution of the growth rates of output (that is, the distribution should have significantly fewer than half its observations below the mean) and the median output growth rate should exceed the mean by an important amount. They argue also that when the kurtosis is significant there may be important outliers.

rity and asymmetries over the Colombian business cycle has been found (Mora, 1997), and also asymmetries over the business cycles of a number of Latin American countries (Mejía-Reyes, 1999).

On the other hand, there are few studies that address international business cycles for Latin America and the results are not conclusive. By using different methodologies, it has been found that significant correlations are limited to small groups of countries (Engel and Issler, 1993; Arnaudo and Jacobo, 1997), that economic fluctuations are highly variable and not time uniform (Arnaudo and Jacobo, 1997), and that most correlations become non-significant when the post-debt crisis period is included in the sample (Iguíñiz and Aguilar, 1998). Also, evidence points to the synchronisation of business cycles regimes for only some countries (Mejía-Reyes, 1999).

These studies represent significant advances for the comprehension of Latin American business cycles. However, only a few of them address issues of nonlinearities and regime characteristics. In this paper, we apply the approach proposed by Hamilton (1989) to get a deeper look at the asymmetries reported above. One of the main characteristics of this approach is that it uses the observed growth rate to identify in a probabilistic sense and without imposing any constraint a priori which of the two regimes the economy is in at each time period. Then each regime is characterised with respect to the magnitude of its mean growth rate, regime persistence, duration, and volatility. When these characteristics differ over regimes, we say that asymmetries are present. In addition, we evaluate the ability of the estimated Markov-switching models to identify periods of expansion and contraction in the process of economic growth by comparing the regimes implied by this kind of model with those obtained by Mejía-Reyes (1999) (hereafter, MR) from the application of a classical business cycles approach. Also, we analyse the international nature of business cycles by applying multivariate Markov-switching models, which allow us to model economic growth across countries as a joint stochastic process, so that we can identify common regime shifts and consequently common cycles.

This paper is structured as follows. In the next section we present the main features of the Markov-switching models applied in this paper. Then we discuss the strategy of specification and estimation, and report the results for eight Latin American countries and the United States. In general, we find evidence of asymmetric behaviour over the business cycle, but we cannot find a Latin American business cycle. Finally, we summarise our results and state some general conclusions.

2. Markov-Switching Processes as Stochastic Business Cycle Models

The Markov-switching autoregressive model proposed by Hamilton (1989) has become increasingly popular for the empirical analysis of macroeconomic fluctuations (for example Goodwin, 1993; Kähler and Marnet, 1992; Krolzig, 1996 and 1997a; and Clements and Krolzig, 1998). In contrast to previous approaches, Hamilton specifies the first difference of the observed series (GNP in particular) as a nonlinear stationary process, where the nonlinearities arise if the process is subject to discrete shifts in regimes –episodes across which the behaviour of the series is markedly different– and where the state of the economy is treated as an unobserved latent variable. In that sense, this innovative approach allows researchers to overcome the shortcoming of linear models to deal with the asymmetry between expansions and contractions that have been documented by Neftci (1984) and Sichel (1993) for the US business cycle, by Artis, Kontolemis, and Osborn (1997) for a group of industrial countries, and by Mora (1997) and Mejía-Reyes (1999) for some Latin American countries.

In particular, Markov-switching autoregressive processes in the growth rate of real GNP are interpreted as stochastic business cycle models, where expansions and contractions are modelled as switching regimes of such a stochastic process.[2] The regimes are associated with different conditional distributions of the growth rate for each regime, whereas it is assumed that the mean of the growth rate is positive in expansions and negative in contractions. Within this framework, the approach consists of solving the actual marginal likelihood function for GNP and maximizing the likelihood function with respect to the population parameters. Also, as a by-product, the approach allows us to calculate optimal inference on the latent state of the economy by assigning probabilities to the unobserved expansion and contraction regimes conditional on the available information set, that is on the observed behaviour of the series.

[2] In this paper it is preferred to refer to declines in real GDP per capita as "contractions" rather than "recessions" in order to include short-run declines as well as dramatic declines.

2.1. Markov-Switching Autoregressive (MS-AR) Models

Hamilton (1989) models the changes in regime as the result of an unobserved state (or regime) variable s_t which is largely unrelated to past realizations of the series. The state variable is a random variable that follows an ergodic Markov chain process and represents different states at which the system can be at each time t. In particular the model suggested by Hamilton for two states of the economy can be described as follows.

Let y_t be a stationary series that is governed by a p^{th} order autoregressive process

$$\left(y_t - \mu\left(s_t\right)\right) = \phi_1\left(y_{t-1} - \mu\left(s_{t-1}\right)\right) + \phi_2\left(y_{t-2} - \mu\left(s_{t-2}\right)\right) + \ldots + \phi_p\left(y_{t-p} - \mu\left(s_{t-p}\right)\right) + \varepsilon_t \quad (1)$$

where ε_t ~ iid N $(0, \sigma^2)$ and the conditional mean, $\mu(s_t)$, switches between two states

$$\mu\left(s_t\right) = \begin{cases} \mu_1 < 0 \\ \mu_2 > 0 \end{cases}$$

when $s_t = 1$ and $st = 2$, respectively. We identify state 1 with contractions and state 2 with expansions. The effect of the regime s_t on the variable y_t is given by the conditional probability density function $p\ (y_t \,|\, s_t)$.

The description of the data-generating process is completed with the formulation of a model for the regime generating process. Specifically, the variable s_t is assumed to be a discrete-valued random variable that can only assume an integer value, {1, 2}. The probability that s_t equals some particular value j depends on the past through the most recent value s_{t-1} according to the following definition:

$$P\left\{s_t = j \middle| s_{t-1} = i, s_{t-2} = k, \ldots\right\} = P\left\{s_t = j \middle| s_{t-1} = i\right\} = p_{ij} \quad (2)$$

Such a process is described as a two-state ergodic Markov chain with transition probabilities $\{p_{ij}\}$ for $i, j = 1, 2$. The latter indicate the probability that the economy switches from regime i in period $t-1$ to regime j in period t. Note that the transition probabilities must satisfy $p_{11} + p_{12} = p_{21} + p_{22} = 1$, which implies that $p_{12} = 1 - p_{11}$ and $p_{21} = 1 - p_{22}$. The transition probability p_{jj} can be interpreted as a regime persistence measure in the sense that gives information about the probability of the economy continuing in the same regime in the next period. Also, the expected duration of regime j is given by $(1 - p_{jj})^{-1}$.

The model for y_t is estimated by maximum likelihood. Following Krolzig (1996, 1997b) we will denote these models as MSM (M)-AR(1) (with M = {1, 2}) to indicate that we are dealing with 2 state Markov-switching mean models, time invariant autoregressive coefficients, and no heteroscedasticity.

The approach above can be generalized to consider the variance of ε_t in (1) also depending on the regime to give account of the differences in contraction and expansion variances that some authors have reported (see Kähler and Marnet, 1992, and Hamilton, 1993). Because this sort of models allows us to have changes in means and variances we denote them as MSH (M)-AR(p) (with M = {1, 2}) to indicate that we are working with heteroscedastic Markov-switching models with an AR(p) component.

Note that the state s_t is not observed directly. Thus the probability that the process is in state 1 at date t, with the inference conditioned on data observed through that date and on the estimated value of the population parameter vector θ, is given by the filter probability defined as $p\left(s_t = 1 \middle| y_t, y_{t-1}, \ldots, y_1; \theta\right)$. We can also calculate the probability conditioned on data observed through the full sample, namely, $p\left(s_t = 1 \middle| y_T, y_{T-1}, \ldots, y_1; \theta\right)$. This probability is called the smoother probability and can also be calculated for any t (see Hamilton, 1989, 1993).

Hamilton's algorithm can also be considered as a formalization of the statistical identification of turning points in a time series with the filter and, particularly, the smoother probabilities used to identify periods of contraction and expansion. The metric proposed by Hamilton is based on the idea that the econometrician would conclude that the economy is more likely than not to be in a contraction in period t when

$$P\left(S_t = 1 \middle| y_T, y_{T-1}, \ldots, y_1\right) > 0.5 \tag{3}$$

On the basis of this rule a binary variable is generated such that it equates 1 when the rule is satisfied and 0 otherwise. Similarly, the rule allows us to define turning points according to the following definitions: date τ is designated a peak if $P\left(S_\tau = 1 \middle| y_T, y_{T-1}, \ldots, y_1\right) < 0.5$ and $P\left(S_{\tau+1} = 1 \middle| y_T, y_{T-1}, \ldots, y_1\right) > 0.5$. Likewise, a date τ is designated trough if $P\left(S_\tau = 1 \middle| y_T, y_{T-1}, \ldots, y_1\right) > 0.5$ and $P\left(S_{\tau+1} = 1 \middle| y_T, y_{T-1}, \ldots, y_1\right) < 0.5$. The observation of Hamilton that this decision rule is not too restrictive because very few values of the smoothed probability lie between 0.3 and 0.7 is still valid for our estimations, as we will see later.

2.2. Markov-Switching Vector Autoregressive (MS-VAR) Models

Markov-switching models have been used to analyse common contemporaneous states of several variables (Engel and Hamilton, 1990; Phillips, 1991; Hamilton, 1993; Krolzig, 1997a). These Markov-switching vector autoregressive models (MS-VAR) generalize the model described in expression (1) to a multivariate context. Conditional on the state of the proces, the N-dimensional vector of stationary time series $y_t = (y_{1t}, y_{2t},..., y_{Nt})$, for $t = 1, 2,..., T$, is generated by a vector autoregression of order p,

$$\left(y_t - \mu\left(s_t\right)\right) = A_1\left(y_{t-1} - \mu\left(s_{t-1}\right)\right) + \ldots + A_p\left(y_{t-p} - \mu\left(s_{t-p}\right)\right) + u_t \qquad (4)$$

where the pre-sample values $y_0,..., y_{1-p}$ are fixed and u_t is a Gaussian white noise process with mean zero and variance-covariance matrix Σ which may be state dependent, namely, u_t ~ NID(0, $\Sigma(s_t)$). Thus we have an MSH(M)-VAR(p) model, where the mean and the variance change when the regime of the process changes.

In MS-VAR models for analysing international business cycles, the dynamic propagation mechanism of impulses to the system consists of two components. The first one is a linear autoregression representing the international (from country m to country n) and inter-temporal (lag k) transmission of country-specific shocks because lagged values of each y_i , for $i = 1, 2,..., N$, enter in the VAR component. Thus, those effects are determined by the elements of the autoregressive matrices $A_k = \{a_{nm,k}\}$. The second component refers to the regime shifts generated by the Markov process, which represents large contemporaneously occurring common shocks. Regime shifts are represented by the switches in the mean vector $\mu\left(s_t\right)$ (and in the variance-covariance matrix). These two sources of fluctuations are not necessarily independent. Thus, changes in regime can simultaneously affect the state of the common business cycles and the international transmission of country-specific shocks.

As before, we assume that the process switches between two states, $s_t = \{1, 2\}$, and that s_t follows an ergodic Markov chain according to expression (2) with joint transition probabilities $\{p_{ij}\}$ for $i, j = \{1, 2\}$, which can be arranged in a transition matrix P. Filter and smoothed probabilities and hence turning points can be obtained from analogous expressions to those defined above.

In the next section we apply the models defined above to analyse the dynamic behaviour of the business cycles of eight Latin American countries and of the United States. Also, we analyse the existence of common cycles.

3. Results

We consider the experience of eight countries: Argentina, Bolivia, Brazil, Chile, Colombia, Mexico, Peru, and Venezuela. We have chosen these countries because they are the largest Latin American economies and because most of them have in common a long period of sustained growth that was interrupted by the international debt crisis in the early 80s. Subsequently, most of them have experienced stabilisation and structural change policies. We analyse the dynamics of the US economy as well in order to analyse the links between its economy and the Latin American ones. To perform our analysis we use series for per capita annual real GDP over the period 1950-1995, which is an updated version of the series real GDP per capita (Laspeyres index, 1985 international prices) from Summers and Heston (1991). Details of the updating are presented in MR.

3.1. Asymmetric Business Cycles and Turning Points

The estimation strategy takes several considerations into account.[3] First, the analysis of unit roots undertaken by MR suggests that all the series of real GDP per capita considered in this study are integrated of order 1, which implies that their first differences are stationary. Thus, we model the series obtained as 100 times the first difference of the natural logarithm of real GDP per capita.

[3] The estimations are undertaken in GAUSS by using numerical optimisation, and the programs are modified versions of those of Hamilton (1989). Although the means of states 1 and 2 are defined to be negative and positive in (1) respectively, the models have been estimated without imposing any restriction *a priori*. Actually, one of the main advantages of this methodology is that it can be employed as an independent objective algorithm for identifying business cycles turning points and, consequently, expansion and recession regimes.

Second, univariate autoregressive linear models were estimated to determine the relevant number of lags, which corresponds to that one that minimizes the Schwarz criterion. This would offer a starting point to determine the value of p.

Third, we have considered various models which are allowed to have an AR(p) component and shifts in mean and variance through the different states. Two criteria were used to select the best model for each country: the statistical significance of the estimated parameters and the ability of the model to track the observed periods of expansions and contractions. In some cases there has been a trade-off between these two criteria.

Then, MSM(2)-AR(p) and MSH(2)-AR(p) models were estimated for the selected number of lags and their performance in tracking periods of positive and negative growth were contrasted. We observed that the autoregressive structure of the model changes when regime shifts are introduced and that in general the number of lags necessary to represent the autoregressive component of the series decreases compared with the linear AR model. Coefficients that were not statistically significant at 10% level were dropped from the model and more parsimonious Markov-switching models were estimated.

Finally, the smoother probabilities were used to determine turning points according to (3). We contrast these turning points with those obtained by MR with a classical business cycles approach, which allows us to judge the performance of Markov-switching models. The results obtained from the application of this strategy are reported below.

In Table 1, the column heads indicate the country and the type of model estimated. The estimates shown are the regime means, transition probabilities, variances, and estimated average durations for states 1 (contraction) and 2 (expansion), respectively, together with the autoregressive coefficients and the maximized log-likelihood values.

Evidence is presented relating to the existence of asymmetries in the performance of some Latin American countries over periods of expansion and contraction. First, the results show that five out of eight Latin American countries have steeper contractions than expansions (the exceptions are Bolivia, Brazil, and Colombia, from which the former two also show the greatest persistence of contractions).[4] Second, the behaviours of the variance of five out of eight countries are

[4] As mentioned above, the transition probability p_{jj} can be interpreted as a regime persistence measure in the sense that gives information about the probability of the economy con-

Table 1. Latin American Countries: Maximum Likelihood Estimates of Parameters for 2-states Markov-Switching Models, 1950-1995

Parameters	*Argentina* MSM(2)-AR(0) $\sigma_1 = \sigma_2$	*Bolivia* MSH(2)-AR(2) σ_1 ? σ_2	*Brazil* MSM(2)-AR(0) $\sigma_1 = \sigma_2$	*Chile* MSH(2)-AR(0) σ_1 ? σ_2	*Colombia* MSH(2)-AR(0) σ_1 ? σ_2	*Mexico* MSH(2)-AR(1) σ_1 ? σ_2	*Peru* MSH(2)-AR(2) σ_1 ? σ_2	*Venezuela* MSM(2)-AR(0) $\sigma_1 = \sigma_2$	*United States* MSM(2)-AR(0) $\sigma_1 = \sigma_2$
μ_1	–6.000	–2.203	–1.637	–7.044	–1.894	–4.684	–7.122	–4.308	–1.676
	(0.806)	(0.477)	(1.148)	(0.548)	(0.564)	(3.101)	(2.497)	(1.773)	(0.748)
μ_2	3.963	3.171	4.560	3.771	3.040	3.492	3.196	2.874	2.830
	(0.518)	(0.261)	(0.649)	(0.712)	(0.496)	(0.637)	(0.919)	(1.031)	(0.326)
p_{11}	0.3380	0.9126	0.7018	0.5481	0.3108	0.4546	0.2679	0.6160	0.3247
	(0.131)	(0.065)	(0.147)	(0.230)	(0.192)	(0.296)	(0.215)	(0.187)	(0.214)
p_{22}	0.6838	0.8860	0.8853	0.9132	0.8539	0.8453	0.8235	0.8164	0.8008
	(0.096)	(0.068)	(0.075)	(0.073)	(0.094)	(0.092)	(0.086)	(0.116)	(0.084)
σ_1^2	6.396	12.421	8.781	53.640	0.860	22.016	39.426	11.666	2.416
	(1.463)	(4.412)	(2.319)	(32.656)	(0.664)	(16.226)	(24.405)	(3.400)	(0.617)
σ_2^2		4.253		11.234	4.368	3.967	9.739		
		(1.460)		(3.720)	(1.523)	(1.609)	(3.896)		
ϕ_1		–0.541				0.365	0.618		
		(0.142)				(0.155)	(0.151)		
ϕ_2		–0.470					–0.262		
		(0.134)					(0.163)		
Duration 1	1.5	11.4	3.4	2.2	1.5	1.8	1.4	2.6	1.5
Duration 2	3.2	8.8	8.7	11.5	6.8	6.5	5.4	5.5	5.0
Log-likelihood	–89.4565	–74.4914	–82.4684	–92.1984	–64.5566	–74.5667	–90.8367	–90.8369	–59.8184

Standard deviations are in parenthesis.

asymmetric (the exemptions being Argentina, Brazil, and Venezuela)[5] with four out of five having higher volatility in contractions than in expansions (Colombia is the exemption). Third, seven out of eight countries have experienced contractions with lower persistence and shorter duration than expansions (only Bolivia experienced a different situation). It is also observed that the estimates for the United States reflect asymmetries between expansions and contractions: the mean growth in expansions is greater in absolute value than that for contractions, while contractions are shorter and less persistent than expansions.[6]

Comparing the results among the Latin American countries with those of the United States, we note that Peru has the largest growth rate of contractions in absolute value (7.1) and that Brazil has the lowest one (1.64), which is even lower than that of the United States (1.68). It is interesting to notice that the average growth rate of expansions for each Latin American country is larger than that of the United States (2.83), and that the highest one corresponds to Brazil as well (4.560). However, the disturbance variance for each country is higher than that of the United States (2.42); Chile has experienced the highest variance in contractions (53.64) while Venezuela has experienced the highest one in expansions (11.67). The lowest variances among the Latin American countries correspond to Colombia for contractions (0.86) and Mexico for expansions (3.97). Finally, we find that the largest duration (and persistence) of contractions corresponds to Bolivia (11.4 years) and that the lowest one corresponds to Peru (1.4 years, which is even less than that of the United Sates, 1.5 years). The longest average period of expansions is that of Chile (11.5 years) while the shortest one is that of Argentina (3.5 years). The corresponding estimation for the United States (5 years) lies between these two numbers.

In summary, compared to the United States, the Latin American countries have experienced both steeper and more volatile growth during expansions and recessions.

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tinuing in the same regime in the next period. Thus the term "persistence" does not refer to the series persistence, but to the regime persistence. Notice that by construction regime persistence and regime duration are positively related, but they are not identical concepts.

5. MSH-AR models were also estimated for these countries. However, estimated variances for each regime were very close each other and statistically insignificant. The same strategy was applied for the United States, but in that case the estimated variances were significant. However, in all these cases MSH-AR models are better in tracking positive and negative growth rates than MSH-AR models.

6. The qualitative implications of these results are consistent in general with those obtained by other authors (Hamilton, 1989; Kähler and Marnet, 1992; Goodwin, 1993; Krolzig, 1997a; and Clements and Krolzig, 1998).

The turning points derived from expression (3) for each country are presented in Table 2, where they are compared with those obtained by MR. In general, it is observed that Markov-switching models identify more turning points than the classical business cycle approach used by MR. The main reason for this is that the latter uses smoothed series as an auxiliary instrument to date turning points, which excludes short-run fluctuations. However, the turning points resulting from Markov-switching models are very close to those identified by MR for the cases of Bolivia, Brazil, Chile (in which case some turning points are shifted at most by two years), Venezuela, and the United States. Although the MS-AR models identify most turning points identified by MR also, the latter generate many more turning points in the cases of Argentina and Peru. The reason might be that these two economies exhibit many short-run fluctuations and hence short regime durations (see Table 1 and Graphs A2 and A8 in Appendix 1).[7]

The corresponding regimes are depicted in Graphs A1 to A9 in Appendix 1. In each graph, panel (a) represents the growth rate of real GDP per capita and a binary variable –obtained from the smoothed probabilities according to expression (3)– that represents the inferred recession regime. Panel (b), in turn, represents the same binary variable and the filter probability in order to show that the same regimes are implied by both sort of inferred probabilities.

3.2. Common Regime Shifts and Common Business Cycles

In this section we analyse the international nature of business cycles by applying Markov-switching vector autoregressive (MS-VAR) models. By modelling economic growth as a joint stochastic process across countries, we can identify common regime shifts and consequently common cycles. The estimation of MS-VAR models is carried out in GAUSS[8] using the Expectation Maximization (EM) algorithm presented in Hamilton (1990) and Krolzig (1997b) to maximize the corresponding likelihood function.

[7] As Goodwin (1993) has pointed out, it is not clear whether a disagreement in the identification of turning points should be taken as evidence against the methodology of classical business cycles or the Markov-switching models. Perhaps the two methods can be used in a complementary manner.

[8] The program used is a modified version of that of Engel and Hamilton (1990). Different sets of simple initial values (to indicate the existence of asymmetry in means, variances, and transition probabilities) have been used.

Table 2. Latin America Countries: Business Cycles Chronologies According to Markov Switching Models for GDP per capita, 1950-1995

Turning Point	*Argentina*	*Bolivia*	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Peru*	*Venezuela*	*United States*
Peak	1950			1953					1953
Trough	1952			*1954*	1951	1953			1954
Peak	1958				1955		1958	1957	1955
Trough	1959	1959			1958		1959	1958	1958
Peak								1959	
Trough								1960	
Peak	1961		1962		1962				
Trough	1963		1965		1963				
Peak					1964				
Trough					1965				
Peak					1966				
Trough					1967				
Peak	1974			*1972*					1973
Trough	1976			*1975*					1975
Peak	1977					1976	*1977*	1978	
Trough	1978					1977	1978		
Peak	1980	1979	1980	1981	1981	1981	*1982*		1979
Trough									1980
Peak									1981
Trough	1982		1983	1983	1983	1983	1983		1982
Peak	1984					1985	1984		
Trough	1985					*1986*	1985	1985	
Peak	1987		1987				1987	1988	1989
Trough	1990	1990					1989	1989	1991
Peak						1992	1991	1992	
Trough			1992			1993	1992		
Peak	1994	1993				1994			

Underlined years correspond to turning points identified as well by Mejía-Reyes (1999) with a classical business cycles methodology; years in italics are different at most by two years with respect to those identified by Mejía-Reyes (1999).

The regimes resulting from MS-AR models, depicted in Graphs A1 to A9 in Appendix 1, and the results reported by MR about business cycle regime synchronization suggest that a Latin American business cycle does not exist. In the graphs it can be observed that, in general, regimes do not match. However, it is still possible that common business cycles exist for some subsets of countries, as some studies have suggested (Engle and Issler, 1993; Arnaudo and Jacobo, 1997; Mejía-Reyes, 1999).

In MS-VAR models a common cycle is identified when there is synchronization in regimes for individual countries. On this basis, we have selected the countries among which there might be a common cycle according to the results obtained from the calculation of the Pearson's corrected contingency coefficient for the regimes resulting from the estimation of MS-AR models for individual countries. The computed Pearson's coefficients are presented in Appendix 2. It is interesting to observe that the results suggest associations that are at most mild.

Based on the analysis of Appendix 2, we estimate MS-VAR models for the following pairs of countries: Argentina-Brazil, Argentina-Peru, Bolivia-Venezuela, Brazil-Peru, and Chile-United States. In addition, we estimate an MS-VAR model for Argentina, Brazil, and Peru to evaluate whether common fluctuations of two countries extend to a third country.[9]

As in the analysis of univariate MS-AR models, the number of autoregressive lags was initially determined according to the Schwarz criterion in a linear VAR model. The experience in the MS-AR estimation in Section 3.1 and the remarks by Krolzig (1997a) suggest that the number of lags in such linear models is the maximum number of lags to be considered for a MS-VAR model. Because the number of lags suggested by the Schwarz criterion in linear VAR models is zero in each case, we have estimated MSH(2)-AR(0) models.

The implication of not including any autoregressive component in the MS-VAR model is that there is neither international nor inter-temporal transmission of country-specific shocks. However, it remains possible the existence of common shifts which might be explained by large contemporaneously occurring common shocks or by coincident, but independent, domestic shocks and policies. Dellas (1985) and

[9] These are the countries for which we have found stronger evidence of synchronization of business cycle regimes on the basis of Pearson's coefficient.

Canova and Dellas (1993) find that this is the case for developed countries while Krolzig (1997a) suggests that this is especially true since the oil-price shock in 1973.

It is plausible to think that international trade and foreign investment within Latin America have not acted as transmission mechanisms. Specifically, largely inwardly oriented policies and restrictions to foreign property implemented until the mid-1980s caused international transactions within the region to be very low. Latin American economies have become more market-oriented since the structural reforms undertaken in the second half of the 1980s. Thus international flows of goods and assets within the region have become important only since the late 1980s.[10] Therefore, although it may be expected that international transactions will become active transmission mechanisms of national fluctuations in the future, it is plausible to think that so far existing common cycles, if any, are due to common policies (import substitution industrialization strategies implemented from the 1950s to the late 1970s as well as restrictive stabilization policies followed during the 1980s and 1990s) and common external shocks (huge capital inflows during the late 1970s and the early 1990s, and abrupt outflows and flights of capital in 1981-1983 and in 1994).

The results of the estimation of MS-VAR models and outlines about some common contraction episodes are presented below. To evaluate how well a model fits common cycles we require the estimated parameters of the MSH(2)-VAR(0) models to be generally consistent with those obtained for the univariate MS-VAR models in the previous section. Also, to accept that there exists a common cycle we expect to observe a reasonable number of common shifts and regimes. In addition, we check how well the MS-VAR models track common shifts and regimes. Finally, we explore the possible causes of common regimes, especially of contractions, to know to what extent economic fluctuations of one country affect those of other, or to what extent they are caused by exogenous common phenomena or by similar economic policies.

In Table 3 we report the estimated parameters, namely, the estimated means vectors $\left(\mu_i \text{ for } i = 1, 2\right)$ and the variance-covariance matrices $\left(\Omega_i \text{ for } i = 1, 2\right)$ as well as the transition probabilities (p_{ii} for i = 1, 2). We chose the estimations that maximize the log-likelihood, which

[10] Some studies have suggested that interregional trade and foreign investment in Latin America have become important just since the mid-1980s and the early 1990s, respectively (see Edwards and Savastano, 1989; Edwards, 1995; CEPAL, 1993, 1998).

Table 3. Latin America: Vector Markov-Switching Models, 1951-1995

Parameter	*Argentina-Brazil*		*Argentina-Peru*		*Bolivia-Venezuela*		*Brazil-Peru*		*Chile-United States*		*Argentina-Brazil-Peru*		
μ_1	–3.257	0.873	–5.250	–2.178	–5.681	–3.561	–2.705	–6.013	–5.888	0.836	–5.617	–3.295	–7.267
	(1.190)	(0.842)	(1.196)	(1.984)	(2.873)	(2.785)	(1.096)	(2.668)	(3.157)	(1.409)	(1.619)	(1.244)	(3.422)
μ_2	4.840	4.686	4.058	3.355	1.770	1.264	4.040	3.123	4.016	2.038	1.733	3.700	2.733
	(0.480)	(0.810)	(0.658)	(0.846)	(0.567)	(0.748)	(0.542)	(0.765)	(0.637)	(0.374)	(0.760)	(0.534)	(0.718)
p_{11}	0.5829		0.3873		4.6×10^{-17}		0.6857		0.584		0.9501		
	(0.127)		(0.132)		(0.155)		(0.173)		(0.193)		(0.035)		
p_{22}	0.5672		0.6612		0.8285		0.9327		0.896		0.6506		
	(0.117)		(0.129)		(0.094)		(0.050)		(0.064)		(0.190)		
Ω_1	18.558	–0.979	10.737	–1.374	31.702	–17.469	8.287	–8.960	49.586	14.103	15.974	–6.478	–7.494
	(7.464)	(3.837)	(5.064)	(6.963)	(19.360)	(17.276)	(4.078)	(8.496)	(25.282)	(9.995)	(9.062)	(5.565)	(13.857)
	–0.979	13.993	–1.374	54.337	–17.469	39.812	–8.960	57.233	14.103	11.838	–6.478	9.390	–15.171
	(3.837)	(4.418)	(6.963)	(19.882)	(17.276)	(23.481)	(8.496)	(28.203)	(9.995)	(5.633)	(5.565)	(5.331)	(12.057)
										–7.494	–15.71	71.361	
										(13.857)	(12.057)	(40.653)	
Ω_2	4.438	–5.220	6.671	0.932	9.819	2.646	9.941	–0.670	9.952	–0.052	22.520	–0.556	5.037
	(1.417)	(1.989)	(2.524)	(2.226)	(2.616)	(2.581)	(2.412)	(2.482)	(2.918)	(1.392)	(5.087)	(2.417)	(3.506)
	–5.220	11.887	0.932	19.216	2.646	16.357	–0.670	18.510	–0.052	4.128	–0.556	11.121	1.320
	(1.989)	(3.748)	(2.226)	(5.500)	(2.581)	(4.563)	(2.482)	(4.963)	(1.392)	(1.067)	(2.417)	(2.520)	(2.422)
										5.037	1.320	20.125	
											(3.506)	(2.422)	(4.562)
Log-likelihood	–170.874		–192.357		–177.430		–179.341		–151.463		–269.146		

Standard deviations are in parenthesis.

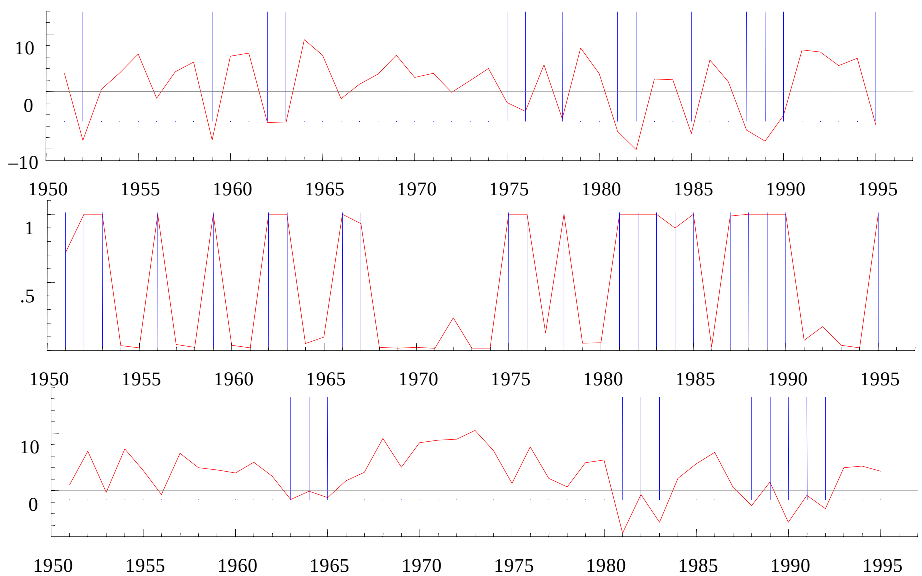
in most cases correspond to the best tracking of declines for the countries in the VAR. In Graphs 1-5 we depict the resulting common regimes from the MSH(2)-VAR(0) models. In each graph we represent the growth rates of real GDP per capita and the inferred regimes from univariate MS-AR models for the two corresponding countries in panels (a) and (c). Also the filter probability and the resulting regimes from the bivariate MSH(2)-VAR(0) model are shown in panel (b).[11]

For Argentina and Brazil, Table 3 indicates that the average growth rate in expansions is greater in magnitude than that in contractions; the estimated transition probabilities imply similar persistence and duration of both regimes, while asymmetric volatility is only found for Argentina. These results are not consistent with those found with univariate MS-AR models. These differences can be explained by the implied regimes. In Graph 1 we observe that common regimes of panel (b) resemble more the regimes of Argentina than those of Brazil. Also, it is interesting to notice that the model tracks some episodes of low growth in both countries as “contractions”, which generates a lower mean in absolute value for state 1 than that estimated by the corresponding MS-AR model for Argentina, and a positive mean for Brazil. According to panels (a) and (c), common contractions for Argentina and Brazil relate only to the early and late 1980s. It is interesting to point out that the former could be considered as an external common shock since it is mainly related to the external debt crisis. The latter, in turn, might be explained by internal efforts to control hyperinflation and by the restrictive stabilization policies implemented to overcome the situation in both countries.[12] Therefore, in a strict sense –considering only the “expansion” and “contraction” regimes for the individual countries– there is not enough evidence to conclude that there is a common cycle between Argentina and Brazil. However, a detailed observation of Graph 1 suggests that it is plausible to think that there exist “low growth” episodes in Brazil associated with contractions in

[11] As above, the binary variable representing recessions and expansions is inferred from the smoothed probability coming from the corresponding model. The binary variable and the filter probability from the MS-VAR model are presented in panel (b) to show that similar regimes are inferred from both sorts of probability.

[12] See Tanner and Sanguinetti (1997) for an analysis of the Argentinean experience over this period and Rabello de Castro and Ronci (1991) for the case of Brazil. Tanner and Sanguinetti argue that the Argentinean situation was additionally complicated by the political instability related to the expectation of probable chaos resulting from the electoral victory of the Peronist candidate Carlos Menem. Kiguel and Liviatan (1995) analyse the hyperinflation of Argentina and Brazil and conclude that the decline in Argentinean GDP during the late 1980s was mainly due to hyperinflation.

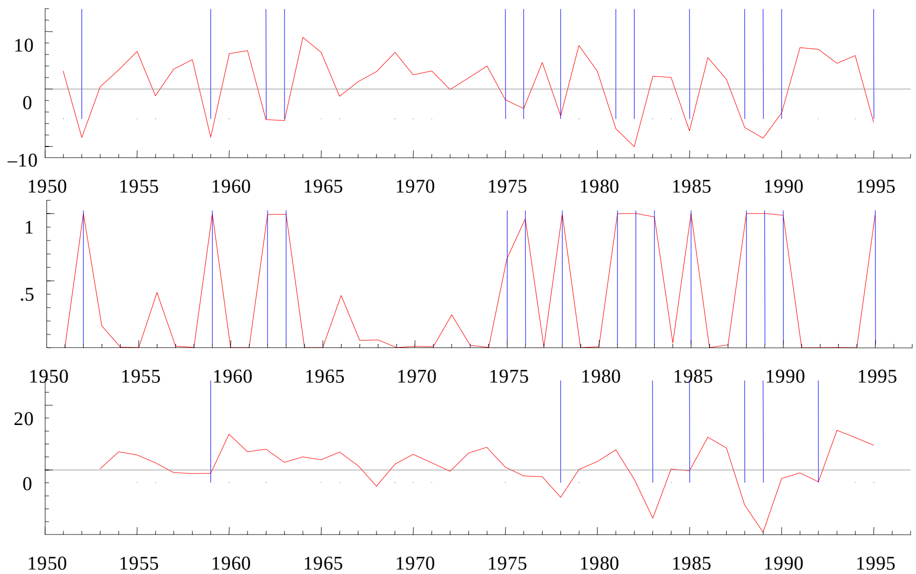
Graph 1. *a)* Argentina: Growth rates of real GDP per capita and regimes from MSM(2)-AR(0) model. *b)* Argentina-Brazil: Filter probability and regimes from MSH(2)-VAR(0) model. *c)* Brazil: Growth rates of real GDP per capita and regimes from MSM(2)-AR(0) model



Argentina. This interpretation makes our results qualitatively consistent with the findings of Engle and Issler (1993), Arnaudo and Jacobo (1997), and Mejía-Reyes (1999).

The results for Argentina and Peru suggest the existence of asymmetric behaviour of real GDP per capita over the business cycles with respect to volatility, duration and persistence, but asymmetric magnitude only for Argentina. The latter result is in clear contradiction with the evidence from the univariate model for Peru in Table 2. The implied common regimes resemble again more closely those of Argentina, but the MSH(2)-VAR(0) model improves the tracking of declines in the Peruvian economy compared with the univariate model. The declines of 1976-1978, 1982-1983, and 1988-1990 are better represented (see Graph 2), for example. However, it is important to point out that once the economic crisis started in the early 1980s, the behaviour of these economies was determined largely by the timing and nature of the implemented stabilization policies as well as by exogenous phenomena. The deepest and longest contraction periods that are common to both

Graph 2. *a)* Argentina: Growth rates of real GDP per capita and regimes from MSM(2)-AR(0) model. *b)* Argentina-Peru: Filter probability and regimes from MSH(2)-VAR(0) model. *c)* Peru: Growth rates of real GDP per capita and regimes from MSH(2)-AR(2) model

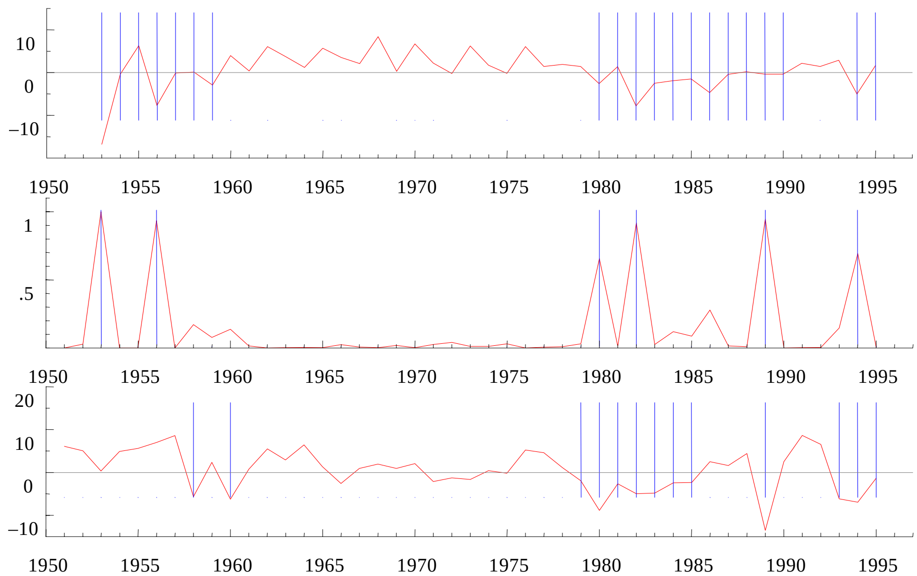


economies are those of the early and late 1980s. The latter was the result of internal issues associated with several failed stabilization efforts as well as political instability in both countries.[13] It seems that the external crisis of the early 1980s was the unique common shock. Thus, although in this case also there is some evidence about common regimes and a common cycle for Argentina and Peru, it seems that the individual business cycles are largely independent.

The estimation results for Bolivia and Venezuela presented in Table 3 exhibit significant asymmetry in all the aspects: magnitude, volatility, persistence, and duration. However, from Graph 3 it seems that

[13] Lago (1991) argues that the decline in the Peruvian output from 1988 to 1992 was associated to the desequilibria generated by the hyperinflation –caused by the over-expansive policies implemented by president Alan García during 1986-1987– as well as by the political instability derived from the attacks of the guerrilla group "Sendero Luminoso" and from the nationalization of banks and private insurance companies. See also Kiguel and Liviatan (1995) for an analysis of the hyperinflation and the stabilization policies in this period; Dancourt, Mendoza, and Vilcapoma (1997) for a characterization of the fluctuations of the Peruvian economy; and Hamann and Paredes (1991) for an analysis of the Peruvian growth strategies.

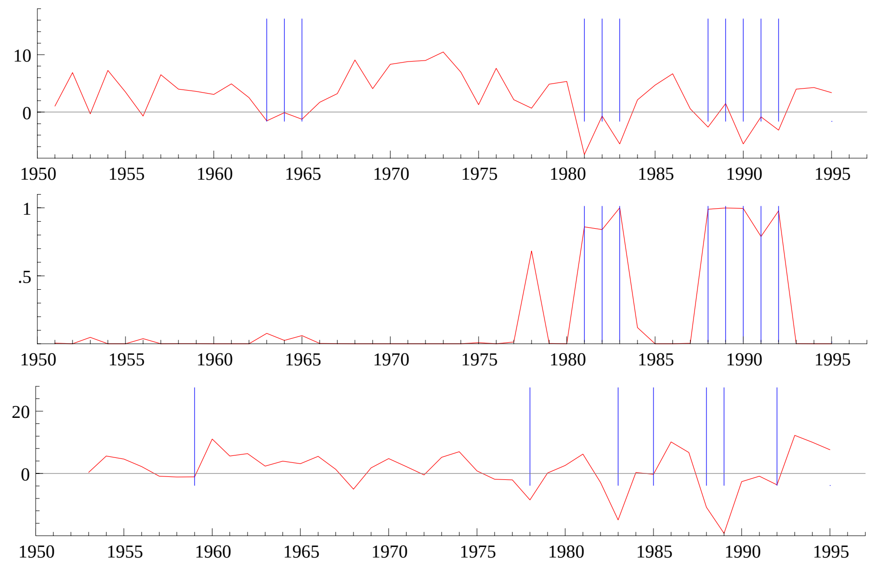
Graph 3. *a)* Bolivia: Growth rates of real GDP per capita and regimes from MSH(2)-AR(2) model. *b)* Bolivia-Venezuela: Filter probability and regimes from MSH(2)-VAR(0) model. *c)* Venezuela: Growth rates of real GDP per capita and regimes from MSM(2)-AR(0) model



the identified regimes do not track adequately the contractions in real GDP per capita. From the regimes for each country it is apparent that both economies were in contraction during the first half of the 1980s and the last two years of the sample. However, the value of the probability of continuing in contraction is almost equal to zero, which implies that the identified contraction regimes last only for one year. Thus, although the regimes identified by the MS-VAR model reflect common years of contraction, they do not give a complete representation of individual regimes. In particular, in panels (a) and (c) we observe that the univariate results imply that Bolivia was in contraction during all years of the 1980s and Venezuela was in contraction during seven years (1979-1985). However, the bivariate model does not identify any long lasting contraction. Therefore, the existence of a common cycle is weak for these two countries.[14]

[14] Sturzenegger (1995) considers that the long recession period of the 1980s in Bolivia was a result of the collapse in tin prices in 1980 –which provoked that the crisis in Bolivia started before the external debt crisis–, the external debt crisis and the difficulties to reduce inflation. Actu-

Graph 4. *a)* Brazil: Growth rates of real GDP per capita and regimes from MSM(2)-AR(0) model. *b)* Brazil-Peru: Filter probability and regimes from MSH(2)-VAR(0) model. *c)* Peru: Growth rates of real GDP per capita and regimes from MSH(2)-AR(2) model



The relationship between the regimes of Brazil and Peru depicted in Graph 4 suggest that there exists a common business cycle. The estimated parameters shown in Table 3 are consistent with those obtained from univariate MS-AR models: the Brazilian economy's growth in expansions is greater than falls in contractions, and the volatility of the economy is similar within each state. On the contrary, in the Peruvian economy falls in contractions are greater than upturns in expansions (in absolute value). Furthermore, the economy is more volatile in contractions than in expansions. The estimated transition probabilities suggest that contractions last less long than expansions, but the value of p_{11} implies that they last for some years. This is rep-

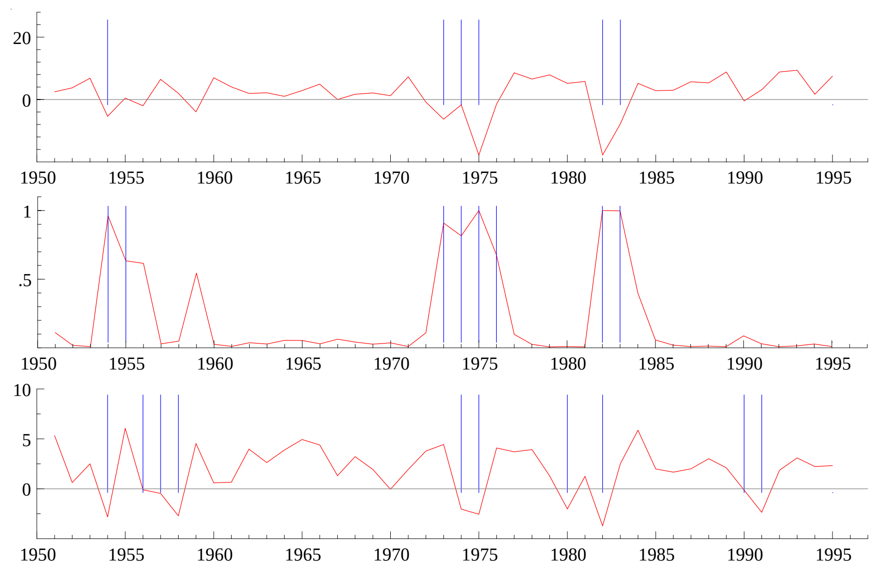
ally, several stabilization programs failed to take inflation down, but their recessionary effects lasted for a long time. In Venezuela, the contraction in the early 1980s was a result of the external debt crisis and the fall in the international price of oil. Whilst the contraction of 1989 is related to the economic uncertainty derived from failed stabilization policies implemented to control the economic chaos provoked by the new fall of oil prices in 1985-1986 (Hausmann, 1990). The recent contraction in Venezuela has been associated with an increase in inflation and the effects of stabilization policies (Little and Herrera, 1995).

resented in Graph 4. It can be observed that the Peruvian contractions of 1982-1983 and 1988-1992 are better represented by the regimes of the bivariate model than by the regimes of the univariate one. Even though the first recession began in 1981 for Brazil, these two recessions were common for both countries and this fact implies the existence of common shifts in regimes and common cycles. However, as mentioned above, it seems that these recessions were largely caused by exogenous shocks and by internal political and economic instability, respectively. Therefore, although the shifts and the cycles are common to both countries, it does seem that the causes of the cycles for each country are largely independent. Similar industrialization and restrictive stabilization policies as well as external shocks might explain common shifts and regimes.

The results for Chile and the United States suggest the existence of some evidence about a common cycle for both countries. In general, the estimated parameters are similar to those obtained from the univariate models, especially to those for Chile. They differ with respect to the United States' ones essentially in the difference of volatility for expansions and contractions. In Graph 5 we note that common regimes in panel (b) resemble more those of Chile: the contraction of 1973-1976 and the huge fall of 1982-1983 are fully represented. These contractions and that of 1954 are the three common contraction regimes represented by the bivariate model. The other US recessions are not tracked. However, even though some recent recessions are common to Chile and the United States, their causes are different. Recent recessions in Chile have been associated with dramatic changes in economic policies and political instability (1972-1976) and with the external debt crisis (1982-1983).[15] In the United States these recessions have been associated with the oil-shock (1974-1975) and the world recession (1981-1982). Actually, the effects of the US economy on the Chilean economy were transmitted abruptly only during the external debt crisis in 1982, and not only to Chile but to the whole region.

[15] The large decline in production in the Chilean growth over the period 1972-1976 was caused by the economic instability associated to the expansionary policy of socialist president Salvador Allende, which ended in high inflation and huge fiscal deficit. On the other hand, investment and savings fell because of the uncertainty provoked by the nationalization of mining, banking, and agricultural sector, as well as most of the manufacturing sectors. Political instability ended in a military coup that was followed by an ambitious reform to change the economy into a world integrated marked-oriented society. Initially, the stabilization program reduced demand even more and caused additional recession. The recession of 1982-1983 was largely related to the external debt crisis and to the worsening in terms of trade, especially to the fall in the copper price (Edwards and Cox Edwards, 1987).

Graph 5. *a)* Chile: Growth rates of real GDP per capita and regimes from MSH(2)-AR(0) model. *b)* Chile-United States: Filter probability and regimes from MSH(2)-VAR(0) model. *c)* United States: Growth rates of real GDP per capita and regimes from MSM(2)-AR(0) model

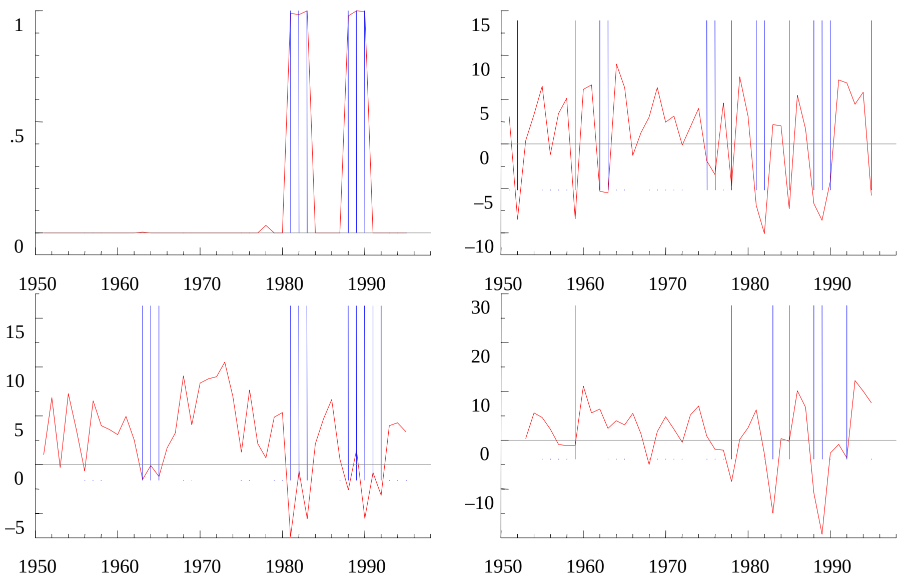


Thus, it seems again that common shifts and cycles have different causes in the individual countries, and, in that sense, business cycles are largely independent.

Finally, we have estimated an MSH(2)-VAR(0) model for Argentina, Brazil, and Peru because we have found some evidence of possible common cycles among them. The results are presented in Table 3. There is evidence of asymmetry in mean with contractions being larger in magnitude than expansions, except for Brazil. According to the estimates of the transition probabilities, we also find that expansions last for more years than contractions. The results are consistent with the estimations of the univariate MS-AR models with respect to existence of asymmetric volatility: there is clear evidence of such asymmetry only in the case of Peru.[16] Thus, in general, our estimations are consistent with those obtained with the univariate models.

[16] We found in Section 3.1 that univariate MS-AR models with common variance for Argentina and Brazil fitted best their cyclical fluctuations.

Graph 6. *a)* Argentina-Brazil-Peru: Filter probability and regimes from MSH(2)-VAR(0) model. *b)* Argentina: Growth rates of real GDP per capita and regimes from MSM(2)-AR(0) model. *c)* Brazil: Growth rates of real GDP per capita and regimes from MSH(2)-AR(0) model. *d)* Peru: Growth rates of real GDP per capita and regimes from MSH(2)-AR(2) model



Estimated regimes are depicted in Graph 6. In panel (a) we represent estimated regimes from a MSH(2)-VAR(0) model while in panels (b), (c), and (d) we present the regimes generated by univariate MS-AR models and the growth rates of real GDP per capita of Argentina, Brazil, and Peru, respectively. It appears that there are two common contractions: in the early and in the late 1980s. We have explained above that the causes of these two contractions are the external debt crisis, common to the region, and domestic economic and political instability, respectively. In addition, the model does not track other contraction episodes because they are not common to this set of countries. Thus, it seems that there are only a few common regimes, with individual business cycles of each country largely independent, and that they are consequence of similar economic policies and common external shocks.

4. Final Considerations

The evidence reported has documented the existence of asymmetric behaviour of business cycle regimes within each country with respect to magnitude, volatility, and/or regime persistence, which might suggest that Latin American economies work differently in expansions and contractions. In general, our results imply that recessions are steeper in absolute magnitude, less persistent, and more volatile than expansions.

Also, it appears that there is some evidence suggesting the existence of a pattern for the Latin American countries, which is qualitatively different to that of the United States. In particular, our results concerning the experience of the United States are consistent with those obtained by other authors for that country and for other developed economies. For example, by using Markov-switching models for developed economies, Kähler and Marnet (1992), Goodwin (1993), and Krolzig (1997a) find evidence of asymmetry in duration and regime persistence between contractions and expansions, while Kähler and Marnet (1992) also find evidence of asymmetric behaviour for the variances. However, their results show neither variances as high as those found here for Latin American countries nor regime persistence as high as that in Latin America. Furthermore, even if they find evidence of asymmetry in magnitude between expansions and contractions, they show that in absolute value the average growth rate in contractions is less than the average growth rate in expansions.[17] Thus, our results support the idea that business cycles in Latin America may be different to business cycles in the United States and (by comparison with the results of the studies mentioned above) other developed countries. However, further work needs to be done on this subject.

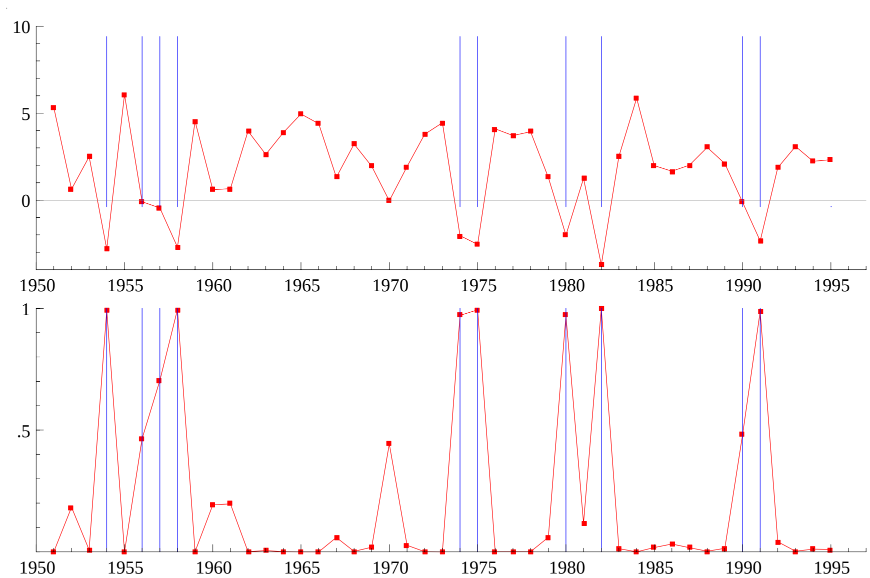
From an international perspective, our Markov-switching vector autoregressive models suggest that, although there is not a common Latin American cycle, there exists some evidence about common regime shifts and common cycles for some countries. We estimated bivariate Markov-switching models for Argentina-Brazil, Argentina-Peru, Bolivia-Venezuela, Brazil-Peru, and Chile-United States and a

[17] This conclusion does not change even if we take into account the evidence reported by Artis, Kontolemis, and Osborn (1997), who find that real industrial production declines by an overall average of 0.9% a month in contractions and increases by an overall average of 0.7% a month in expansions. Furthermore, they find that a similar pattern is present in nine out of twelve developed countries.

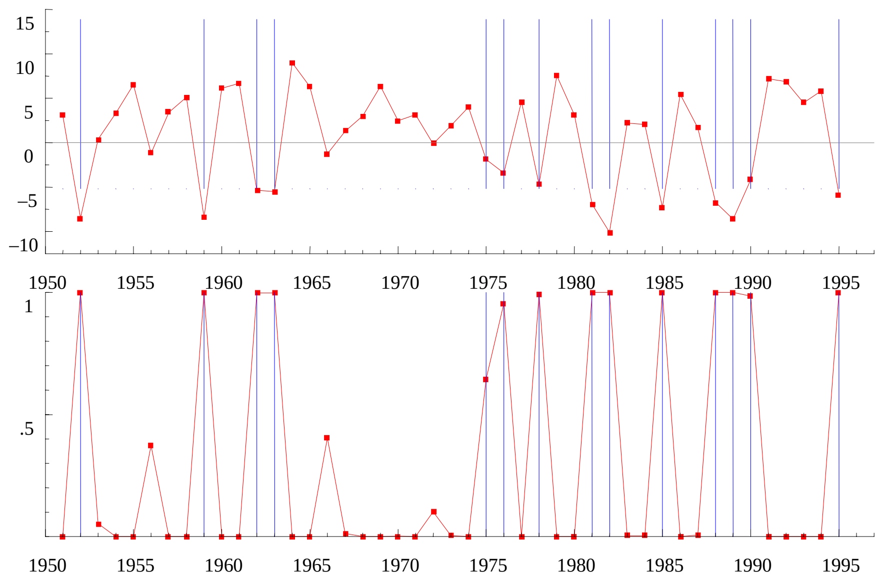
three-country MAS-VAR model for Argentina, Brazil, and Peru. We find evidence of common cycles only for Brazil-Peru and Chile-United States. However, when we look for the explanation of contraction periods, we observe that their causes appear to be different across countries. During the 1980s and the 1990s, in most cases contractions were related to external shocks and to inflation and hyperinflation processes as well as the recessionary effects of stabilization policies pursued by the separate countries. On the other hand, the synchronized long period of expansion during the 1950s and most of the 1960s was a consequence of similar industrialization strategies followed in Latin American countries. It is important to point out that the inward-oriented nature of these policies makes it hard to think about them as co-ordinated policies. Therefore, it seems that individual business cycles are largely independent in Latin America. It might be expected, however, that after the recent structural reforms, especially the liberalization of trade and financial markets, a major integration and correlation among business cycles and business cycles regimes may occur within the region in the future.

Appendix 1. Growth rates of real GDP per capita, filter probability and regimes from Markov-switching autoregressive models

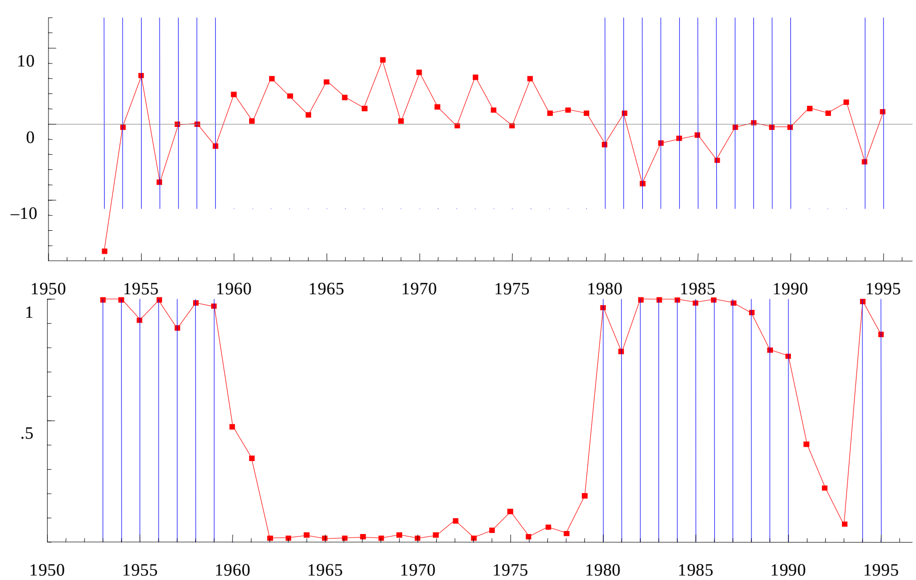
Graph A1. United States. MSM(2)-AR(0) model. *a)* Growth rates of real GDP per capita and regimes. *b)* Filter probability and regimes



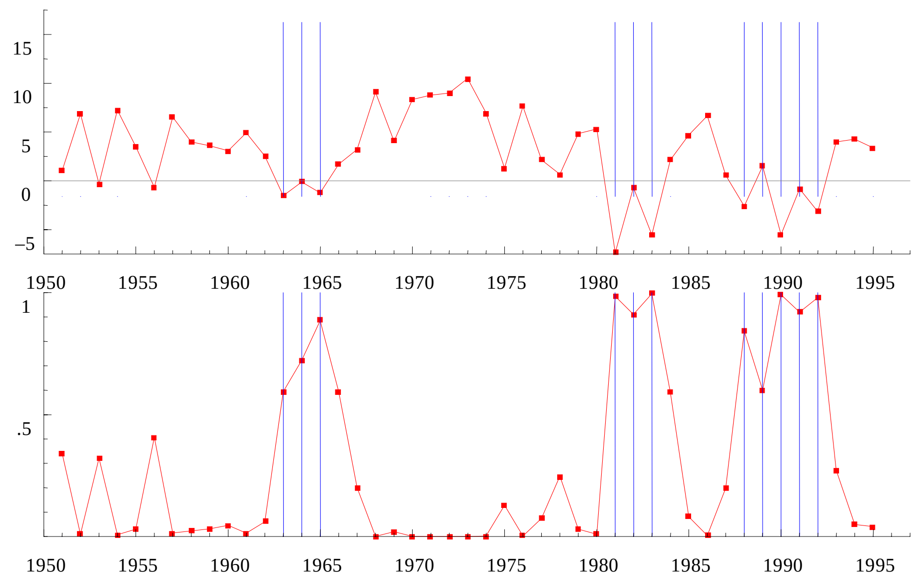
Graph A2. Argentina. MSM(2)-AR(0) model. *a)* Growth rates of real GDP per capita and regimes. *b)* Filter probability and regimes



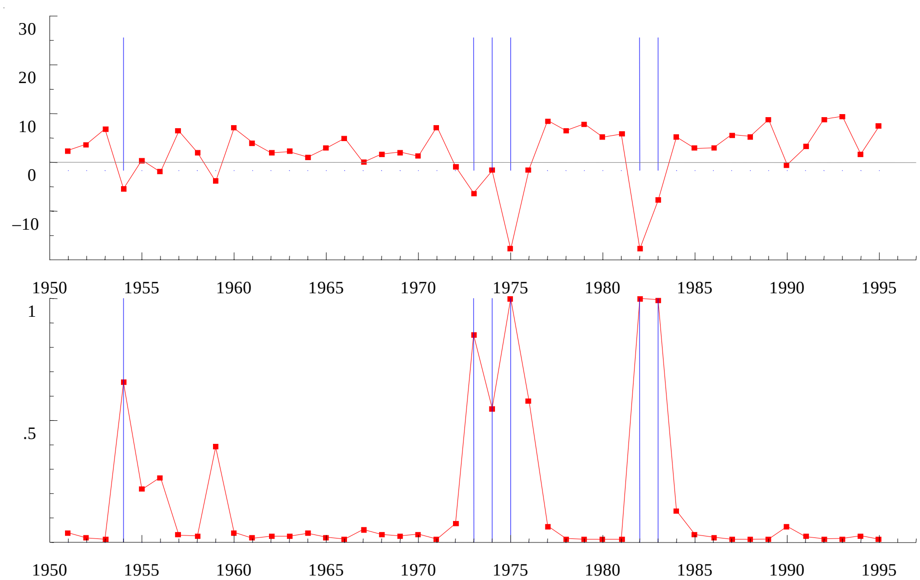
Graph A3. Bolivia. MSH(2)-AR(2) model. *a)* Growth rates of real GDP per capita and regimes. *b)* Filter probability and regimes



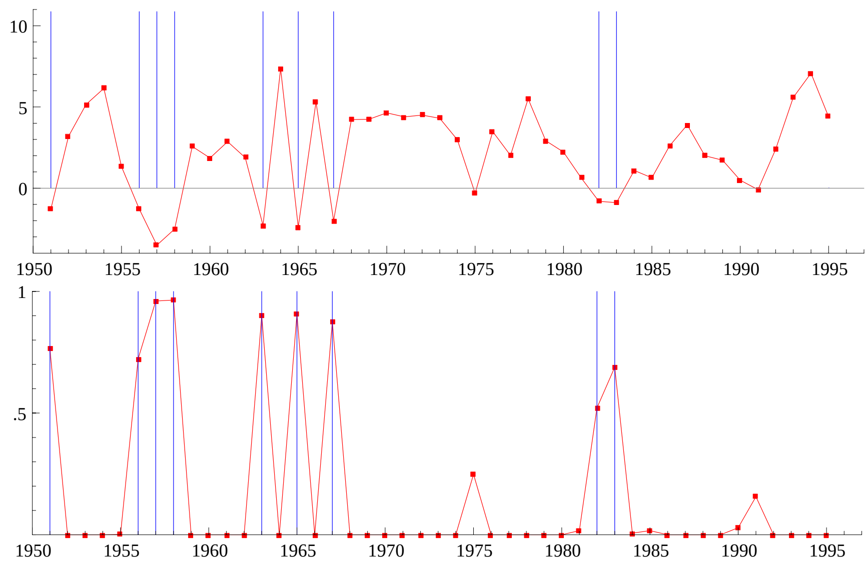
Graph A4. Brazil. MSM(2)-AR(0) model. *a)* Growth rates of real GDP per capita and regimes. *b)* Filter probability and regimes



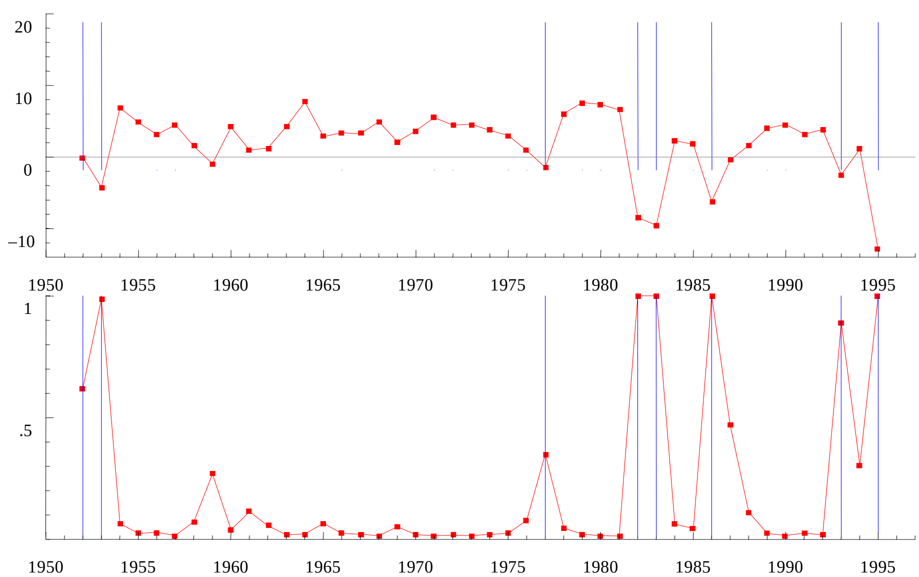
Graph A5. Chile. MSH(2)-AR(0) model. *a)* Growth rates of real GDP per capita and regimes. *b)* Filter probability and regimes



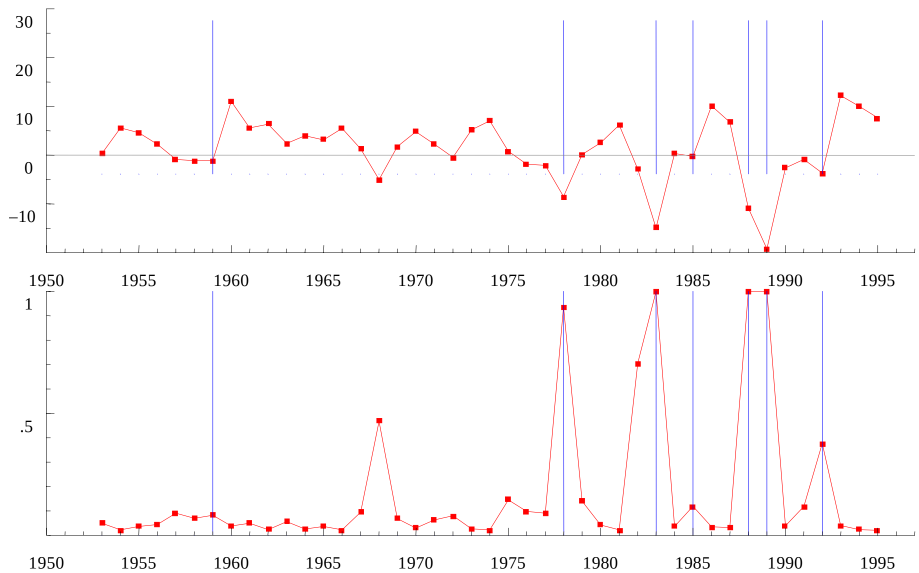
Graph A6. Colombia. MSH(2)-AR(0) model. *a)* Growth rates of real GDP per capita and regimes. *b)* Filter probability and regimes



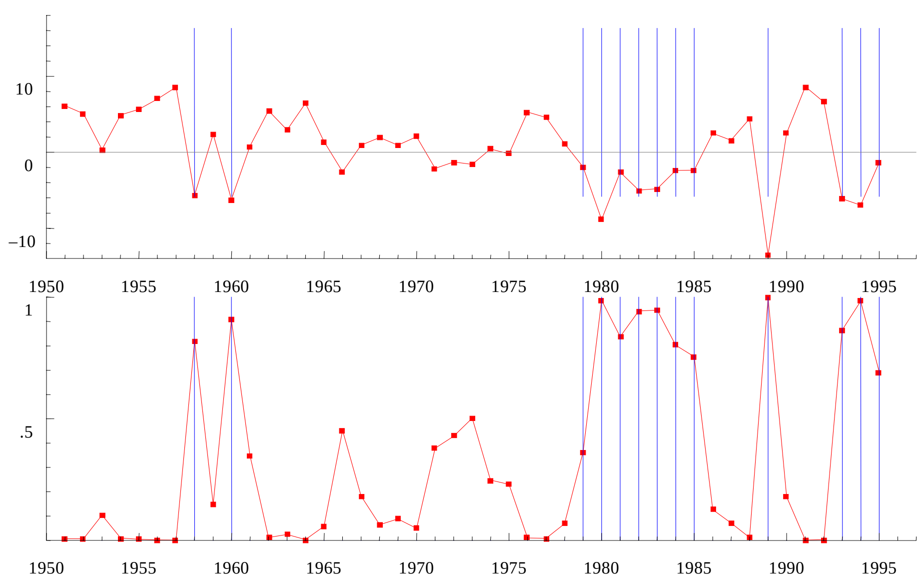
Graph A7. Mexico. MSH(2)-AR(1) model. *a)* Growth rates of real GDP per capita and regimes. *b)* Filter probability and regimes



Graph A8. Peru. MSH(2)-AR(2) model. *a)* Growth rates of real GDP per capita and regimes. *b)* Filter probability and regimes



Graph A9. Venezuela. MSM(2)-AR(0) model. *a)* Growth rates of real GDP per capita and regimes. *b)* Filter probability and regimes



Appendix 2. Pearson's corrected contingency coefficient for regimes derived from Markov-switching autoregressive models

We follow the methodology suggested by Artis, Kontolemis, and Osborn (1997) (AKO hereafter). The regimes determined according to univariate MS-AR models in Section 2 are used to create a binary time series variable for each country, denoting years during expansion by zeros and recessions by ones. Then the Pearson's corrected contingency coefficient was calculated for each pair of countries.

For the subject analysed in this study, independence implies that there is no contemporaneous relationship between the business cycle regimes (expansion/recession) for the two countries. At the other extreme, complete dependence indicates that the two countries are in the same regime for every time period and hence have identical business cycle turning point dates. The former results in a corrected Person's coefficient of zero and the latter one of 100.

We follow the same criteria as in Mejía-Reyes (1999) and consider the association between the regimes of country *i* and country *j* as "mild" and "strong" when the Pearson's corrected contingency coefficient lies in the intervals 40-60 and 61-100, respectively. The computed coefficients are reported in Table A1.

MS-VAR models have been estimated for the cases when the association between countries is at least a mild one. Four pairs of countries satisfy this requirement, but this is not satisfied for third countries. For example, there are mild associations between Peru and Argentina and Peru and Brazil, but the association between Argentina and Brazil is low. We estimate a MS-VAR model for Argentina and Brazil as well because of three reasons. First, some authors (Engle and Issler, 1993, and Arnaudo and Jacobo, 1997) have found strong links between these two economies. Second, the Pearson's coefficient is very close to our lower limit for mild associations (39.1). Third, the Pearson's coefficient estimations reported by MR suggest a strong association (67.3). Thus, we estimate MS-VAR models for the following pairs of countries: Argentina-Brazil, Argentina-Peru, Bolivia-Venezuela, Brazil-Peru, and Chile-United States. In addition, we estimate a MS-VAR model for Argentina, Brazil, and Peru as these are the countries for which we have found stronger evidence of synchronization of business cycle regimes.

Table A1. Latin America: Pearson's Corrected Contingency Coefficient (States Determined by Markov-Switching Models), 1951-1995

	Argentina	*Bolivia*	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Peru*	*Venezuela*	*United States*
Argentina	—	28.2	39.1	2.7	13.5	8.1	52.0	14.2	1.8
Bolivia	28.2	—	13.3	4.0	21.4	30.4	30.4	52.7	35.5
Brazil	39.1	13.3	—	11.4	32.0	0.0	43.0	13.2	9.7
Chile	2.7	4.0	11.4	—	18.3	21.8	0.6	6.9	54.7
Colombia	13.5	21.4	32.0	18.3	—	11.7	6.9	6.9	36.5
Mexico	8.1	30.4	0.0	21.8	11.7	—	3.4	29.2	16.2
Peru	52.0	30.4	43.0	0.6	6.9	3.4	—	17.0	33.4
Venezuela	14.2	52.7	13.2	5.4	6.9	29.2	17.0	—	1.8
United States	1.8	35.5	9.7	54.7	36.5	16.2	33.4	1.8	—

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