NBER WORKING PAPER SERIES

REAL BUSINESS CYCLES IN EMERGING COUNTRIES?

Javier Garcia-Cicco
Roberto Pancrazi
Martin Uribe

Working Paper 12629
http://www.nber.org/papers/w12629

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
October 2006

We are grateful to Stephanie Schmitt-Grohe and Vivian Yue for comments and to Alejandro Gay for sharing his data on private consumption shares. The views expressed herein are those of the author(s) and do not necessarily reflect the views of the National Bureau of Economic Research.

Real Business Cycles in Emerging Countries?
Javier Garcia-Cicco, Roberto Pancrazi, and Martin Uribe
NBER Working Paper No. 12629
October 2006
JEL No. F41

ABSTRACT

This paper investigates the hypothesis that an RBC model driven by permanent and transitory productivity shocks accounts well for business cycles in emerging markets. Existing studies that make this claim use short time series to estimate the parameters of the underlying driving forces. This practice is problematic, particularly because a central goal in this literature is to ascertain the role of permanent shocks to productivity. One contribution of the present study is to use a data set consisting of almost a century of aggregate data from Argentina. We conduct a GMM estimation of the parameters of a small open economy RBC model. We find that the RBC model does a poor job at explaining the Argentine business cycle. The difficulties of the model are most evident along five dimensions: (a) the RBC model counterfactually predicts that consumption growth is less volatile than output growth. (b) The volatility of the trade balance-to-output ratio implied by the RBC model is four times as large as its empirical counterpart. (c) The volatility of investment growth is half as large in the model as it is in the data. (d) The RBC model predicts the wrong sign for the autocorrelation function of output growth. (e) A robust prediction of the model is that the trade-balance share in output is a near random walk, with an autocorrelation function close to unity, whereas in the data, the highest autocorrelation coefficient of this variable takes place at the first order and is less than 0.6, with higher-order autocorrelations converging quickly to zero.

Javier Garcia-Cicco
Duke University
Department of Economics
213 Social Science Building
Box 90097
Durham, NC 27708
jg55@duke.edu

Roberto Pancrazi
Department of Economics
Duke University
Durham, NC 27708
rp21@duke.edu

Martin Uribe
Department of Economics
Duke University
Durham, NC 27708-0097
and NBER
uribe@duke.edu

1 Introduction

A central characteristic of business cycles in developed countries is their remarkable dampening after the second world war. This phenomenon is often attributed to improved policy management. Policymakers and policy institutions are generally credited for avoiding large economic depressions like the one that took place in the interwar period. By contrast, business cycles in many emerging countries display no signs of smoothing in the past fifty years. Large swings in aggregate activity are as likely to occur now as they were a century ago. For instance, following the debt crisis of the 1980s most countries in Latin America underwent output contractions of enormous dimensions, in many cases comparable to the one that took place during the U.S. Great Depression. Not surprisingly, misplaced fiscal and monetary policies have been blamed for the failure to achieve and maintain aggregate stability in the region.

Recently, however, a number of studies have departed from the mainstream view that in order to understand economic fluctuations in emerging markets models must take explicitly into account the role of money, fiscal deficits, and other policy variables. This line of research argues that business cycles in emerging countries can be explained well using a frictionless neoclassical model driven solely by appropriately estimated shocks to total factor productivity. Kydland and Zarazaga (2002), for instance, find that the RBC model can replicate satisfactorily the 'lost decade' of the 1980s in Argentina. More recently Aguiar and Gopinath (2004) have suggested that an RBC model driven by permanent shocks to productivity can explain most of the qualitative and quantitative differences in business cycles observed in developed and developing countries.

In this paper, we undertake an investigation of the hypothesis that an RBC model driven by a combination of permanent and transitory shocks to total factor productivity can account satisfactorily for the observed aggregate dynamics in developing countries. This investigation is motivated by what we believe is an important drawback of the existing studies advocating the ability of the RBC model to explain business cycles in developing countries. Namely, the use of short samples both for characterization of observed business cycles and for estimation of the parameters of the theoretical model. Aguiar and Gopinath (2004), for instance use Mexican data starting in the early 1980s to estimate a productivity shock process containing a unit root. By nature, driving forces with unit roots are associated with low frequency cycles. As a result, short time series are particularly ill-suited to estimate this type of uncertainty.

The main departure of the present study from the existing body of work is the use of long time series. We build a data set covering almost a century of Argentine business cycles.

The frequency of our data is annual and covers the period 1913-2005. Although our dataset contains as many observations as two and a half decades of quarterly data—the typical length and frequency of time series used in the related studies—it is more informative about the nature of the underlying stochastic process for total factor productivity. For it is well known that in estimating the stochastic process of macroeconomic time series containing unit roots, given the number of observations the length of the sample period is more important than the frequency of observations.

The theoretical framework that we use as the basis of our analysis is a standard neoclassical open economy model. We allow the production technology to be shifted by permanent and temporary shocks to productivity and estimate the structural parameters of the driving forces by GMM. In comparing the implications of the model to the data, we emphasize five dimensions: the volatility of consumption growth relative to the volatility of output growth, the volatility of the trade balance-to-output share, the volatility of investment growth, the autocorrelation function of output growth, and the autocorrelation function of the trade balance-to-output ratio. These five statistics are intimately related to one another in the RBC model. Consider, for instance, the volatility of consumption growth relative to that of output growth. In the data, consumption growth is 43 percent more volatile than output growth. In order for a model driven by productivity shocks to be able to explain this fact, it must be the case that the typical productivity shock produces an increasing pattern in future expected output. For in this case, consumption-smoothing household will find it optimal to increase consumption by more than the increase in current income. This excess of spending over income, or trade deficit, is financed by borrowing against future expected income. It follows that high volatility in consumption is likely to be associated with a volatile trade balance, serially correlated output growth rates, and serially correlated trade deficits.

We find that the RBC model fails dramatically along the five dimensions described above. Specifically, the estimated RBC model is unable to induce higher volatility in consumption growth than in output growth. Second, the volatility of the trade-balance share in output implied by the RBC model is four times as large as its empirical counterpart. Third, the observed volatility of investment is twice as large as the one predicted by the RBC model. Fourth, the model misses the sign of all of the autocorrelation coefficients of output growth up to fourth order. Finally, the model predicts that the trade balance-to-output ratio is a near random walk, with an autocorrelation function close to unity up to fourth order. By contrast, in the data the highest autocorrelation coefficient of this variable takes place at the first order and is less than 0.6, with higher-order autocorrelations converging quickly to zero.

The remainder of the paper is organized in six sections. Section 2 presents a number of stylized facts associated with the Argentine business cycle over the period 1913-2005. Section 3 lays out the theoretical model. Section 4 discusses the estimation of the structural parameters of the model. Section 5 contains the main results of the paper. It compares the predictions of the model with the data. Section 6 presents a robustness analysis. One of the sensitivity tests conducted in this section consists in estimating the model using a recently compiled set of Argentine data covering the period 1865-2004. Section 7 concludes.

2 Business Cycles in Argentina: 1913-2005

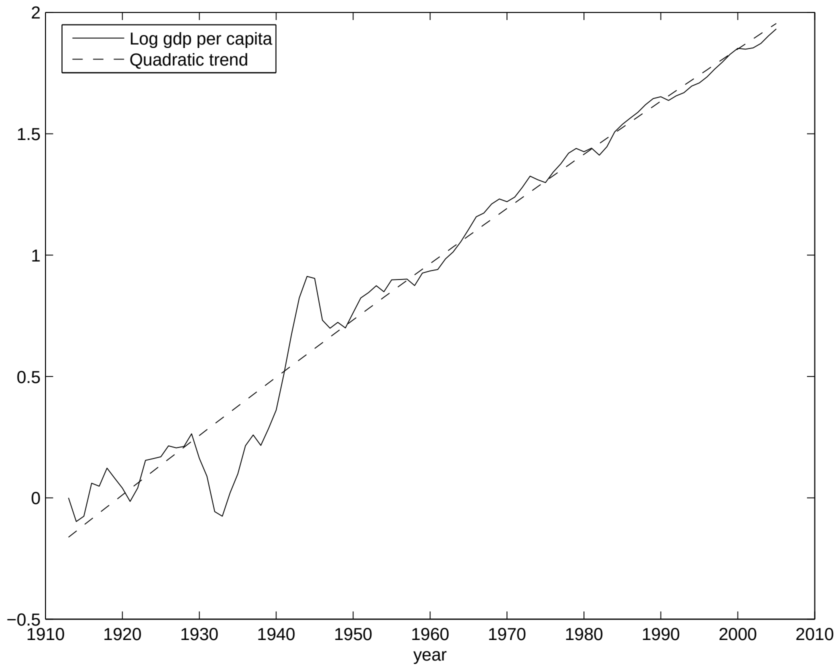
Because in the present study we entertain the hypothesis that business cycles in emerging countries are driven in part by permanent shocks to productivity, we are particularly interested in empirical evidence spanning many years. Accordingly, we put together a data set containing almost one century of aggregate data from Argentina. Our sample ranges from 1913 to 2005. Analyzing the pre- and postwar periods jointly represents a departure from the usual practice in studies focused on developed countries. Typically, these studies concentrate either on the prewar period—often emphasizing the Great Depression years—or on the postwar period—as do most of the many papers spurred by the work of Kydland and Prescott (1982) and King, Plosser, and Rebelo (1988). And there is a good reason for separating the pre- and postwar periods when examining developed-country data. For the volatility of business cycles in industrialized countries fell sharply in the second half of the twentieth century. In sharp contrast, in emerging countries business cycles do not appear to dampen after the second world war. This fact is clearly illustrated by figure 1, which shows with a solid line the logarithm of GDP per capita between 1913 and 2005 and with a broken line the associated quadratic trend for the United States (panel a) and Argentina (panel b). In the United States, the first half of the twentieth century is dominated by the Great Depression and appears as highly volatile. By comparison, the half century following the end of World War II appears as fairly calm, with output evolving smoothly around its long-run trend.[1] On the other hand, in Argentina output fluctuations appear equally volatile in the prewar period as in the postwar period.[2]

[1]A similar pattern emerges for each of the remaining G7 countries. See Basu and Taylor 1999, table A3.

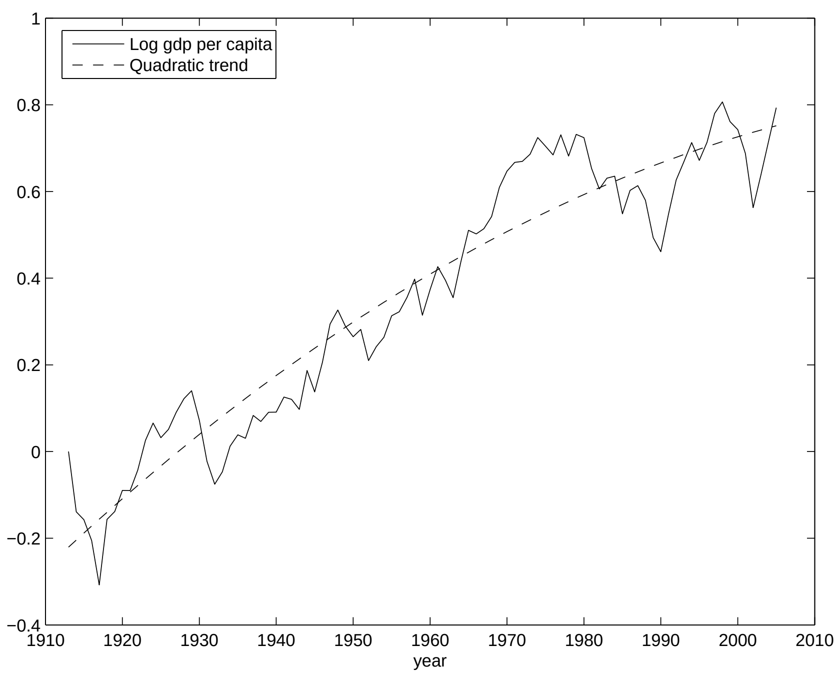
[2]Basu and Taylor (1999) also find no differences in output volatility in Argentina in the interwar and postwar eras (see their table A3). Using data from Argentina for the period 1884 to 1990, Sturzenegger and Moya (2003) find that business cycles in the pre world war II period were more volatile than in the postwar period. This different result is due to the fact that their sample does not include the years 1991-2005, which are among the most volatile of the postwar era and that their sample includes the period 1884-1912, which was particularly volatile (see Basu and Taylor, 1999, table A3). Later in section 6.1 we work with a data set of Argentine data ranging from 1865 to 2004 and find that the pre-world-war-one period contributed substantially to the observed aggregate volatility, particularly in investment growth.

Figure 1: Output Per Capita in Argentina and the United States: 1913-2005

(a) United States



(b) Argentina



The message conveyed by figure 1 is confirmed by table 1. Over the whole sample, 1913-

Table 1: Standard Deviation of Per-Capita Output Growth in Argentina and the United States

Period	Argentina	United States
1913-2005	5.4	5.4
1913-1945	5.9	8.0
1946-2005	5.2	3.4

Note: In percentage points.

2005, the United States and Argentina display similar volatilities in per-capita output growth of 5.4 percent. However, in the United States the volatility of output growth falls significantly from 8 percent to less than 4 percent between the prewar and postwar periods. In Argentina the volatility of output growth falls insignificantly from 5.9 in the earlier subperiod to 5.2 percent after the war.[3] It is this fact that motivates us to pursue a joint analysis of the Argentine pre- and postwar periods.

Table 2 displays a number of statistics from Argentina over the period 1913-2005.[4] Notably, unlike in developed countries, per-capita consumption growth in Argentina is significantly more volatile than per-capita output growth. Others have documented this fact for Argentina and other emerging countries using postwar data (Neumeyer and Perri, 2006, Aguiar and Gopinath, 2004, and Uribe, 2006). Our contribution here is to show that the high volatility of consumption relative to output remains present after augmenting the sample to include the first half of the twentieth century. Gross investment growth is enormously volatile. Its standard deviation is about four times as large as that of output growth. At the same time, the trade balance-to-output ratio is about as volatile as output growth. The observed correlation between the trade balance-to-output ratio and output growth is negative but insignificantly different from zero. By contrast the domestic components of aggregate demand, private consumption growth and investment growth, are significantly negatively correlated with the trade balance. Finally, both output growth and consumption growth are positively serially correlated, but these correlations are small and insignificant, while investment growth and the trade-balance share display significant serial correlation.

[3]Modeling the conditional volatility of the cyclical component of output per capita as a GARCH(1,1) process and applying Chow and QLR tests, we found no structural breaks in volatility over the sample. The same result obtains under alternative ways of detrending the data, such as HP filtering or taking growth rates. By contrast, the volatility of U.S. per-capita output growth presents a significant decline after the second war.

[4]The discrepancy between the standard deviation of output growth shown on tables 1 and 2 is due to the fact that the second moments shown in table 2 were estimated jointly with the autocorrelation function of order four shown in figure 2 below, which entails a loss of 4 degrees of freedom.

Table 2: Argentina 1913-2005: Summary Statistics

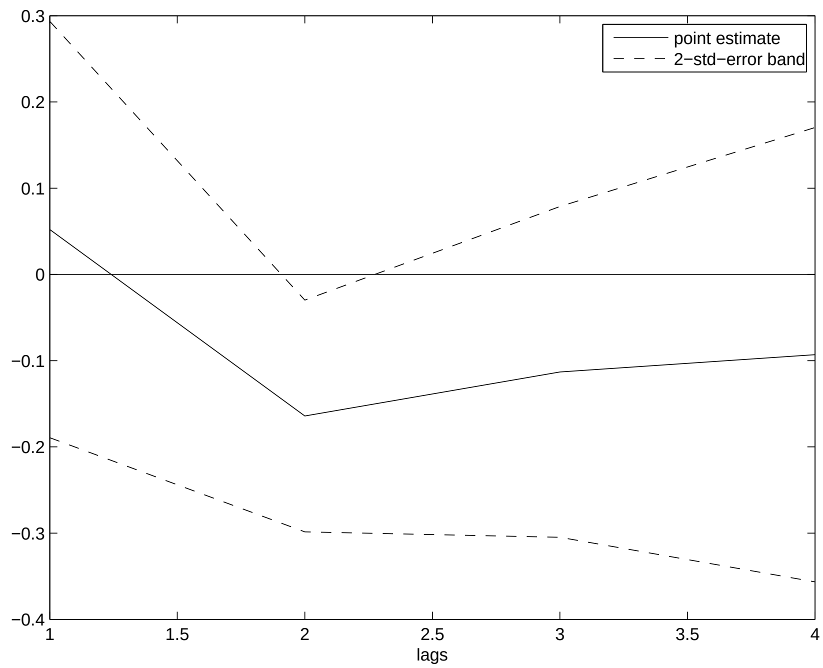
Statistic	g^Y	g^C	g^I	tby
Standard Deviation	5.1	7.3	19.0	4.9
	(0.4)	(0.7)	(1.8)	(0.7)
Correlation with g^Y	1.00	0.77	0.66	-0.07
	(0.00)	(0.07)	(0.12)	(0.09)
Correlation with tby	-0.076	-0.30	-0.21	1.00
	(0.09)	(0.06)	(0.09)	(0.00)
Serial Correlation	0.05	0.04	0.21	0.58
	(0.12)	(0.09)	(0.10)	(0.07)

Note: g^Y, g^C, and g^I denote the growth rates of output per capita, consumption per capita, and investment per capita, respectively, and tby denotes the trade balance-to-output ratio. Standard deviations are reported in percentage points. Standard errors are shown in parenthesis.

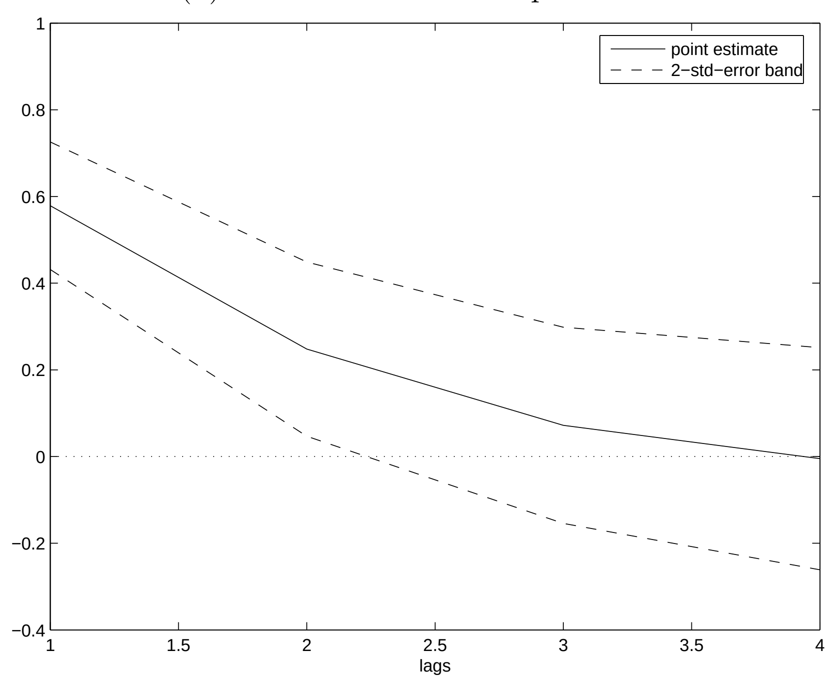
In evaluating the hypothesis that business cycles in emerging countries are driven by stochastic variations in the rate of technical progress, it is of particular interest to examine the autocorrelation function of output growth. The reason is that under that hypothesis, a productivity shock produces a deterioration in the trade balance and large increases in consumption (as is required for consumption growth to be more volatile than output growth) if current increases in the level of output are accompanied by further increases in output. For, if in response to a productivity shock future output is expected to be even larger than current output, then households will wish to borrow against future expected income thereby increasing today's consumption beyond the current increase in income. Panel (a) of figure 2 displays with a solid line the point estimate of the autocorrelation function of output growth up to fourth order and with broken lines a two-standard-error band. As mentioned earlier, the first-order autocorrelation is positive but small and not significantly different from zero. The second-order autocorrelation of output is negative, large, and significantly different from zero. Higher-order autocorrelations are negative but not highly significant. Panel (b) of figure 2 displays the autocorrelation function of the trade balance-to-output ratio along with its associated two-standard-error confidence band. This autocorrelation function is similar to that of an AR(1) process with a positive autoregressive coefficient of about 1/2. The first- and second-order autocorrelations are significantly above zero. Later in the paper we will ask whether a theoretical model driven by permanent shocks to productivity can account for the autocorrelation functions of output growth and the trade share observed in the data. This question has not been addressed in the literature on business cycles in

Figure 2: Autocorrelation Function of Output Growth

(a) Output



(b) Trade Balance-to-Output Ratio



emerging countries.[5]

3 The Model

The theoretical framework is the small open economy model presented in Schmitt-Grohé and Uribe (2003) augmented with permanent and transitory productivity shocks as in Aguiar and Gopinath (2004). The production technology takes the form

$$Y_t = a_t K_t^{\alpha} (X_t h_t)^{1-\alpha}, \tag{1}$$

where Y_t denotes output in period t, K_t denotes capital in period t, h_t denotes hours worked in period t, and a_t and X_t represent productivity shocks. We use upper case letters to denote variables that contain a trend in equilibrium, and lower case letters to denote variables that do not contain a trend in equilibrium.

The productivity shock a_t is assumed to follow a first-order autoregressive process in logs. That is,

$$\ln a_{t+1} = \rho_a \ln a_t + \epsilon^a_{t+1}.$$

The productivity shock X_t is nonstationary. Let

$$g_t \equiv \frac{X_t}{X_{t-1}}$$

denote the gross growth rate of X_t. We assume that the logarithm of g_t follows a first-order autoregressive process of the form

$$\ln(g_{t+1}/g) = \rho_g \ln(g_t/g) + \epsilon^g_{t+1}.$$

The innovations ϵ^a_t and ϵ^g_t are assumed to be i.i.d. processes with mean zero and variances σ_a^2 and σ_g^2, respectively. The parameter g measures the deterministic gross growth rate of the productivity factor X_t. The parameters $\rho_a, \rho_g \in [0, 1)$ govern the persistence of a_t and g_t, respectively.

Household face the following period-by-period budget constraint:

$$\frac{D_{t+1}}{1+r_t} = D_t - Y_t + C_t + I_t + \frac{\phi}{2}\left(\frac{K_{t+1}}{K_t} - g\right)^2 K_t, \tag{2}$$

[5]The autocorrelation function of output growth is at center stage in studies suggesting the failure of the RBC model to explain business cycles in industrialized countries. See, for example, Cogley and Nason (1995).

where D_{t+1} denotes the stock of debt acquired in period t, r_t denotes the domestic interest rate on bonds held between periods t and $t+1$, C_t denotes consumption, I_t denotes gross investment, and the parameter ϕ introduces quadratic capital adjustment costs. The capital stock evolves according to the following law of motion:

$$K_{t+1} = (1-\delta)K_t + I_t,$$

where $\delta \in [0,1)$ denotes the depreciation rate of capital.

In order to induce independence of the deterministic steady state from initial conditions, we assume that the country faces a debt-elastic interest-rate premium as in Schmitt-Grohé and Uribe (20003). Specifically, the domestic interest rate is assumed to be the sum of the world interest rate $r^* > 0$, assumed to be constant, and a country premium that is increasing in a detrended measure of aggregate debt as follows:

$$r_t = r^* + \psi\left(e^{\widetilde{D}_{t+1}/X_t - \bar{d}} - 1\right),$$

The variable $\tilde{D}_t$ denotes the aggregate level of external debt per capita, which the household takes as exogenous. In equilibrium, we have that $\tilde{D}_t = D_t$.

Consumers are subject to a no-Ponzi-scheme constraint of the form $\lim_{j\to\infty} E_t \frac{D_{t+j}}{\prod_{s=0}^{j}(1+r_s)} \leq 0$. The household seeks to maximize the utility function

$$E_0 \sum_{t=0}^{\infty} \beta^t \frac{\left[C_t - \theta\omega^{-1}X_{t-1}h_t^{\omega}\right]^{1-\gamma} - 1}{1-\gamma}, \tag{3}$$

subject to (1), (2), and the no-Ponzi-game constraint, taking as given the processes a_t, X_t, and r_t and the initial conditions K_0 and D_{-1}. Appendix B presents the first-order conditions associated with the household's optimization problem, and appendix C presents the equilibrium conditions of this economy expressed in terms of stationary variables.

4 Calibration and Estimation

The time unit in the model is meant to be a year. We assign values to the structural parameters using a combination of calibration and econometric estimation techniques.

We calibrate the parameters α, δ, ψ, $\bar{d}$, θ, ω, and γ using long-run data relations from Argentina as well as parameter values that are common in related business-cycle studies. Table 3 presents the calibrated parameter values. We set the parameter $\bar{d}$ to ensure an average trade balance-to-output ratio of about 1 percent, as observed in Argentina over the

Table 3: Calibration

Parameter	Value
γ	2
δ	0.1255
α	0.32
ψ	0.001
ω	1.6
θ	2.24
β	0.9224
$\bar{d}$	0.03

period 1913-2005. We follow Schmitt-Grohé and Uribe (2003) and assign a small value to the parameter ψ, measuring the sensitivity of the country interest-rate premium to deviations of external debt from trend, with the sole purpose of ensuring independence of the deterministic steady state from initial conditions, without affecting the short-run dynamics of the model. The value assigned to the depreciation rate δ implies an average investment ratio of about 19 percent, which is close to the average value observed in Argentina of about 17 percent. The value assumed for the discount factor β implies a relatively high average real interest rate of about 8.5 percent per annum, which is empirically plausible for an emerging market like Argentina. There is no reliable data on factor income shares for Argentina. We therefore set the parameter α, which determines the average capital income share, at 0.32, a value commonly used in the related literature. We follow the calibration of an emerging economy performed in Aguiar and Gopinath (2004) to set values to the remaining calibrated parameters.

We estimate econometrically the five parameters defining the stochastic process of the productivity shocks, g, ρ_g, σ_g, ρ_a, and σ_a, and the parameter governing the degree of capital adjustment costs, ϕ. To estimate these six parameters, we apply the generalized method of moments (GMM) using annual data from Argentina. The sample period is 1913 to 2005. Appendix A lists the data sources. We include 16 moment conditions: the variances and first- and second-order autocorrelations of output growth (g^Y), consumption growth (g^C), investment growth (g^I), and the trade balance-to-output ratio, (tby), the correlation of g^Y with g^C, g^I, and tby, and the unconditional mean of g^Y. Appendix D provides more details on the estimation procedure. The estimated parameters are presented in table 4. The permanent shock, g_t, is estimated to be more volatile and persistent than the transitory shock a_t. The standard deviation and autoregressive coefficient of g_t are estimated with precision. The same is true for the standard deviation of a_t. However, the autoregressive

Table 4: Estimated Structural Parameters

Parameter	Point Estimate	Standard Error
g	1.0013	0.0040
σ_g	0.0283	0.0046
ρ_g	0.4165	0.0537
σ_a	0.0119	0.0052
ρ_a	0.2122	0.3299
ϕ	0.9036	0.2995
Test of model fit	p value	0.13

Note: GMM estimates.

coefficient of a_t is not significantly different from zero. The p value of the test of model fit, shown at the bottom of table 4, indicates that the null hypothesis that the 16 theoretical moments included in the GMM estimation are equal to the corresponding sample moments is rejected with a confidence level of 13 percent.

5 Model Performance

Table 5 reports second moments implied by the theoretical model. To facilitate comparison, the table reproduces the empirical counterparts and their associated standard errors from table 2. In the RBC model, consumption growth is less volatile than output growth. This prediction represents a significant difficulty of the estimated RBC model, as excess consumption volatility is a key characteristic of emerging economies. The reason why consumption is less volatile than output in the model is the presence of trend stationary shocks. In effect, shutting off the stationary shock by setting $\sigma_a = 0$ results in a volatility of consumption growth 15 percent higher than the volatility of output growth. Intuitively, a positive but transitory shock produces an increase in the level of current output follow by a gradual decline toward its pre-shock level. Because output is expected to fall in the future, it is optimal for consumption-smoothing households to save, causing consumption to increase by less than income. In this way, there is a tendency for consumption to be less volatile than output. By contrast, in response to a positive and persistent shock to productivity growth, current output increases on impact and is expected to continue to grow in the future. This increasing profile for future expected income induces households to consume beyond the increase in current output. In this way, consumption tends to be more volatile than output in response to permanent productivity shocks. When both, trend stationary and permanent

Table 5: Comparing Model and Data: Summary Statistics

Statistic	g^Y	g^C	g^I	tby
Standard Deviation				
– Model	4.3	3.8	10.0	16.0
– Data	5.1	7.3	19.0	4.9
	(0.4)	(0.7)	(1.8)	(0.7)
Correlation with g^Y				
– Model	1.00	0.97	0.80	-0.05
– Data	1.00	0.77	0.66	-0.08
	(0.00)	(0.07)	(0.12)	(0.09)
Correlation with tby				
– Model	-0.05	-0.06	-0.06	1.00
– Data	-0.08	-0.30	-0.21	1.00
	(0.09)	(0.06)	(0.09)	(0.00)
Serial Correlation				
– Model	-0.13	-0.06	-0.13	0.99
– Data	0.05	0.04	0.21	0.58
	(0.12)	(0.09)	(0.10)	(0.07)

Note: g^Y, g^C, and g^I denote the growth rates of output per capita, consumption per capita, and investment per capita, respectively, and tby denotes the trade balance-to-output ratio. Standard deviations are reported in percentage points. Standard errors of sample-moment estimates are shown in parenthesis.

shocks to productivity are present, whether consumption growth turns out to be more or less volatile than output becomes a quantitative issue. In our estimated RBC model the tradeoff is resolved, counterfactually, in favor of consumption smoothing. As far as the volatility of consumption is concerned, therefore, the estimated model appears to assign too much importance to transitory shocks.

Yet, there is a dimension along which permanent productivity shocks appear to be over emphasized in the estimated model. Namely, the volatility of the trade balance. In the data, the trade-balance share in output is about as volatile as output growth, where as in the model the trade-balance share is around four times as volatile as output growth. Much of this discrepancy between data and model is due to the permanent component of productivity shocks. Shutting off the permanent productivity shock by setting σ_g equal to zero while holding constant all other parameter values results in a significant reduction in the volatility of the trade-balance share relative to the volatility of output growth.

Furthermore, there is a sense in which both the permanent and the transitory components of total factor productivity are estimated to be too insufficient volatile. For the model predicts too little volatility in investment growth, almost half the one observed in the data. Higher volatility in either source of uncertainty would contribute to ameliorating this significant mismatch between data and model.

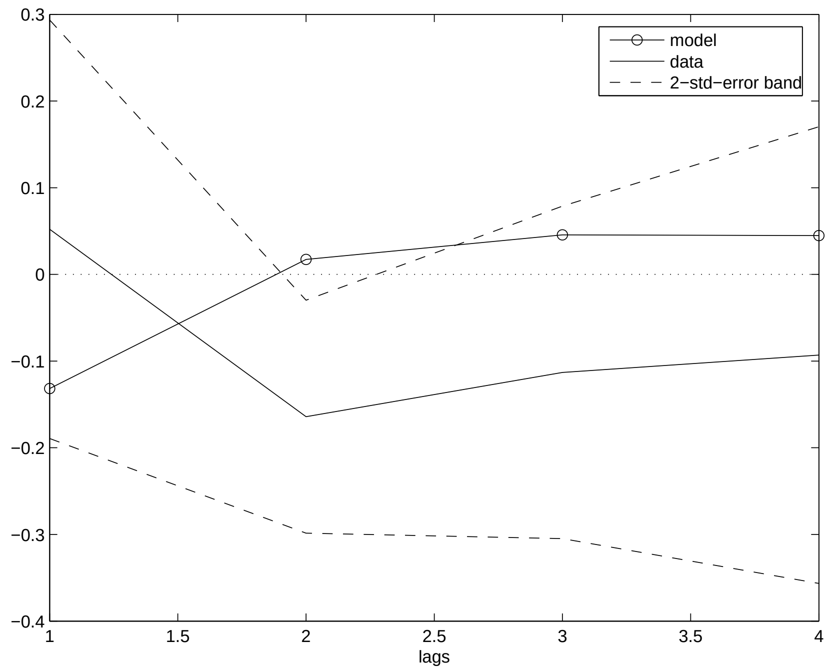
The RBC model correctly predicts a negative but mild correlation between output growth and the trade balance-to-output ratio. However, it significantly underpredicts the negative correlation between the trade-balance share and private consumption growth as well as the also negative correlation between the trade-balance share and investment growth. The fact that the trade-balance share is more correlated with the components of aggregate demand than with output growth may be an indication that shocks other than movements in total factor productivity could be playing a role in driving business cycles in Argentina.

Panel (a) of figure 3 depicts with a circled line the autocorrelation function of output growth implied by the model up to fourth order. For comparison, the figure also reproduces from figure 2 the estimated autocorrelation function with a solid line and a two-standard-error band with broken lines. The model fails to predict the sign of all four autocorrelations. The most significant departure of the model from the data occurs at order two, as the value implied by the model falls outside of the two-standard-error confidence band. This mismatch is notable because the second-order autocorrelation is the only element of the estimated autocorrelation function of output growth that is statistically significant.

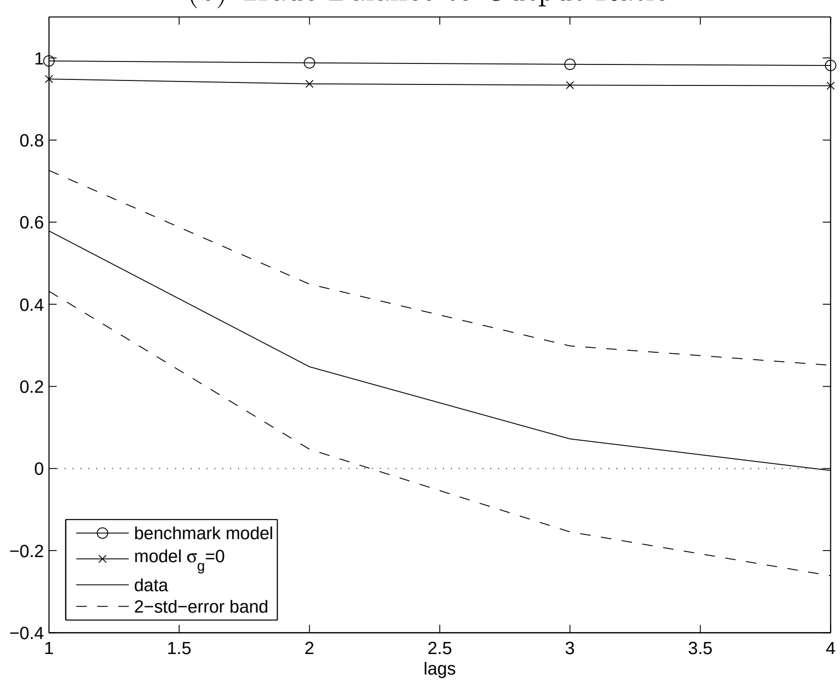
Panel (b) of figure 3 displays the theoretical and estimated autocorrelation functions of the trade balance-to-output ratio. In the RBC model, the trade-balance share follows a near random walk, with all four autocorrelations taking values close to unity. The theoretical au-

Figure 3: Comparing Model and Data: Autocorrelation Functions

(a) Output Growth



(b) Trade Balance-to-Output Ratio



tocorrelation function lies outside the confidence interval around the estimated counterpart. The point estimate of the autocorrelation function is downward slopping, takes a value below 0.6 at order one and then declines quickly toward zero. To understand why the trade balance displays quasi-random-walk dynamics in equilibrium it is useful to think of an endowment economy facing a constant world interest rate and in which the endowment follows a random walk process. In this environment, consider the response to an unanticipated innovation in output. In response to this shock, consumption experiences a once-and-for-all increase about equal in size to the increase in income, as households realizing that the increase in endowment is permanent, feel no need to increase savings. It follows that the trade balance is more or less unaffected by the shock and, as a result, the ratio of the trade balance to output inherits the random walk nature of the latter.

One might think that an implication of the analysis of the previous paragraph is that the autocorrelation function of the trade balance-to-output ratio should fall significantly in the absence of permanent shocks. This is not the case, however. Panel (b) of figure 3 displays with a crossed line the autocorrelation function of the trade balance-to-output ratio implied by the RBC model after shutting off the permanent shock by setting $\sigma_g = 0$, while keeping all other parameters at their baseline values. The autocorrelation function does not fall significantly; all of the autocorrelation coefficients remain above 90 percent and the autocorrelation function continues to lie much above its empirical counterpart. The reason for this result is that with stationary productivity shocks, although output and investment become stationary variables, consumption follows a near random walk typical of small open economies.[6]

Based on the analysis of this section, we conclude that the RBC model driven by permanent and transitory productivity shocks does a poor job at explaining business cycles in Argentina over the period 1913-2005.

6 Robustness

To establish the robustness of our results, we examine the sensitivity of the predictions of the model to a number of perturbations to the theoretical structure and estimation procedure. We begin by allowing for a more flexible stochastic process for the underlying productivity shocks. Specifically, we consider the case in which both the temporary and permanent technology shocks follow autoregressive processes of order 2. Formally, we assume that the

[6]The mismatch between the theoretical and empirical autocorrelation functions of the trade balance-to-output ratio remains even under purely transitory productivity shocks. Specifically, setting $\sigma_g = \rho_a = 0$ results in a flat autocorrelation function of the trade-balance share at above 0.87, a level still outside the confidence band around the empirical autocorrelation function.

laws of motion of a_t and g_t are now given by

$$\ln a_{t+1} = \rho_a \ln a_t + \rho_a' \ln a_{t-1} + \epsilon_{t+1}^a.$$

and

$$\ln(g_{t+1}/g) = \rho_g \ln(g_t/g) + \rho_g' \ln(g_{t-1}/g) + \epsilon_{t+1}^g.$$

Table 6 displays the predictions of the model driven by these processes (see the column labeled 'AR(2) Process'). For comparison, the table also shows sample moments and the model predictions under the benchmark (AR(1)) specification. Under the AR(2) specification, the model continues to fit the data poorly. The hypothesis that the theoretical and empirical moments employed in the estimation equal each other is rejected at a confidence level of 6 percent (see panel F). In addition, over all the model fails to match the data along the same lines as the version with AR(1) shocks. In particular, the model underestimates the volatility of consumption and investment growth and overestimates the volatility of the trade balance share. In addition the model displays great difficulty matching the autocorrelation functions of output growth and the trade balance share.

As a second robustness check, we eliminate all autocorrelations of order two from the moment conditions used in the GMM estimation. The results appear in the column labeled 'No AC Order 2' in table 6. This econometric specification yields a much poorer fit. The hypothesis that theoretical and sample moments are equal is now rejected at a confidence level of 2 percent.

Our third sensitivity experiment consists in replacing the growth rates of consumption and investment by the shares of these variables in GDP in the GMM estimation. This modification is motivated by the fact that in the model consumption and investment are cointegrated with output. The results are shown in the column labeled 'Shares' in table 6. Both the fit and business-cycle implications of the model are similar whether it is estimated in growth rates or shares.

The column labeled 'Cobb Douglas Preferences' in table 6 displays the equilibrium dynamics implied by a version of the model featuring a Cobb Douglas specification for the aggregator of consumption and leisure in preferences. Formally, we assume that the utility function is given by

$$E_0 \sum_{t=0}^{\infty} \beta^t \frac{[C_t(1-h_t)^{\omega}]^{1-\gamma} - 1}{1-\gamma}.$$

Unlike the preference specification assumed in the baseline model, the Cobb Douglas preference implies a nonzero wealth elasticity of labor supply. We set the parameter ω at 3.4 to ensure that in the deterministic steady state households supply 20 percent of their time to

Table 6: Robustness Analysis

	Data	Bench-mark	AR(2) Process	No AC Order 2	Shares for Estimation	Cobb Douglas	$\rho_a = 0$	$g = 1.0125$
				A. Standard Deviation				
g^Y	5.1	4.3	4.5	4.3	4.3	3.3	4.5	4.3
g^C	7.3	3.8	3.6	3.8	3.7	3.3	3.8	3.6
g^I	19	10	11	10	10	10	10	10
tby	4.9	16	14	16	16	19	16	21
				B. Correlation With g^Y				
g^Y	1	1	1	1	1	1	1	1
g^C	0.77	0.97	0.96	0.97	0.97	0.4	0.96	0.97
g^I	0.66	0.8	0.61	0.78	0.82	0.092	0.72	0.80
tby	-0.076	-0.045	-0.022	-0.043	-0.044	-0.11	-0.033	-0.026
				C. Correlation with *tby*				
g^Y	-0.076	-0.045	-0.022	-0.043	-0.044	-0.11	-0.033	-0.026
g^C	-0.3	-0.06	-0.051	-0.06	-0.058	-0.22	-0.054	-0.036
g^I	-0.21	-0.064	-0.083	-0.066	-0.061	-0.21	-0.068	-0.047
tby	1	1	1	1	1	1	1	1
				D. Autocorrelation Function of g^Y				
1st Order	0.052	-0.13	-0.22	-0.14	-0.12	0.25	-0.20	-0.13
2nd Order	-0.16	0.017	-0.043	0.024	0.0085	0.24	0.045	0.017
3rd Order	-0.11	0.046	0.076	0.048	0.04	0.17	0.048	0.046
4th Order	-0.093	0.045	0.039	0.045	0.042	0.12	0.043	0.046
				E. Autocorrelation Function of *tby*				
1st Order	0.58	0.99	0.98	0.99	0.99	0.97	0.99	0.99
2nd Order	0.25	0.99	0.98	0.99	0.99	0.95	0.99	0.99
3rd Order	0.072	0.98	0.98	0.98	0.98	0.93	0.98	0.99
4th Order	-0.0049	0.98	0.97	0.98	0.98	0.92	0.98	0.99
				F. Test of Fit				
p value	–	0.13	0.061	0.021	0.13	0.13	–	–

Note: g^Y, g^C, and g^I denote the growth rates of output per capita, consumption per capita, and investment per capita, respectively, and *tby* denotes the trade balance-to-output ratio. Standard deviations are reported in percentage points.

Table 7: Robustness Analysis Continued

	Data	Bench-mark	$\alpha = 0.4$	$\delta = 0.1$	$\gamma = 5$	$\beta = 0.89$	$\beta = 0.96$	1946-2005 Data	1946-2005 Model
					A. Standard Deviation				
g^Y	5.1	4.3	3.7	4.3	4.3	4.3	4.3	5.1	4.2
g^C	7.3	3.8	3.2	3.7	3.8	3.7	3.8	5.97	3.8
g^I	19	10	9.2	12	10	11	10	16	10.3
tby	4.9	16	18	17	25	24	7.4	2.1	16.1
					B. Correlation with g^Y				
g^Y	1	1	1	1	1	1	1	1	1
g^C	0.77	0.97	0.95	0.97	0.97	0.97	0.96	0.92	0.98
g^I	0.66	0.8	0.72	0.8	0.81	0.81	0.74	0.88	0.86
tby	-0.08	-0.05	-0.05	-0.05	-0.03	-0.02	-0.12	-0.30	-0.05
					C. Correlation with *tby*				
g^Y	-0.08	-0.05	-0.05	-0.05	-0.03	-0.02	-0.12	-0.30	-0.05
g^C	-0.30	-0.06	-0.07	-0.06	-0.04	-0.03	-0.17	-0.32	-0.06
g^I	-0.21	-0.06	-0.08	-0.06	-0.04	-0.04	-0.17	-0.31	-0.06
tby	1	1	1	1	1	1	1	1	1
					D. Autocorrelation Function of g^Y				
1st	0.05	-0.13	-0.11	-0.14	-0.14	-0.12	-0.17	0.08	-0.10
2nd	-0.16	0.02	0.06	0.01	0.02	0.02	0.009	-0.25	0.02
3rd	-0.11	0.05	0.08	0.04	0.05	0.05	0.04	0.03	0.05
4th	-0.09	0.05	0.07	0.05	0.05	0.05	0.04	0.05	0.05
					E. Autocorrelation Function of *tby*				
1st	0.58	0.99	0.99	0.99	1	1	0.95	0.68	0.99
2nd	0.25	0.99	0.99	0.99	1	1	0.92	0.28	0.99
3rd	0.07	0.98	0.98	0.98	0.99	0.99	0.90	0.06	0.99
4th	-0.00	0.98	0.98	0.98	0.99	0.99	0.88	-0.04	0.98
					R. Test of Fit				
p value	–	0.13	0.13	0.13	0.13	0.13	0.16	–	0.48

Note: g^Y, g^C, and g^I denote the growth rates of output per capita, consumption per capita, and investment per capita, respectively, and *tby* denotes the trade balance-to-output ratio. Standard deviations are reported in percentage points.

the labor market. We reestimate the structural parameters of the model following the same procedure as in the benchmark case.

Under the Cobb Douglas specification, the standard deviation of consumption growth is only slightly smaller than that of output growth. At first, this may appear as an improvement upon the baseline model. But the larger relative volatility of consumption growth occurs not because the volatility of consumption growth itself is higher under Cobb Douglas preferences than under the baseline preferences (indeed it is smaller), but rather because the volatility of output falls significantly. As a result, the model with Cobb Douglas preferences does a poorer job than the baseline model at matching both the volatility of consumption growth and the volatility of output growth observed in the data. The reason why the volatility of output growth falls under Cobb Douglas preferences is quite intuitive: A positive permanent productivity shock produces a negative wealth effect on labor supply. As a result, such a shock induces a smaller increase in hours and output under the Cobb Douglas specification than under the baseline specification.[7]

Under Cobb Douglas preferences, the correlation of investment growth with output growth drops significantly to a value close to zero. This decline is also related to wealth effects. In effect, a positive permanent productivity shock increases the marginal productivity of capital inducing firms to invest more. But at the same time, as explained before with Cobb Douglas preferences labor increases by less than under the benchmark preference formulation. Because the marginal product of capital is increasing in labor, we have that the incentives to invest in response to a positive permanent productivity shock are weaker under Cobb Douglas preferences than under the baseline preferences.

Of the six structural parameters of the model that are estimated econometrically, the autoregressive coefficient of the stationary productivity shock, ρ_a, is the one estimated with the greatest degree of uncertainty. Although the point estimate of this parameter is relatively high, 0.21, its large standard deviation of 0.33 makes it statistically not different from zero. This result motivated us to simulate the model setting ρ_a equal to zero while keeping all other parameters at their baseline values. The result is shown in the penultimate column of table 6. The predictions of the model are not much affected by restricting ρ_a to be zero. Similar results emerge if instead of fixing all parameters other than ρ_a at their baseline values one reestimates the model holding ρ_a fixed at zero. The robustness of the results along this dimension is to some extent not surprising. For the mere fact that the GMM method assigns a high standard deviation to a parameter estimate means precisely that the fit of the model

[7]Indeed in response to a positive innovation in g_t hours fall under the Cobb Douglas specification, but increases under the baseline specification. Note also that in our intuition we ignore the role of the stationary shock a_t. The reason is that this shock is estimated to have a small autoregressive coefficient and therefore produces negligible wealth effects.

is insensitive to variations in the parameter in question.

We also performed robustness checks involving the parameter g, measuring the deterministic gross growth rate of the economy. As noted earlier in the paper, the baseline estimation of the model delivers a point estimate of g close to unity, implying virtually zero growth in the nonstochastic steady state. By contrast, the average growth rate of the Argentine economy over the sample period is about 1.25 percent. To gauge the sensitivity of the results to perturbing g, we fixed this structural parameter at the elevated value 1.0125 (implying a nonstochastic growth rate of 1.25 percent) and simulated the model holding all other parameters at their baseline values. The last column of table 6 shows that the implications of the model under this parameterization are not altered in any significant way. (We arrived at a similar conclusion after reestimating the model under the restriction $g = 1.01$.) We also explored the possibility that the low point estimate of g may be due to an asymmetry in the approximation accuracy across moment conditions. This asymmetry arises because in generating theoretical moments, we resort to a linear approximation of the equilibrium conditions of the model. This approximation delivers a second-order accurate approximation to all second moments but only a first-order accurate approximation of all first moments, including the average growth rate of output. Accordingly, we proceeded to approximate the average growth rate of output implied by the model using a second-order approximation to the equilibrium conditions. We applied the procedure and computer code developed by Schmitt-Grohé and Uribe (2004). The resulting point estimate of g continues to be close to unity, as under the baseline estimation.

All of the sensitivity experiments discussed up to this point involve the set of six parameters that we estimate econometrically. We also explored the robustness of our results to perturbing the values assigned to the nonestimated structural parameters. Table 7 displays the model predictions under five alternative parameterizations. (a) Low labor share: The baseline calibration assumes a share of labor income in GDP of 68 percent. Although there is no data on national income accounts for Argentina, some have argued that the labor share may be lower than the value typically used in business-cycle studies. See the discussion of this issue in Kydland and Zarazaga, (2002). Following this authors, we set $\alpha = 0.4$. (b) Low depreciation rate: The baseline calibration assumes a depreciation rate of 12.5 percent per year. We lower this value to 10 percent, a number more widely used in the literature. (c) Low elasticity of intertemporal substitution: Some authors argue that emerging-country aggregate data are consistent with relatively interest-inelastic consumption growth rates. Reinhart and Végh (1995), for example, using data from Argentina estimate γ to be 5. We compute the dynamics of the model using this parameter value. (d) High country premium: The baseline model assumes a value of β of 0.92, which is consistent with a real interest rate of 8.5 percent per annum. Given a value of about 4 percent typically assumed for the average real interest rate in developed countries, our baseline value of β implies a long-run country premium of about 4.5 percent. We explore the properties of the model under a value of β of 0.89, which implies a high country premium of about 8.5 percent. (e) We also consider the case of a low country premium by setting $\beta = 0.96$. This discount factor is consistent with no country premium in the long run. Our reading of the simulations reported in table 7 is that the results obtained under the baseline parameterization are generally robust to the parameter perturbations considered. Paradoxically, the best fit of the model obtains when we counterfactually set the long-run country premium at zero by making households relatively patient ($\beta = 0.96$).

Finally, the last two columns of table 7 display empirical and theoretical moments obtained using postwar data (1946-2005). The empirical regularities identified using the sample 1913-2005 continue to characterize the data when only the postbellum period is taken into account. In addition, the model estimated using the shorter sample fails to mimic the data along the same dimensions as the model estimated using data from the early twentieth century. In spite of these similarities, the p-value of the J test of model fit increases notably after shortening the sample, which we interpret as a reflection of the low power of the test.

6.1 Argentina 1865-2004

Thus far, we have focused on Argentine data ranging from the early twentieth century to the present. We now expand the data set to the period 1865 to 2004. The source of this longer data set is a recent compilation published by Ferreres (2005).

As a preliminary check, we begin by exploring the subsample 1913-2004 as a point of comparison with the data set we have been using up to this point, to which we will refer as the baseline data set. The time series for output in the Ferreres and baseline data sets are virtually identical. The correlation between these two time series is 0.999. This is not quite the case for the other variables we use. In effect, the correlations of private consumption growth, investment growth, and the trade-balance-to-output ratio in the Ferreres data set with their respective counterparts in the baseline data set are 0.81, 0.88, and 0.89.

Because of the discrepancies in the Ferreres and the baseline data sets, we first estimate the model using the subsample 1913-2004 from the Ferreres data set. The results are shown in table 8. Comparing the empirical and theoretical second moments displayed in the table leads to identical conclusions to those drawn using the baseline data set. In particular, the model counterfactually predicts consumption growth to be less volatile than output growth, investment growth to be half as volatile as its empirical counterpart, and the trade-balance

Table 8: Argentina 1865-2004

	1913-2004		1865-2004	
	Data	Model	Data	Model
		A. Standard Deviation		
g^Y	5.0	4.2	6.4	6.2
g^C	5.5	3.8	7.2	4.3
g^I	18.0	10.0	26.0	14.0
tby	2.5	16.0	5.4	21.0
		B. Correlation with g^Y		
g^Y	1	1	1	1
g^C	0.90	0.97	0.63	0.96
g^I	0.78	0.85	0.68	0.52
tby	-0.19	-0.051	-0.11	0.01
		C. Correlation with tby		
g^Y	-0.19	-0.05	-0.11	0.01
g^C	-0.23	-0.06	-0.06	-0.02
g^I	-0.26	-0.06	-0.16	-0.07
tby	1	1	1	1
		D. Autocorrelation Function of g^Y		
1st Order	0.03	-0.11	-0.01	-0.30
2nd Order	-0.19	0.02	-0.12	0.02
3rd Order	-0.08	0.05	0.04	0.03
4th Order	-0.07	0.05	-0.06	0.02
		E. Autocorrelation Function of tby		
1st Order	0.70	0.99	0.87	0.99
2nd Order	0.35	0.99	0.76	0.98
3rd Order	0.08	0.99	0.68	0.98
4th Order	-0.08	0.98	0.61	0.98
		F. Test of Fit		
p-value	–	0.13	–	0.01

Note: g^Y, g^C, and g^I denote the growth rates of output per capita, consumption per capita, and investment per capita, respectively, and tby denotes the trade balance-to-output ratio. Standard deviations are reported in percentage points. Data Source: Own calculations based on data from Ferreres (2005).

ratio to be much more volatile than it is in the data. Also the model does a poor job replicating the autocorrelation functions of output and the trade-balance-to-output ratio. The fit of the model using the Ferreres data set is the same as the one obtained using the baseline data, with a p value of 0.13 in both cases.

Having established the robustness of our main results to using the Ferreres data set, we proceed to expand the sample to include the period 1865-2004. The longer sample displays higher volatility in all time series, especially in aggregate investment. The RBC model has a hard time fitting the longer time series. The estimated p-value is only 0.01, implying that the hypothesis that the theoretical and empirical moments used for estimation are identical is rejected with a confidence level of about 1 percent. The model fails to match the data along the same dimensions as is the case under the shorter sample. Namely, the model produces counterfactual predictions for the volatilities of output growth, consumption growth, and the trade-balance ratio, and the autocorrelation functions of output and the trade-balance ratio.

7 Conclusion

The present study scrutinizes the hypothesis that business cycles in developing economies are driven by permanent and/or transitory exogenous shifts in total factor productivity and transmitted through the familiar mechanism of the frictionless RBC model.

The starting point of our investigation is the notion that if permanent shocks are to play an important role in the macroeconomy, then long time series are called for both for characterizing business cycles, as well as for identifying the parameters defining the stochastic process of the underlying shocks. Accordingly, we build a data set covering almost a century of aggregate data from Argentina. We use these data to estimate a battery of statistics that provide a fairly complete picture of the Argentine business cycle. We then formulate a standard RBC model of the small open economy driven by permanent and transitory productivity shocks. We estimate the parameters of these productivity shock processes and other structural parameters of the model using our data from Argentina.

Comparing the predictions of the model with the data, we arrive at the conclusion that the RBC model does a poor job at explaining business cycles in Argentina. One challenge for the model is the empirical fact that in Argentina, as in many other emerging countries, private consumption growth is more volatile than output growth. In order for the RBC model to explain this fact, permanent shocks to productivity must be the predominant driving force. The econometric estimation does not assign permanent shocks this predominance. On the other hand, there is a sense in which permanent shocks are too important in the model. In effect, in the model the trade balance-to-output ratio is about four times as volatile as output growth, whereas in the data these two variables are equally volatile. This mismatch between model and data is due to the permanent component of productivity shocks. For shutting off this shock in the theoretical model causes the volatility of the trade balance-share to fall by about 75 percent.

Moreover, there are aspects of the data that the RBC model cannot match quite independently of the relative importance of permanent and transitory productivity shocks. This is the case with the autocorrelation function of the trade balance-to-output ratio. In the model, the trade-balance share follows a near random walk, with an autocorrelation function close to one. In the data, the autocorrelation of the trade-balance share is far below unity. The presence of permanent shocks certainly contributes to this discrepancy. But not much. Shutting off the permanent shock results in only a small downward shift in the autocorrelation function of the trade-balance share. Furthermore, the problem cannot be attributed to the fact that the transitory productivity shock is too persistent, because shutting off the permanent shock and at the same time making the stationary shock purely transitory still results in an autocorrelation function of the trade-balance share much above its empirical counterpart.

Our findings suggest that the RBC model driven by productivity shocks does not provide an adequate explanation of business cycles in emerging countries like Argentina. We note, however, that all of the results reported in the present study are based on a joint hypothesis that the productivity shocks and the transmission mechanism built in the RBC model fit the data. Consequently, discerning whether the failure of the RBC model is to be blamed on productivity shocks or on the model's propagation mechanism remains a subject for future research.

Appendix

A. Data Sources

Argentina

GDP, Investments, Exports, and Imports

1913 - 1980: Domenech, Roberto L. (1986), table 2.

1981 - 1992: Dirección Nacional de Cuentas Nacionales (1996). Available at `http://www.mecon.gov.ar/secpro/dir_cn/ant/contenido.htm`.

1993 - 2005: Secretearia de Politica Economica (2006). Available at `http://www.mecon.gov.ar/peconomica/informe/indice.htm`.

Private Consumption:

1913 - 1980: Domenech, Roberto L. (1986), table 2.

1980 - 1992: República Argentina, Ministerio de Economía y Obras y Servicios Públicos (1994).

1993 - 2005: Secretearia de Politica Economica (2006). Available at `http://www.mecon.gov.ar/peconomica/informe/indice.htm`.

Population:

1913 - 1949: Domenech, Roberto L. (1986), table 4.

1950 - 2005: CEPAL and INDEC (2004). Available at `http://www.indec.gov.ar/principal.asp?id_tema=165`.

United States:

GDP

1913 - 1928: Balke, N. S. and Robert J. G. (1989).

1929 - 2005: Bureau of Economic Analysis. Available at `www.bea.gov`.

Population

1900 - 2005: U.S. Census, Statistical Abstract of the United States. Available at `http://www.census.gov/compendia/statab/population/`.

B. Optimality Conditions of the Household's Problem

Letting $\lambda_t X_{t-1}^{-\gamma}$ denote the Lagrange multiplier associated with the sequential budget constraint, the optimality conditions associated with this problem are (1), (2), the no-Ponzi-game constraint holding with equality, and

$$\left[C_t/X_{t-1} - \theta\omega^{-1}h_t^{\omega}\right]^{-\gamma} = \lambda_t$$

$$\left[C_t/X_{t-1} - \theta\omega^{-1}h_t^{\omega}\right]^{-\gamma} \theta h_t^{\omega-1} = (1-\alpha)a_t \left(\frac{K_t}{X_{t-1}h_t}\right)^{\alpha} \left(\frac{X_t}{X_{t-1}}\right)^{1-\alpha} \lambda_t$$

$$\lambda_t = \beta\frac{1+r_t}{g_t^{\gamma}}E_t\lambda_{t+1}$$

$$\begin{aligned}\left[1+\phi\left(\frac{K_{t+1}}{K_t}-g\right)\right]\lambda_t &= \frac{\beta}{g_t^{\gamma}}E_t\lambda_{t+1}\left[1-\delta+\alpha a_{t+1}\left(\frac{X_{t+1}h_{t+1}}{K_{t+1}}\right)^{1-\alpha}\right.\\ &\left.+\phi\left(\frac{K_{t+2}}{K_{t+1}}\right)\left(\frac{K_{t+2}}{K_{t+1}}-g\right)-\frac{\phi}{2}\left(\frac{K_{t+2}}{K_{t+1}}-g\right)^2\right]\end{aligned}$$

C. Equilibrium Conditions

Define $y_t = Y_t/X_{t-1}$, $c_t = C_t/X_{t-1}$, $d_t = D_t/X_{t-1}$, and $k_t = K_t/X_{t-1}$. Then, a stationary competitive equilibrium is give by a set of processes stationary solution to the following equations:

$$[c_t - \theta\omega^{-1}h_t^{\omega}]^{-\gamma} = \lambda_t$$

$$\theta h_t^{\omega-1} = (1-\alpha)a_t g_t^{1-\alpha}\left(\frac{k_t}{h_t}\right)^{\alpha}$$

$$\lambda_t = \frac{\beta}{g_t^{\gamma}}\left[1+r^* + \psi\left(e^{d_t-\bar{d}}-1\right)\right]E_t\lambda_{t+1}$$

$$\begin{aligned}\left[1+\phi\left(\frac{k_{t+1}}{k_t}g_t-g\right)\right]\lambda_t &= \frac{\beta}{g_t^{\gamma}}E_t\lambda_{t+1}\left[1-\delta+\alpha a_{t+1}\left(\frac{g_{t+1}h_{t+1}}{k_{t+1}}\right)^{1-\alpha}\right.\\ &\left.+\phi\frac{k_{t+2}}{k_{t+1}}g_{t+1}\left(\frac{k_{t+2}}{k_{t+1}}g_{t+1}-g\right)-\frac{\phi}{2}\left(\frac{k_{t+2}}{k_{t+1}}g_{t+1}-g\right)^2\right]\end{aligned}$$

$$\frac{d_{t+1}}{1+r_t}g_t = d_t - y_t + c_t + i_t + \frac{\phi}{2}\left(\frac{k_{t+1}}{k_t}g_t - g\right)^2 k_t,$$

$$r_t = r^* + \psi\left(e^{d_t-\bar{d}}-1\right),$$

$$k_{t+1}g_t = (1-\delta)k_t + i_t$$

$$y_t = a_t k_t^{\alpha}(g_t h_t)^{1-\alpha}$$

D. GMM Estimation Procedure

Let $b \equiv [g\, \sigma_g\, \rho_g\, \sigma_a\, \rho_a\, \phi]'$ be the 6×1 vector of structural parameters to be estimated. We write the moment conditions as:

$$u_t(b) = \begin{bmatrix} E_{gy}(b) - g_t^Y \\ \sigma^2_{gy}(b) - (g_t^Y - \bar{g}^Y)^2 \\ \sigma^2_{gc}(b) - (g_t^C - \bar{g}^C)^2 \\ \sigma^2_{gi}(b) - (g_t^I - \bar{g}^I)^2 \\ \sigma^2_{tby}(b) - (tby_t - \overline{tby})^2 \\ \rho_{gy,gc} - \frac{(g_t^Y - \bar{g}^Y)(g_t^C - \bar{g}^C)}{\sigma_{gy}(b)\sigma_{gc}(b)} \\ \rho_{gy,gi} - \frac{(g_t^Y - \bar{g}^Y)(g_t^I - \bar{g}^I)}{\sigma_{gy}(b)\sigma_{gi}(b)} \\ \rho_{gy,tby} - \frac{(g_t^Y - \bar{g}^Y)(tby_t - \overline{tby})}{\sigma_{gy}(b)\sigma_{tby}(b)} \\ \rho_{gy1}(b) - \frac{(g_t^Y - \bar{g}^Y)(g_{t-1}^Y - \bar{g}^Y)}{\sigma^2_{gy}(b)} \\ \rho_{gy2}(b) - \frac{(g_t^Y - \bar{g}^Y)(g_{t-2}^Y - \bar{g}^Y)}{\sigma^2_{gy}(b)} \\ \rho_{gc1}(b) - \frac{(g_t^C - \bar{g}^C)(g_{t-1}^C - \bar{g}^C)}{\sigma^2_{gc}(b)} \\ \rho_{gc2}(b) - \frac{(g_t^C - \bar{g}^C)(g_{t-2}^C - \bar{g}^C)}{\sigma^2_{gc}(b)} \\ \rho_{gi1}(b) - \frac{(g_t^I - \bar{g}^I)(g_{t-1}^I - \bar{g}^I)}{\sigma^2_{gi}(b)} \\ \rho_{gi2}(b) - \frac{(g_t^I - \bar{g}^I)(g_{t-2}^I - \bar{g}^I)}{\sigma^2_{gi}(b)} \\ \rho_{tby1}(b) - \frac{(tby_t - \overline{tby})(tby_{t-1} - \overline{tby})}{\sigma^2_{tby}(b)} \\ \rho_{tby2}(b) - \frac{(tby_t - \overline{tby})(tby_{t-2} - \overline{tby})}{\sigma^2_{tby}(b)} \end{bmatrix},$$

where $Ex(b)$ denotes the expected value of the variable x_t implied by the theoretical model, $\sigma_x(b)$ denotes the standard deviation of x_t implied by the theoretical model, $\rho_{xy}(b)$ denotes the correlation between x_t and y_t implied by the theoretical model, and ρ_{xj} denotes the autocorrelation of order j of x_t implied by the theoretical model. All of these statistics are functions of the vector b of structural parameters. We denote by $\bar{x} \equiv \sum_{t=1}^T x_t$ the sample mean of x_t, where T is the sample size. We compute moments implied by the theoretical model by solving a linearized version of the system of equilibrium conditions with respect to the logarithm of all variables except the trade-balance share in GDP, which we keep in levels.

Define $J(b, W) = \bar{u}'W\bar{u}$, where $\bar{u}(b)$ denotes the sample mean of $u_t(b)$ and W is a symmetric positive definite matrix compatible with $\bar{u}(b)$. The GMM estimate of b, denoted $\hat{b}$, is given by

$$\hat{b} = \underset{b}{\operatorname{argmin}}\, J(b, W).$$

The matrix W is estimated in two steps. For more details see Burnside (1999).

References

Aguiar, Mark, and Gita Gopinath, "Emerging Market Business Cycles: The Cycle is the Trend," NBER working paper No. 10734, August 2004.

Balke, N. S. and Robert J. G., "The Estimation of Prewar Gross National Product: Methodology and New Evidence," *Journal of Political Economy 97,* February 1989, 38-92.

Basu, S. and A. Taylor, "Business Cycles in International Historical Perspective," NBER working paper 7090, April 1999.

Burnside, Craig, "Real Business Cycle Models: Linear Approximation and GMM Estimation," manuscript, The World Bank, May 1, 1999.

CEPAL and INDEC, "Estimaciones y proyecciones de población. 1950-2015," Buenos Aires, 2004.

Cogley, Timothy and James M. Nason, "Output Dynamics in Real Business Cycle Model," *American Economic Review 85,* June 1995, 492-511.

Dirección Nacional de Cuentas Nacionales, "Oferta y Demanda Globales 1980-1995," Secretaría de Programación Económica, Ministerio de Economía, Buenos Aires, September 1996.

Domenech, Roberto L., "Estadísticas de la Evolución Económica de Argentina 1913-1984," *Estudios (de IEERAL) 39,* July-September 1986, 103-185.

Ferreres, Orlando J.,, *Dos Siglos de Economía Argentina (1810-2004). Historia Argentina en Cifras,* Fundación Norte y Sur: Buenos Aires, 2005.

King, R.G., C.L. Plosser and S.T. Rebelo, "Production, Growth and Business Cycles I: The Basic Neoclassical Model," *Journal of Monetary Economics 21,* 1988, 191-232.

Kydland, Finn E. and Edward C. Prescott, "Time to Build and Aggregate Fluctuations," *Econometrica 50,* November 1982, 1345-70.

Kydland, Finn E. and Carlos E. J. M. Zarazaga, "Argentina's Lost Decade," *Review of Economic Dynamics 5,* January 2002, 152-165.

Neumeyer, Pablo A. and Fabrizio Perri, "Business Cycles in Emerging Economies: The Role of Interest Rates," *Journal of Monetary Economics 52,* March 2005, 345-380.

Reinhart, Carmen M., Végh, Carlos A. , "Nominal Interest Rates, Consumption Booms, and Lack of Credibility: A Quantitative Examination ," *Journal of Development Economics,* April 1995, 46(2), 357-78.

República Argentina, Ministerio de Economía y Obras y Servicios Públicos, "Argentina en crecimiento, 1994- 1996," Buenos Aires, 1994.

Schmitt-Grohé, Stephanie and Martín Uribe, "Closing Small Open Economy Models," *Journal of International Economics,* 61, October 2003/ pp. 163-185 .

Schmitt-Grohé, Stephanie and Martín Uribe, "Solving Dynamic General Equilibrium Models Using a Second-Order Approximation to the Policy Function," *Journal of Economic Dynamics and Control 28,* January 2004, 755-775.

Secretaría de Política Económica, "Informe Económico Trimestral No. 54," Buenos Aires, 2006.

Sturzenegger A. and R. Moya, "Economic Cycles," in G. della Paolera and A. Taylor editors, *A New Economic History of Argentina,* Cambridge University Press, November 2003.

Uribe, Martín, "Lectures in Open Economy Macroeconomics," manuscript, Duke University, Spring 2006.