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THE IMPACT OF FINANCIAL DEVELOPMENT AND TRADE ON THE ECONOMIC GROWTH OF BOLIVIA

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The relationship of economic growth with financial development and trade openness is analyzed with annual time-series data for Bolivia during the 1940-2010 period. The analysis is an advance over previous work in several ways. First, the hypothesis of a long-run relationship between these variables is tested using bivariate cointegrated systems and employing the methodology of cointegration analysis. Second, causality tests utilizing standard Granger regressions and ECM models are carried out to determine the direction of causality between indicators of economic growth and financial development, and economic growth and trade openness. Lastly, the study comprises a period of seventy years, a first for a study of this kind on Bolivia. The empirical results demonstrate that there is indeed a long-run equilibrium relationship, and that unidirectional Granger causality runs from the indicators of financial development and trade openness to economic growth.

JEL classification codes: C10, E01, F43, G00, O54
Key words: financial development, trade openness, economic growth, cointegration test, Granger causality test

I. Introduction

The existence of correlation between financial development and economic growth has been documented in a number of empirical studies. Some have found a positive association between these variables while others have discovered that financial development may in some cases hamper economic growth. Likewise, the relationship between trade openness and economic growth has been thoroughly analyzed, and the findings in most papers support the notion that greater openness to trade generates positive growth effects.

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The literature on the relationship between financial development and economic growth is vast. The empirical work on the issue of causality between these variables is not as abundant. Goldsmith (1969) was the first to document a positive correlation between growth and indicators of financial development. In line with Goldsmith's findings, King and Levine (1993) demonstrate that better financial systems improve the probability of successful innovation and thereby accelerate economic growth, while financial sector distortions reduce the rate of economic growth by reducing the rate of innovation; Rajan and Zingales (1998) find that industrial sectors in countries with relatively better developed financial markets grow faster than equivalent sectors in less developed financial settings. Beck and Levine (2004), using a panel data set for the 1976-1998 period, find that stock markets and banks positively influence economic growth; and Levine (2005) concludes that the preponderance of evidence suggests that both financial intermediaries and markets matter for growth.

Representative papers that found little or no evidence of a positive correlation between financial development and growth are: Shan and Morris (2002), who for a set of nineteen OECD countries and China find meager evidence that financial development 'leads' economic growth, either directly or indirectly, which, in their view, casts doubt on claims that financial development is a necessary and perhaps sufficient precursor to economic growth; and, Boulila and Trabelsi (2004) who explore the issue of causality in the Middle East and North Africa and find little support to the view that finance is a leading sector in the determination of long-run growth in the countries of the region.

In studies that have focused on Latin America, De Gregorio and Guidotti (1995) examined the empirical relationship between long-run growth and financial development proxied by the ratio between bank credit to the private sector and GDP. They find that this proxy is positively correlated with growth in a large cross-country sample, but its impact changes across countries and is negative in a panel data for Latin America, due –they argue– to financial liberalization taking place in a poor regulatory environment. Moreover, Bittencourt (2010) analyzes the period 1980-2007 for a set of four Latin American countries, including Bolivia, and based on panel time-series data confirms the Schumpeterian prediction which suggests that finance authorizes the entrepreneur to invest in productive activities, and therefore to promote economic growth. He highlights the importance of macroeconomic stability as a necessary pre-condition for financial development.

In country-specific studies, Murinde and Eng (1994), analyzing Singapore for the period 1979-1990, find unidirectional causality from financial development to economic growth, which they thought justified the deliberate financial restructuring policy implemented by the government of that country during the 1980s. Yucel (2009), analyzing Turkey for the 1989-2007 period, finds bidirectional causality between financial development, trade openness and growth, indicating that economic policies aimed at financial development and trade openness have a statistically significant impact on economic growth. Chang and Caudill (2005), analyzing Taiwan for the period 1962-1998, find unidirectional causality running from financial development to growth, highlighting the importance of financial development in Taiwan's growth, particularly during the 1990s.

The literature on the relationship between trade openness and economic growth is also plentiful. The evidence obtained for a cross-section of countries and for individual ones is mixed, though the preponderance of findings suggests a positive association between these variables. Issues of disagreement and debate concern the construction of an appropriate trade openness index, the use of cross-section analysis, and the direction of causality. Rodrik (1997) argues that in most studies of openness and growth, the indicators used do not appropriately reflect the trade regimes. Furthermore, and in a closely related study, Rodriguez and Rodrik (2000) criticize the selection of several trade openness indicators and find little evidence that open trade policies –in the sense of lower tariff and nontariff barriers to trade– are significantly associated with economic growth. In the same line, Yanikkaya (2003) finds that trade liberalization does not have a simple and straightforward relationship with growth, and Dowrick and Golley (2004), analyzing the winners and losers of trade, find that since 1980 the benefits of trade accrued mostly to richer economies. Skeptic positions are not new. Jung and Marshall (1985) were among the first to question conventional wisdom assumptions about the purported positive association between trade and growth. Analyzing the period 1950-1981, they find the direction of causality between exports and growth to be inconclusive for a set of Asian nations.

At the other end of the spectrum, Frankel and Romer (1999) find that trade has a quantitatively large and robust, though only moderately statistically significant, positive effect on income for a set of countries. With this paper they have become leading advocates of the view that trade does indeed play a positive role on growth. Levine and Renelt (1992), in line with the advocates for free trade, identify a positive, robust correlation between growth and the share of investment in GDP, and between the investment share and the ratio of international trade to GDP. Edwards (1992) finds that countries that liberalize their international trade and become more open will tend to grow faster. Sachs and Warner (1995), in a detailed historical description of the evolution of societies, provide strong evidence of convergence among open economies during the period 1970-89, as well as evidence of accelerated growth in countries that have undertaken market-oriented reforms. Van Den Berg (1996) addresses the causality issue in six Latin American countries by comparing results from single equation and simultaneous equation models. He finds that both imports and exports have positive and distinct effects on economic growth and that the impact of openness on growth is higher and more significant in simultaneous equation models. Harrison (1996) studies the effects of trade openness on growth using panel data and compares predictions of several measures on trade openness. Granger causality tests show that causality between openness and growth runs in both directions. Billmeier and Nannicini (2008), after some econometric adjustments to cross-section and individual data sets, confirm a positive correlation between trade openness and growth in selected regions. Lee et al. (2004) also find a positive, though small, effect of trade openness on growth, while Winters (2004) determines that the most plausible conclusion is that liberalization generally induces a temporary, but possibly long-lived, increase in growth. More recently, Lucas (2007) proposes a model to describe the evolution of real GDPs in open economies, and using the Sachs-Warner definition of openness predicts convergence of income levels and growth rates; and Wacziarg and Welch (2008) find that liberalization has, on average, robust positive effects on growth, openness and investment rates within countries.

In country-specific studies, Edwards and Lederman (1998) illustrate the Chilean experience during the period 1974 through the 1990s. Their findings highlight the positive effects of unilateral trade liberalization on the country's economic growth. Jenkins (1996), in a slightly different paper that analyzes the impact of trade liberalization on Bolivian manufactured goods during the 1980s and part of the 1990s, concludes that the improved export performance in manufactured goods was largely the result of a more realistic and stable real exchange rate after 1985, and not of trade policy reforms, which, according to him, had little impact. Finally –though by no means exhaustively– Utkulu and Özdemir (2004) examine the long-run relationship between trade and economic growth in Turkey and find that a positive association between these variables is indeed plausible. Their most important finding, however, is that a reduction in trade distortions is linked to growth, highlighting the importance of trade policy on the economic performance of that country.

The motivation and main innovation of this work is threefold: first, this study analyzes whether there is a long-run economic association of economic growth with financial development and trade openness. Second, causality tests are performed to determine the direction of causality between these variables. Lastly, where other authors analyzing Bolivia tend to concentrate only on the recent past (see, for instance, Sachs and Warner 1995, Jenkins 1996, and Bittencourt 2010, who study the years after 1980 and do not extend the analysis beyond 2007), in this paper I analyze a period of seventy years, from 1940 to 2010. In all three counts, the findings presented here represent a first for this country.

Testing the validity of the assumptions concerning economic growth, financial development and trade openness in Bolivia should be particularly interesting, as this country has experienced complete sets of political and economic cycles, particularly during the period analyzed here. During the 1940s the country dabbled in populist military dictatorships; the 1950s and 1960s brought nationalizations in the most important sectors of the economy and the first episode of severe inflation (in the mid 1950s); then followed the 1970s, a decade of debt-induced growth that preceded the first recorded episode of hyperinflation (in the 1980s) which was not the result of civil war, foreign war or an internal political revolution. The latter part of the 1980s and the 1990s –a period of structural adjustments– was the first time the country modernized its institutions.[1] Trade was liberalized, price controls and a wide array of subsidies largely disappeared, and new, forward-looking legislations for banks and other financial intermediaries (in 1993), the Central Bank (in 1996), and the Securities Exchange (in 1998) came into law. A backlash to these reforms began at the turn of the new century and has lasted until the present. Many of the reforms undertaken during the 1990s have been reversed, particularly those concerning trade policy and regulation of the finance industry, and the results of these policies have yet to be fully understood.

This work is organized as follows: Section II presents the empirical strategy; Section III introduces the data; the results are presented in Section IV; Section V concludes.

II. Empirical strategy

Many macroeconomic time series contain unit roots, which tend to be dominated by stochastic trends.[2] The presence of a stochastic trend is determined by testing

[1] This assertion refers to events comprised between 1940 and 2010. The latter part of the 19th century and first two decades of the 20th century were characterized by high and sustained growth. It was also the last time that the country was governed by liberal governments, in the classical, economic sense. During the latter part of the 1980s and the 1990s, Bolivia once again was governed by (largely) progressive, semi-liberal governments.

[2] Unit roots are important in determining the stationarity of a time series because the presence of non-stationary regressors invalidates many standard hypothesis tests. Among other things, the F-statistic calculated from a regression involving nonstationary time-series data does not follow the standard distribution, hence the significance of the test is overstated and spurious results are obtained.

the presence of unit roots. Several tests for the presence of unit roots in time series have appeared in the literature (see, for instance, Dickey and Fuller, 1979, 1981; Phillips and Perron, 1988, and Kwiatkowski et al. 1992). In this study, unit roots are tested using both the Augmented Dickey-Fuller (ADF) and the Phillips and Perron (PP) tests.

Once a unit root has been confirmed for a data series, the question is whether there exists some long-run equilibrium relationship among variables. The existence of a long-run equilibrium relationship among economic variables is referred to as cointegration.

According to Engle and Granger (1987), a set of variables, Y_t, is said to be cointegrated of order (d,b), denoted CI(d,b), if Y_t is integrated of order d and there exists a vector, β, such that βY_t is integrated of order $(d-b)$. Cointegration tests are conducted using the method of Johansen (1988) and Johansen and Juselius (1990).

The Johansen method applies the maximum likelihood principle to determine the presence of cointegrating vectors in nonstationary time series. Following Johansen (1988) and Johansen and Juselius (1990), a two-dimensional (2×1) vector autoregressive model is employed with Gaussian errors given by

$$Y_t = A_1 Y_{t-1} + A_2 Y_{t-2} + \dots + A_k Y_{t-k} + \mu + \varepsilon_t,\ t = 1, 2, \dots, T, \tag{1}$$

where Y_t is, in turn, real GDP (real GDP per capita) and the financial development or trade openness indicators, and ε_t is i.i.d. $N(0,\Sigma)$. After first-differencing on the vector level, the model in error correction form is written as:

$$\Delta Y_t = \Gamma_1 \Delta Y_{t-1} + \Gamma_2 \Delta Y_{t-2} + \dots + \Gamma_{k-1} \Delta Y_{t-k+1} - \Pi Y_{t-1} + \mu + \varepsilon_t, \tag{2}$$

where $\Gamma_i = -I + A_1 + A_2 + \dots + A_i$, for $i = 1, 2, \dots, k$-1, and $\Pi = I - A_1 - A_2 - \dots - A_k$.

The Π matrix conveys information about the long-run relationship between the Y_t variables, and the rank of Π is the number of linearly independent and stationary linear combinations of real GDP (real GDP per capita) and the financial development and trade openness indicators.

Testing for cointegration involves testing for the rank, r, of the Π matrix, which is achieved by examining whether the eigenvalues of Π are significantly different from zero. The test has three possible outcomes: (a) the Π matrix has full column rank implying that the Y_t was stationary in levels, (b) the Π matrix has zero rank, in which case the system is a traditional first-differenced VAR, and (c) the Π matrix has rank r with $0 < r < 2$, implying that there exists r linear combinations of Y_t that are stationary or cointegrated. If condition (c) prevails, Π can be decomposed into two $2\times r$ matrices, α and β, such that $\Pi = \alpha\beta'$. The β vectors represent the r linear cointegrating relationships such that $\beta' Y_t$ is stationary. By testing the significance of the β-coefficients, it can be determined whether the variables are entering the cointegrating relationship significantly. The loading matrix, α, represents the error-correction mechanism and can be interpreted as speed of adjustment parameters.

Johansen (1988) and Johansen and Juselius (1990) propose two statistics for the determination of the number of cointegrating vectors, or, equivalently, the rank of Π: the trace statistic (T_r) and the maximum eigenvalue statistic. In this work, the trace statistic is utilized.[3] Its likelihood ratio statistic is:

$$T_r = -T\sum_{i=r+1}^{p=2} \ln(1-\lambda), \tag{3}$$

where $\lambda_{r+1},\ldots\lambda_p$ are the estimated $p - r$ smallest eigenvalues. The null hypothesis to be tested is that there are at most r cointegrating vectors. That is, the number of cointegrating vectors is less than or equal to r, where r is 0 or 1. In each case, the null hypothesis is tested against the general alternative.

It is well known that cointegration tests are very sensitive to the choice of lag length. Here, the Schwartz Criterion (SC) is used to select the number of lags required in the cointegration test. The SC is defined as follows:

$$\mathrm{SC} = \ln(y'My/T) + K\ln T/T, \tag{4}$$

where $M = 1 - X(X'X)^{-1}X'$, T is the sample size. Here, K_I is chosen so as to numerically minimize SC.

Engle and Granger (1987) demonstrate that if two non-stationary variables are cointegrated, then a vector autoregression in first differences is misspecified. Hence, cointegration must be tested before running causality tests. Tano (1993) proposes the use of cointegration and error-correction modeling (ECM) in Granger causality models because of the possibility of spurious co-movement of variables, while Granger (1986) and Engle and Granger (1987) have proposed the error-correction model (ECM) as a more comprehensive method for testing causality when variables are cointegrated. Cointegration analysis attempts to identify conditions under which existing relationships are not spurious. Unlike standard Granger causality which

[3] Even though the trace statistic is the one utilized in this work, the number of cointegrating equations according to the maximum eigenvalue test are also reported.

may not detect any causal relationship between variables, with ECM, cointegration ensures that Granger causality exists, at least in one direction.

The cointegrated ECMs of real GDP (real GDP per capita) with financial development or trade openness indicators are as follows:

$$\Delta Y_{1t} = \theta_{11}^{m}(L)\Delta Y_{1t} + \theta_{12}^{n}(L)\Delta Y_{2t} + \propto ECT_{1t-1} + a + \mu_{1t}, \tag{5}$$

$$\Delta Y_{2t} = \theta_{21}^{m}(L)\Delta Y_{1t} + \theta_{22}^{n}(L)\Delta Y_{2t} + \beta ECT_{2t-1} + b + \nu_{2t}, \tag{6}$$

where

$$\theta_{ij}^{m}(L) = \Sigma_{l=1}^{Mij}\theta_{ijl}^{n}L^{l} \text{ and } \theta_{ij}^{n}(L) = \sum_{l=1}^{Nij}\theta_{ijl}^{n}L^{l}, \tag{7}$$

and Δ is the first-difference operator, L is the lag operator, Y_{1t} and Y_{2t} are real GDP (real GDP per capita) and either financial development or trade openness indicators –which are first-differenced stationary time series– respectively, and, μ_{1t} and, ν_{2t} are disturbance terms without serial correlation, where E[μ_{1t}, μ_{2s}] = 0, E[ν_{1t}, ν_{2s}] = 0, E[μ_{1t}, ν_{2s}] = 0 for all $t \neq s$. $ECT_{it\text{-}1}$ is the error-correction term, lagged one period, which is derived from the long-run cointegrating relationship and included to capture short-run dynamics. The inclusion of these error correction terms, which must be stationary if the variables are cointegrated, differentiates the ECM from the standard Granger causality regressions.

On the basis of equations (5) and (6), unidirectional causality from Y_{2t} to Y_{1t} is implied if not only the estimated coefficients on the lagged Y_{2t} variables in equation (5) are statistically different from zero as a group (based on standard F-statistics), but also the coefficient on the error-correction term in equation (5) is significant, and if the set of estimated coefficients on the lagged Y_{1t} variables in equation (6) are not statistically different zero. On the other hand, Y_{1t} causes Y_{2t} if the estimated coefficients on the lagged Y_{1t} variable in equation (6) are statistically different from zero as a group, the coefficient on the error-correction term in equation (6) is significant, and if the set of estimated coefficients on the lagged Y_{2t} variables in equation (5) are not statistically different zero. Bidirectional causality or feedback between Y_{2t} and Y_{1t} exists if the set of estimated coefficients on the lagged Y_{2t} variables in equation (5) are statistically significant as a group, the set of estimated coefficients on the lagged Y_{1t} variables in equation (6) are statistically significant as a group, and the coefficients of error-correction terms in both equations are significant. As to long-run causality in equations (5) and (6), if the coefficient α in equation (5) (or β in equation 6) is significantly different from zero, it would indicate long-run causality from Y_{2t} to Y_{1t} (or from Y_{1t} to Y_{2t}).

As the Granger-causality tests are known to be very sensitive to lag length, some care must be taken when making this choice. Lag lengths are determined using Hsiao's (1979a, 1979b, 1981) sequential procedure. This procedure is based on the Granger definition of causality and Akaike's (1974) minimum final prediction error (FPE) criterion. The FPE criterion is specified as follows:

$$FPE = \left[\frac{T+k}{T-k}\right]\left(\frac{SSR}{T}\right), \tag{8}$$

where T is the number of observations, k is the number of parameters estimated, and SSR is the sum of squared residuals. Hsiao (1981) points out that 'the FPE criterion balances the risk due to the bias when a lower order is selected and the risk due to the increase of variance when a higher order is selected, and choosing the order of the lags by minimum FPE is equivalent to applying an approximate F-test with varying significance levels. This procedure is known as the stepwise Granger-causality technique.

However, if real GDP (real GDP per capita) and the financial development and trade openness indicators are found to be not cointegrated, then the intertemporal causality relationships in equations (5) and (6) must be estimated without the error-correction terms.

III. Data

This study analyzes Bolivia and the impact that financial development and trade openness have had on the country's economic growth. As was explained in Section II, the indicators of economic growth are tested against indicators of financial development and trade openness to account for a long-run relationship. Once a long-run relationship has been established (or rejected), the issue of causality is explored to determine whether the economic growth indicators are the result of financial development and trade openness indicators, whether causality runs the other direction, or whether causality between these variables cannot be determined.

Though the term 'economic growth' can be understood in different ways, in this work the two most commonly indicators of this concept are utilized: real GDP and real GDP per capita. The financial development indicators are the ratio of M2 (money in circulation) to GDP, also known as the Finance Depth Ratio (FDR), and the ratio of M2 minus currency to GDP, as suggested by Demetriades and Hussein (1996). The FDR indicator is the one most commonly used in the literature to account for financial development; the second is a slightly adjusted indicator and it represents an improvement over the traditional FDR indicator in that, by subtracting currency from M2, a larger ratio is avoided, which may have been the result of more intensive utilization of currency rather than an increase in the volume of deposits. The trade openness indicators are the ratio of total exports and imports to GDP, and the ratio of exports and imports of Bolivia's four most important economic partners to GDP.[4] The significance of the trade and financial indicators rests on their perceived relation with the chosen indicator of economic growth. High finance ratios may indicate relatively well developed financial markets, or, in times of high inflation, may serve as an indicator of short-run finance of government programs. High trade indicators may indicate a relatively open economy, and hence one with competition and a more efficient assignment of resources.

The empirical analysis uses annual data during the 1940-2010 period on real GDP, real GDP per capita, and the ratios of real total trade, real trade of the four biggest economic partners, real M2, and real M2 minus currency to GDP.[5] The base year for all variables is 2000. All data have been obtained from the Statistical bulletins and the Annual Reports of the Central Bank of Bolivia (www.bcb.gob.bo). Finally, all data series have been transformed to the logarithmic form to achieve stationarity in variance.

The four empirical versions analyzed for the two economic growth indicators are the following:

$$\begin{array}{ll} \text{(a) } lrgdp = f(lrtotaltrade/GDP), & \text{(e) } lrgdp/capita = f(lrtotaltrade/GDP), \\ \text{(b) } lrgdp = f(lrtradefour/GDP), & \text{(f) } lrgdp/capita = f(lrtradefour/GDP), \\ \text{(c) } lrgdp = f(lrM2/GDP), & \text{(g) } lrgdp/capita = f(lrM2/GDP), \\ \text{(d) } lrgdp = f(lrM2\text{-}currency/GDP), & \text{(h) } lrgdp/capita = f(lrM2\text{-}currency/GDP), \end{array} \tag{9}$$

[4] For the period under consideration, the four biggest trade partners for Bolivia are the United States, United Kingdom, Argentina, and Brazil.

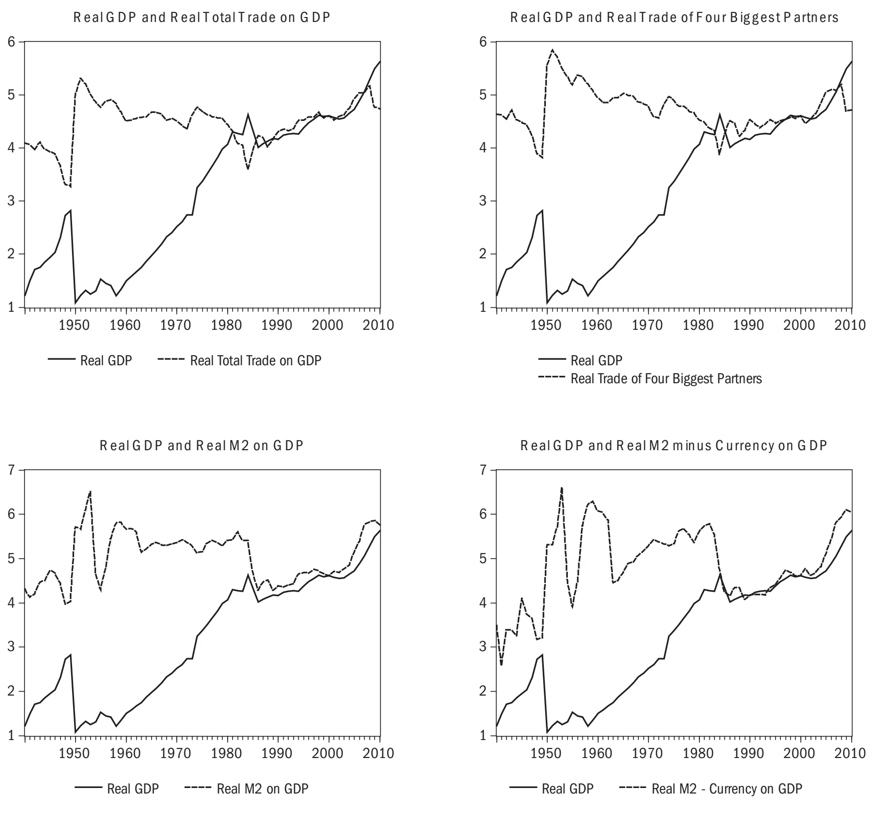
[5] M2 comprises the sum of currency and coins in circulation; checking accounts in domestic currency (Bolivianos); checking accounts in Bs adjusted to the value of US dollars; checking accounts in UFVs (unit of account in Bs adjusted to inflation and created to induce housing ownership); saving accounts in Bs; saving accounts in Bs adjusted to the value of US dollars; and saving accounts in UFVs. Other measures of financial development, such as private credit/GDP or an indicator of stock market development, were not used due to lack of reliable, continuous data. Values for 2010 are preliminary.

where *lrgdp* = log real GDP, *lrgdp/capita* = log real GDP per capita, *lrtotaltrade/GDP* = log real total trade over GDP, *lrtradefour/GDP* = log real trade of four biggest economic partners over GDP, *lrM2/GDP* = log real M2 over GDP, and *lrM2-currency/GDP* = log real M2 minus currency over GDP.

IV. Results

A preliminary exploration of the data is presented in Figure 1, which illustrates how economic growth, as measured by real GDP, relates to the financial development and trade openness indicators over the period of interest.

Figure 1. Economic growth, financial and trade openness indicators, 1940-2010



Generally, there seems to be a high degree of correlation in all specifications. However, the correlation seems stronger in all cases after the hyperinflation years of the early 1980s. Pairwise correlations reported in Table 1 confirm the graphical intuition as all variables are indeed correlated.

The correlation values shown on Table 1 correspond to the entire period under analysis. It is interesting to note that the correlation between economic growth –be it real GDP or real GDP per capita– and trade and financial development is, with the exception of the finance indicator *lrM2-currency/GDP*, always negative. However, when the same correlation values are estimated for the period 1986-2010 –that is, the time after the hyperinflation period– all correlation values are positive and above 0.60, denoting the highly positive influence that seems to exist between the variables.

Because cointegration equations require the use of nonstationary variables and error-correction equations require the use of stationary variables, each data series is first examined for the probable order of difference stationarity. Table 2 reports the results of nonstationary tests for real GDP, real GDP per capita, real total trade, real trade of four biggest economic partners, real M2, real M2 minus currency, real total trade on GDP, real trade of four biggest economic partners on GDP, real M2 on GDP and real M2 minus currency on GDP, using the augmented Dickey Fuller test (ADF test) and the Phillips and Perron test (PP test).[6]

A constant is included but no time trend, as recommended by Dickey et al. (1986). According to the test statistics by MacKinnon (1991), nonstationarity cannot be rejected for the levels of all variables at the 5% significance level.[7] In contrast, when the data are differenced, nonstationarity can be rejected for all data series studied. This indicates that all the data series are integrated of order one, or I(1). Table 2 also reports the minimum AIC(n), which indicates the appropriate order of the autoregressive process, n. The results from the PP test generally support the ADF test and confirm that all the data series are integrated of order one.[8]

Once unit roots have been confirmed for all data series, it remains to be determined whether there exists some long-run equilibrium relationship between real GDP (real GDP per capita) and the financial development and trade openness indicators.

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[6] As one reviewer correctly noted, the finance and openness ratios are likely to be stationary by default, as they are bounded within closed intervals. To account for this, the variables used to estimate those ratios are also tested for unit roots.

[7] The variable for real M2 over GDP is stationary in levels at the 5% level of significance.

[8] According to the PP test, the variables for real trade of four biggest partners over GDP and real M2 over GDP are stationary in levels at the 5% level of significance.

Table 1. Pairwise correlations

	Real GDP (*lrgdp*)	Real GDP per capita (*lrgdp/capita*)	Real total trade over GDP (*lrtotaltrade/GDP*)	Real trade four biggest partners over GDP (*lrtradefour/GDP*)	Real M2 over GDP (*lrM2/GDP*)	Real M2 minus currency over GDP (*lrM2-currency/GDP*)
Real GDP (*lrgdp*)	1.000					
Real GDP per capita (*lrgdp/capita*)	0.986	1.000				
Real total trade over GDP (*lrtotaltrade/GDP*)	-0.020	-0.149	1.000			
Real trade four biggest partners over GDP (*lrtradefour/GDP*)	-0.466	-0.558	0.868	1.000		
Real M2 over GDP (*lrM2/GDP*)	-0.116	-0.174	0.610	0.592	1.000	
Real M2 minus currency over GDP (*lrM2-currency/GDP*)	0.093	0.019	0.641	0.496	0.919	1.000

Table 2. Unit root tests

	Augmented Dickey-Fuller test				Phillips and Perron test	
	Level	AIC (*n*)	First difference	AIC (*n*)	Level	First difference
Real GDP	0.0025	0.1913	-4.8797**	0.2070	-0.3080	-8.1358**
Real GDP per capita	-0.4177	0.1822	-4.8925**	0.2008	-0.7488	-8.0983**
Real total trade	0.1119	-0.8811	-4.7958**	-0.8976	0.0250	-7.0902**
Real trade of four biggest partners	0.2203	-0.7587	-4.9108**	-0.7645	-0.0257	-7.3378**
Real M2	-0.4108	0.6915	-5.1195**	0.7050	-0.6235	-7.0058**
Real M2 minus currency	-0.8200	12.635	-5.2158**	1.2788	-0.9453	-7.8037**
Real total trade over GDP	-2.8964	0.0420	-5.9377**	0.1501	-2.8788	-7.0907**
Real trade of four biggest partners over GDP	-2.7607	0.0834	-5.9751**	0.1837	-2.9519*	-7.2600**
Real M2 over GDP	-3.0792*	0.7873	-5.4845**	0.9333	-3.0311*	-7.3469**
Real M2 minus currency over GDP	-2.7672	1.4398	-5.2383**	1.5637	-2.7166	-8.0597**

Note: ** and * denote significance at the 1% and 5% level, respectively

Table 3: Cointegration test results for real GDP and real GDP per capita

	Real GDP				Real GDP per capita		
	Trace statistic	Critical value (5%)	Critical value (1%)		Trace statistic	Critical value (5%)	Critical value (1%)
1) Real total trade over GDP (VAR lag = 1)				1) Real total trade over GDP (VAR lag = 1)			
H_0: r = 0	22.14*	19.96	24.6	H_0: r = 0	16.79	19.96	24.6
H_0: r ≤ 1	10.08*	9.24	12.97	H_0: r ≤ 1	5.53	9.24	12.97
2) Real trade of four biggest partners over GDP (VAR lag = 1)				2) Real trade of four biggest partners over GDP (VAR lag = 1)			
H_0: r = 0	21.88*	19.96	24.6	H_0: r = 0	17.35	19.96	24.6
H_0: r ≤ 1	8.44	9.24	12.97	H_0: r ≤ 1	4.09	9.24	12.97
3) Real M2 over GDP (VAR lag = 1)				3) Real M2 over GDP (VAR lag = 1)			
H_0: r = 0	18.94*	18.17	23.46	H_0: r = 0	18.85*	18.17	23.46
H_0: r ≤ 1	4.35*	3.74	6.4	H_0: r ≤ 1	4.47*	3.74	6.4
4) Real M2 minus currency on GDP (VAR lag = 1)				4) Real M2 minus currency over GDP (VAR lag = 1)			
H_0: r = 0	17.43	18.17	23.46	H_0: r = 0	17.44	18.17	23.46
H_0: r ≤ 1	4.39*	3.74	6.4	H_0: r ≤ 1	4.57*	3.74	6.4

Notes: * and ** denote significance at the 5% and 1% level, respectively *r* denotes the number of cointegrating vectors specifications 1, 2 follow a non-deterministic trend and specifications 3, 4 follow a quadratic, deterministic trend.

Following the Johansen and Juselius method, a VAR model is estimated first to find an appropriate lag structure. The Schwartz Criterion (SC) suggests one lag for most of the VAR models examined.

Table 3 presents the results from the Johansen cointegration tests for real GDP and each of the financial development and trade openness indicators.

Cointegration results for three of the four specifications demonstrate that there is indeed a long-run equilibrium relationship between economic growth, as measured by real GDP, both trade openness indicators (real total trade over GDP and real trade of four biggest partners over GDP), and real M2 over GDP. The trace statistics for specifications 1 through 3 produce results suggesting two cointegrating equations (specifications 1 and 3) and one cointegration equation (specification 2) at the 5 percent level of significance.[9] A long-run equilibrium relationship between real GDP and real M2 minus currency over GDP is not found.

Table 3 also presents the results from the Johansen cointegration tests for real GDP per capita and each of the financial development and trade openness indicators. Cointegration results between this alternative indicator of economic growth and the financial development and trade openness indicators are less conclusive. There only seems to be a long-run equilibrium relationship between real GDP per capita and real M2 over GDP, but the other three specifications do not produce evidence of cointegration.[10]

Depending on the results of the cointegration tests, two approaches to testing causality are followed. If the variables are not cointegrated, causality tests are conducted based on the standard Granger regressions, that is, Equations (5) and (6) without the error-correction terms. If, on the other hand, cointegration is found to exist, then error-correction models, as indicated in Equations (5) and (6), can be estimated as a basis for determining causality. Based on the cointegration results, causality is investigated using the standard Granger regressions for real GDP and real M2 minus currency over GDP, real GDP per capita and real total trade over GDP, real GDP per capita and real trade of four biggest partners over GDP, and real GDP per capita and real M2 minus currency over GDP. Causality is tested based on the error correction model for real GDP and real total trade over GDP, real GDP and real trade of four biggest partners over GDP, real GDP and real M2 over GDP, and real GDP per capita and real M2 over GDP. Table 4 presents the main findings for these causality tests.

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[9] The maximum eigenvalue test for all specifications indicate no cointegration equations at the 5% and 1% levels of significance.

[10] The trace test indicates the presence of 2 cointegration equations at the 5% level of significance. The maximum eigenvalue test indicates no cointegration equations for this specification.

Table 4. Test statistics and probabilities for causality tests

(a-b) Real GDP (*lrgdp*) and real trade over GDP (*lrtotaltrade/GDP*, *lrtradefour/GDP*)				
Cointegrated, hence causality using ECM model: *lrtotaltrade/GDP* causes *lrgdp*				
Lrgdp	1	-0.0079		
		(-0.09)		
lrtotaltrade/GDP	-126.7684**	1		
	(-3.03)			
Cointegrated, hence causality using ECM model: *lrtradefour/GDP* causes *lrgdp*				
Lrgdp	1	0.1086		
		(1.43)		
lrtradefour/GDP	9.2115**	1		
	(3.57)			
(c-d) Real GDP (*lrgdp*) and real M2 over GDP (*lrM2/GDP*, *lrM2-currency/GDP*)				
Cointegrated, hence causality using ECM model: *lrM2/GDP* causes *lrgdp*				
Lrgdp	1	0.1297		
		(1.08)		
lrM2/GDP	7.7116**	1		
	(2.55)			
Not cointegrated, hence causality using traditional Granger test			F-Statistic	Probability
lrM2-currency/GDP does not Granger cause *lrgdp*			2.3943	0.0994
lrgdp does not Granger cause *lrM2-currency/GDP*			0.3954	0.675
lrM2-currency/GDP Granger causes *lrgdp*				
(e-f) Real GDP per capita (*lrgdp/capita*) and real trade over GDP (*lrtotaltrade/GDP*, *lrtradefour/GDP*)				
Not cointegrated, hence causality using traditional Granger test			F-Statistic	Probability
lrtotaltrade/GDP does not Granger cause *lrgdp/capita*			4.281	0.018
lrgdp/capita does not Granger cause *lrtotaltrade/GDP*			0.0785	0.9246
lrtotaltrade/GDP Granger causes *lrgdp/capita*				
Not cointegrated, hence causality using traditional Granger test			F-Statistic	Probability
lrtradefour/GDP does not Granger cause *lrgdp/capita*			3.5409	0.0348
lrgdp/capita does not Granger cause *lrtradefour/GDP*			0.8251	0.4428
lrtradefour/GDP Granger causes *lrgdp/capita*				
(g-h) Real GDP per capita (*lrgdp/capita*) and real M2 over GDP (*lrM2/GDP*, *lrM2-currency/GDP*)				
Cointegrated, hence causality using ECM model: *lrM2/GDP* causes *lrgdp/capita*				
lrgdp/capita	1	0.1859		
		(1.1)		
lrM2/GDP	5.3792**	1		
	(2.57)			
Not cointegrated, hence causality using traditional Granger test			F-Statistic	Probability
lrM2-currency/GDP does not Granger cause *lrgdp/capita*			1.5109	0.2109
lrgdp/capita does not Granger cause *lrM2-currency/GDP*			0.6851	0.6052
Both null hypothesis are accepted: independent				

Notes: t-statistics in parenthesis. ** indicates significant at the 5% level of significance. The FPE criterion determined 1 lag length for all causality tests.

Several findings are worth to note. First, unidirectional Granger causality running from trade openness indicators to economic growth holds regardless of which economic growth indicator is used. Additionally, unidirectional Granger causality running from financial development indicators to economic growth holds, with the exception of the specification real GDP per capita and real M2 minus currency over GDP, where causality is indeterminate. Finally, in those instances where causality was found, it only ran in one direction, leaving little room for ambiguity –at least for the period analyzed– as to which variable Granger causes economic growth.

V. Conclusions

This study analyzes the impact of trade openness and financial development on the economic growth of Bolivia for the period 1940-2010. The analysis is an advance over previous studies in several respects. First, the hypothesis of a long-run relationship between indicators of economic growth and financial development, and economic growth and trade openness, is tested using bivariate cointegrated systems and employing the methodology of cointegration analysis suggested by Johansen (1988) and Johansen and Juselius (1990). Second, the issue of causality between the variables of interest is analyzed with standard Granger regressions and with error-correction models. Lastly, this paper analyzes a period of seventy years, the longest for a study of this kind on Bolivia. The results show that there is indeed a long-run equilibrium relationship between economic growth, financial development and trade openness indicators. Further, unidirectional Granger causality is found running from financial development and trade openness indicators to economic growth. These results are believed to be quite reliable due to the use of recently developed statistical methods for choosing the lag structure, the use of error-correction models to determine causality, and the utilization of a large dataset

The policy implications of these findings are particularly relevant today, as the current government is trying to revert many of the reforms that were painfully implemented during the 1980s and 1990s. If greater openness to the outside world and a healthy financial system are as important to affecting economic growth as this study demonstrates, then current attempts to close the economy do not seem reasonable. Furthermore, the recent nationalization of the two privately-run pension funds institutions cannot augur well to the health of the financial system. Hence, sooner or later, the repercussions of such ill-advised policies are to be reflected on Bolivia's economic growth.

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