Export-output Causality in Bolivia during the Years of

by

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March 2009

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Keywords – South America; Export-led growth; Export instability; Stabilization policies; Regional markets

JEL classification code – F14, C32, F43
Export-output Causality in Bolivia

Abstract. As a small open economy, Bolivia provides a good example of a developing country trying to achieve economic growth by pursuing outward-oriented policies. After an overview of the economy, we examine the causal relationship between output and three measures of export revenues from agricultural, manufacturing, and mining sectors. Our findings suggest that export revenues play an important role in determining current aggregate activity but have not yet been an engine for future growth. Given these findings, we discuss the strength and limitations of outward-oriented (and macroeconomic stabilization) policies for promoting economic activity as well as the importance of preferential export agreements in regional and US markets.

1. Introduction

Many developing economies have long realized that they cannot solely rely on their domestic markets to achieve adequate and sustainable economic growth. Following the economic success of newly-industrialized countries, these economies have been implementing an export-led growth strategy, looking for markets beyond their national borders to attain the long-needed growth in economic activity. Unfortunately, trade has not yet been a source of growth for many of these economies. Reasons cited include the heavy dependence on the export of primary products with inherent supply instability and historically volatile world prices, the lack of a viable manufacturing sector strong enough to compete in the global export market, and the heavy dependence on preferential treatment in regional markets that are subject to both economic and political instability.

A case in point is Bolivia. With a population of 8.5 million and real GDP per capita of $943 in 2004, Bolivia has a poverty index similar to some sub-Saharan countries. Growing at only 2.9% average annual rate for 2000-2005, the recent economic performance of Bolivia is not encouraging either. Obstacles to growth include (i) limited physical infrastructure, (ii) lack of necessary physical capital, (iii) lack of a well-educated and trained labor force, (iv) slow path of
technological progress, (v) large size of the informal sector, (vi) existence of a highly dollarized financial system, (vii) lack of well-developed financial institutions, and (viii) fragile institutional and legal system.¹ More specifically, investing in physical and human capital as well as technology has been limited by the increasing budget deficit which reduces the availability of loanable funds for private investment, and by the inability of financial institutions to efficiently channel savings into the hands of investors. A large portion of the Bolivian economy is informal, with 95% of output produced outside the legal system. This makes it difficult for the government to collect the necessary tax revenue to improve its fiscal situation. In addition, industries operating within the informal sector tend to lack the necessary documents for securing credit for investment. As a result, banks often impose excessive collateral requirements for making loans. This problem is reinforced by the fact that the Bolivian economy is highly dollar-based, and most loans are made in dollars. Companies conducting business in the domestic currency are considered by lenders as risky (due to the perceived exchange rate risk), justifying the excessive collateral requirements for loans and high interest charges.²

To achieve the long-needed growth in economic activity, Bolivia has been implementing an export-led growth strategy. Despite being geographically landlocked with a poor transportation infrastructure, the share of export revenues in GDP has been rather large, with an annual average of 25% for 2000-2003. For this period, however, real exports revenue was growing at only 3.0% average annual rate. One of Bolivia’s most important foreign outlets is the Andean Market, in which its exports receive preferential treatment (i.e., Bolivia’s exports are

¹ Such obstacles also prevent Bolivia from efficiently utilizing such abundant resources as natural gas and fertile soil. Natural gas reserves in Bolivia are estimated at 56 trillion cubit feet, second in South America after Venezuela.

² For 2002, the interest rates were 20.6% and 12.1%, respectively, on loans made in Bolivian pesos and in US dollars. For the same year, the Bolivian peso was depreciated by 8.5%.
subject to very limited, if any, tariff or non-tariff barriers). The Andean trading countries are Bolivia, Colombia, Ecuador, Peru, and Venezuela. Colombia (which is in the process of signing a Free Trade Agreement, FTA, with the US) has announced that in line with this agreement, it will soon end preferential treatment for Bolivia’s exports. Peru has already signed an FTA with the US and thus will have to end preferential treatment for Bolivia’s exports. Another concern is the non-renewal of the Andean Pact Trade and Drug Enforcement Agreement (APTDZA). This agreement between the US and the Andean countries allows over 6,000 Bolivian products to enter the US with preferential treatment. Its non-renewal by the U.S. Government may adversely affect Bolivia’s ability to compete in US markets.

Given this overview, we set out to examine the causal relationship between real GDP and three measures of real export revenues from agricultural, manufacturing, and mining sectors in Bolivia. In so doing, we utilize the procedure developed by Haugh (1976) and Feige and Pearce (1976). The test results suggest that these export revenues do not cause real GDP, and real GDP does not cause the alternative measures of real exports. Despite the lack of causality, a strong and significant contemporaneous relationship between real GDP and export revenues is detected. That is, export revenues play an important role in determining current aggregate activity but have not yet been an engine for future growth and employment in Bolivia.

2. Theoretical and Empirical Background

Three alternative hypotheses characterize the dynamic relationship between real output and real exports: the export-led growth hypothesis views real exports as the engine of economic growth, the growth-driven export hypothesis views economic growth as the driving force for export expansions, and the last hypothesis sees a bi-directional causation between output and exports. The export-led growth hypothesis basically maintains that the expansion of exports and
foreign competitive pressures improve resource allocation, facilitate technological progress, enhance adaptability and managerial efficiency, promote economies of scale, and create positive externalities for various sectors through specialized inputs (Keesing, 1967; Balassa, 1978; Krueger, 1980; Helpman and Krugman, 1985; Grossman and Helpman, 1991; and Greenaway and Sapsford, 1994a). The model by Rivera-Batiz and Romer (1991), for instance, indicates that the expansion of international trade boosts economic growth by increasing the number of specialized inputs. Export expansion also helps ease binding foreign exchange constraints and improve the country’s ability to import the necessary capital and intermediate goods unavailable domestically (McKinnon, 1964).

In support of the growth-driven export hypothesis, the study by Kaldor (1967) emphasizes the positive impact of output growth on productivity growth through such factors as economies of scale, learning curve effects, increased division of labor, and the formation of new production methods and subsidiary industries. Improved productivity as the prime cause of output growth in the industrial sector, for instance, reduces the unit labor cost and makes it easier to compete and sell in the global export market (Kaldor, 1967, p. 40). Thus, this theory predicts a causal relationship from output growth to exports through productivity growth.

Trade theories incorporating economies of scale and imperfect competition suggest the possibility of a bi-directional causation between output and exports (see, for instance, Marin, 1992). The basic assertion underlying these theories is that export expansion is more likely to lead to productivity improvements if the entry of new firms causes competition, forcing inefficient firms to exit the market. Then, the exploitation of scale economies through export

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3 See the collection of essays in McCombie and Thirlwall (2004) and Kaldor (1989) on the importance of export growth in easing the balance of payments constraint.
expansion leads to productivity improvements, which, in turn, stimulate further economies of scale and cost reductions, making it easier for the few surviving firms to compete and sell in the global export market.

There are numerous empirical studies examining the relationship between output and exports for both developed and developing countries. In support of the export-led growth hypothesis, Marin (1992) provides evidence for Germany, Japan, the UK, and the US. The growth-driven hypothesis is, however, supported for Australia (Kunst and Marin, 1989), Canada (Henriques and Sadorsky, 1996), and Greece (Panas and Vamvoukas, 2002). Chow (1987) examines the causal relationship between exports of manufacturing goods and manufacturing output for newly-industrialized countries. His findings support a bi-directional causation for Hong Kong, Korea, Israel, Singapore, Taiwan, and Brazil. For developing countries, empirical evidence on the causal relationship between output and exports are also mixed. For instance, Balassa (1985) presents evidence in support of the export-led hypothesis for forty-three developing countries. However, this hypothesis is supported for only four (Indonesia, Egypt, Costa Rica, and Ecuador) out of thirty-seven developing countries examined by Jung and Marshall (1985). Bahmani-Oskooee et al. (1991) also find mixed results for twenty developing countries, with evidence in support of export-led growth particularly for the newly industrialized countries. In another study, Bahmani-Oskooee and Alse (1993) find further evidence of a bi-directional causality for eight developing countries.\(^4\)

3. Data and Methodology

In this study, we utilize quarterly data on real Gross Domestic Product (GDP) and real export revenues from agricultural, manufacturing, and mining sectors of the Bolivian economy for 1988.1-2002.4. The data are obtained from the Central Bank of Bolivia. Throughout this study, \( Y \) denotes real GDP, \( A, Mf \), and \( Mg \) denote, respectively, the real export revenues from the agricultural, manufacturing, and mining sectors. Finally, \( X \) denotes the sum of real export revenues from the three sectors.\(^5\)

Granger’s (1969) notion of causality is based on temporal ordering and incremental predictability criteria. That is, \( X \) Granger-causes \( Y \), if the past history of \( X \) contains information useful to predict future \( Y \). To empirically examine the causal relationship between \( X \) and \( Y \), we utilize the procedure suggested by Haugh (1976), and Feige and Pearce (1976). Specifically, Feige and Pearce (1976) point out that univariate models, such as the autoregressive integrated moving-average (ARIMA), efficiently utilize the information in the past history of the time-series, and, therefore, generate forecasts that can be considered weakly rational. In other words, the residuals from the ARIMA model of \( Y \), referred to as the pre-whitened \( Y \), represent the part of \( Y \) that cannot be predicted from its own past history. In order for \( X \) to Granger-cause \( Y \), the information available in the past history of \( X \) must be correlated with the pre-whitened \( Y \). If the pre-whitened \( Y \) happens to be orthogonal to such information in \( X \), then, we can conclude that \( X \) does not contain useful information necessary to improve the predictions of future \( Y \), and, therefore, \( X \) fails to Granger-cause \( Y \). Now the question is: How do we measure the information

\(^{5}\) The data on total export revenues is not available on a quarterly basis. However, we believe the sum, \( X \), is a good approximation of the total export revenues for the sample period under examination. There are also other limitations that we face in this study. First is the lack of available quarterly data (on the series examined here) after 2002. Second is the lack of quarterly data on other relevant macroeconomic variables which force us to restrict ourselves to the bivariate models presented in the text.
in the past history of $X$? Both Haugh (1976), and Feige and Pearce (1976) suggest measuring the information in the past history of $X$ by the residuals from the ARIMA model of $X$, referred to as the pre-whitened $X$. Utilizing the pre-whitened $X$, instead of $X$ itself, helps avoid misleading results due to the serial correlation present in $X$ (Feige and Pearce, 1976, 509-510).

We shall denote the error terms from the ARMA models of $X$ and $Y$, respectively, by $u_X$ and $u_Y$, and the population cross-correlation between $u_{X,t}$ and $u_{Y,t+k}$ by $\rho(u_{X,t} \ u_{Y,t+k})$. Then, $X$ and $Y$ are independent, if $\rho(u_{X,t} \ u_{Y,t+k}) = 0$ for all $k$. $X$ does not cause $Y$, if $\rho(u_{X,t} \ u_{Y,t+k}) = 0$ for all $k > 0$. $Y$ does not cause $X$, if $\rho(u_{X,t} \ u_{Y,t+k}) = 0$ for all $k < 0$. There is no causality in either direction, if $\rho(u_{X,t} \ u_{Y,t+k}) = 0$ for both $k > 0$ and $k < 0$. Finally, there is no contemporaneous relationship between $X$ and $Y$, if $\rho(u_{X,t} \ u_{Y,t+k}) = 0$ for $k = 0$.

Using the sample cross-correlation function estimates of $\hat{u}_{X,t}$ and $\hat{u}_{Y,t+k}$, denoted $r(\hat{u}_{X,t} \ \hat{u}_{Y,t+k})$, we perform two tests. The first test compares individual cross-correlations to their approximate standard errors. As indicated by Feige and Pearce (1976, p. 510), under the null hypothesis of independence, individual sample cross-correlations are each distributed normally with a mean zero and standard error $1/\sqrt{N}$, where $N$ is the sample size. Thus, an individual sample cross-correlation, $r(\hat{u}_{X,t} \ \hat{u}_{Y,t+k})$, is significantly different from zero, if $r(\hat{u}_{X,t} \ \hat{u}_{Y,t+k}) \geq 2/\sqrt{N}$. The second test is based on the $S^*$-statistic,

$$S^* = N^2 \sum_{k=-m}^{m} (N-k)^{-1} r^2(\hat{u}_{X,t} \ \hat{u}_{Y,t+k})$$

where $m = 1, 2, \ldots, M$, with $M$ defined as the maximum number of cross-correlations estimated for both $k < 0$ and $k > 0$, and $r^2(\hat{u}_{X,t} \ \hat{u}_{Y,t+k})$ is the square of the estimated cross-correlation coefficient at lag or lead $k$. Under the null hypothesis of independence, $S^*$ is distributed
approximately as $\chi^2(2M+1)$. We reject (i) the null hypothesis of independence if the calculated $S^*$ is greater than the 10% critical value of $\chi^2(2M+1)$, (ii) the null hypothesis that $X$ does not cause $Y$ if the calculated $S^*$ (for $k > 0$) is greater than the 10% critical value of $\chi^2(M)$, (iii) the null hypothesis that $Y$ does not cause $X$ if the calculated $S^*$ (for $k < 0$) is greater than the 10% critical value of $\chi^2(M)$, and (iv) the null hypothesis of no causality in either direction if the calculated $S^*$ (for $k \neq 0$) is greater than the 10% critical value of $\chi^2(2M)$. Finally, we reject the null hypothesis of no contemporaneous relationship between $X$ and $Y$ if the calculated $S^*$ (for $k = 0$) is greater than the 10% critical value of $\chi^2(1)$.

4. Empirical Results

Following the methodology described above, we now need to identify the ARIMA models for real GDP ($Y_t$) and for alternative real export revenues ($A_{t}, M_{ft}, M_{gt}$, and $X_t$). The general ARIMA model, say, for the logarithm of $z_t$, is written as

$$\varnothing_p(B)(1 - B)^d \log(z_t) = \theta_q(B)u_t$$

where $B$ is the backward operator; $(1 - B)^d$ is the $d$th difference of the series, with $d$ defined as the order of integration; $\varnothing_p(B)$ is a polynomial in $B$ with $p$ autoregressive roots; $\theta_q(B)$ is a polynomial in $B$ with $q$ moving-average roots; and $u_t$ is a white noise series. The transformation of the series to logarithms is intended to ensure homoscedasticity. Identification of the ARIMA model involves the determination of $d$, $p$, and $q$. In identifying $d$, we employ the augmented Dickey-Fuller (ADF) test equation (Dickey and Fuller, 1981) to test the null hypothesis of a unit root for both real GDP and alternative measures of real export revenues. The test equations include a time trend ($T$) and an adequate number of augmented lags to ensure the absence of autocorrelation in the error term. As seen in Table 1, the calculated ADF test statistics are all significant, indicating that the null hypothesis of a unit root is rejected in favor of trend-stationarity for all series.
With \( d = 0 \) for the logarithm of \( Y_t, A_t, Mf_t, Mg_t, \) and \( X_t \), the general ARIMA model reduces to the following general ARMA model,
\[
\Phi_p(B) z'_t = \theta_q(B) u_t,
\]
where \( z'_t = \log(z_t) - \alpha - \beta T \) is the de-trended series of \( \log(z_t) \). Using the estimates of the sample autocorrelation and partial autocorrelation functions of the de-trended series (in addition to the Schwarz information criterion), we have determined the autoregressive order \((p)\) and the moving-average order \((q)\). Table 2 reports the estimates of the ARMA models for all series. As indicated by the insignificant calculated Ljung-Box Q-statistics, the residual series are white noise, and thus the univariate ARMA models are all correctly specified.\(^6\)

The residual series from the univariate models in Table 2, \( \hat{u}_Y, \hat{u}_A, \hat{u}_{Mf}, \hat{u}_{Mg}, \) and \( \hat{u}_X \), are, respectively, the pre-whitened series of the logarithm of \( Y_t, A_t, Mf_t, Mg_t, \) and \( X_t \). The pre-whitened real GDP, \( \hat{u}_Y \), is then cross-correlated with each pre-whitened measure of real export revenues, \( \hat{u}_A, \hat{u}_{Mf}, \hat{u}_{Mg}, \) and \( \hat{u}_X \). Figure 1 illustrates the sample estimates of these cross-correlation functions. Inspection of the individual cross-correlations reveals very few significant spikes, with no pattern implying any kind of Granger-causality between real GDP and alternative export revenues. Based on the calculated \( S^*\)-statistics for \( k > 0 \), reported in Table 3, we cannot reject the individual null hypotheses that alternative measures of real export revenues do not Granger-cause real GDP. Also, based on the calculated \( S^*\)-statistics for \( k < 0 \), we cannot reject the individual null hypotheses that

\(^6\) For the estimates in Table 2, both the inverted autoregressive and moving-average roots are less than one. For example, for \( \log(Y) \), the inverted autoregressive root is 0.51 and the inverted moving-average roots are \( 0.65 \pm 0.27i, 0.11 \pm 0.83i, \) and \( -0.76 \pm 0.48i \); for \( \log(A) \), the inverted moving-average root is \(-0.53\); for \( \log(Mf) \), the inverted moving-average root is \(-0.64\); for \( \log(Mg) \), the inverted autoregressive root is 0.52 and the inverted moving-average roots are \( 0.89, 0.13 \pm 0.92i \) and \( -0.62 \pm 0.22i \); for \( \log(X) \), the inverted moving-average root is \(-0.56.\)
real GDP does not Granger-cause alternative measures of real export revenues.\(^7\) We thus conclude that the alternative measures of real export revenues are not leading indicators for real GDP, and real GDP is not a leading indicator for the alternative measures of real export revenues.\(^8\)

Despite the lack of causality in the sense of Granger (1969), the sample cross-correlation function estimates reveal a strong and significant contemporaneous cross-correlation between the pre-whitened series. For instance, for \(k = 0\), \(r(\hat{u}_{A,t}, \hat{u}_{Y,t}) = 0.68\), \(r(\hat{u}_{Mf,t}, \hat{u}_{Y,t}) = 0.71\), \(r(\hat{u}_{Mg,t}, \hat{u}_{Y,t}) = 0.76\), and \(r(\hat{u}_{X,t}, \hat{u}_{Y,t}) = 0.72\). With the significant calculated \(S^*\)-statistics (in Table 3), we reject the individual null hypotheses of no contemporaneous relationship between alternative measures of real export revenues and real GDP at the less than 1% level of significance.\(^9,10\)

5. Implications and Conclusions

The case of contemporaneous cross-correlation in Bolivia, while important, does not reveal any information on the direction of causality within a quarter. Thus, one can only speculate about the direction of the contemporaneous effect. Before doing so, however, consider the following points:

\(^7\) See Appendix A for a different approach in testing export-output causality.

\(^8\) These findings are consistent with Love and Chandra (2005) who report no causality in either direction for Pakistan and Sri Lanka. We may argue that our findings (for Bolivia) coupled with those of Love and Chandra (for Pakistan and Sri Lanka) give credence to the view that a minimum level of development is required before the benefits of export promotion on economic growth can be realized (see Helleiner, 1986; Greenaway and Sapsford, 1994b).

\(^9\) As seen, the calculated \(S^*\)-statistics for \(k \geq 0\) in Table 3 are all significant at the 1% level of significance, leading to the rejection of the null hypothesis of independence. We should, however, note that these results are mainly due to the strong contemporaneous relations.

\(^10\) The sample estimates of the cross correlation functions of (i) \(\hat{u}_A\) and \(\hat{u}_{Mf}\), (ii) \(\hat{u}_A\) and \(\hat{u}_{Mg}\), and (iii) \(\hat{u}_{Mf}\) and \(\hat{u}_{Mg}\), not reported here, reveal the absence of Granger-causality in either direction. Nonetheless, there still exist strong and significant contemporaneous relations among these series.
1. Agricultural products generally experience great short-run supply instability due to, for instance, climate changes and fluctuating world prices. Bolivia, with a small share of the world market for agricultural products, faces relatively elastic foreign demand curves for such products. Accordingly, supply instability tends to generate greater fluctuations in real agricultural export revenues (see Massell, 1970, p. 621 for an explicit illustration of this point).

2. Instability in real mining export revenues may, however, be due to demand-related factors. For instance, foreign demand for raw materials is often unstable because of cyclical changes in foreign incomes and short-run fluctuations in world prices.

3. Given the absence of effective government stabilization policies in Bolivia and the small size of the economy with a relatively large foreign sector, it is plausible to argue that the fluctuations in real GDP contemporaneously reflect the instabilities in real agricultural and mining export revenues.

4. The fluctuations in real GDP, however, create short-run instability in domestic demand (due to changes in purchasing power) for agricultural, manufacturing, and mining products. As indicated by Massell (1970, p. 622), instability in the domestic demand curve for a product produces instability in the net export supply curve, which, like supply instability discussed in point 1, creates instability in real export revenues.

5. Manufacturing products with higher short-run income elasticity tend to be much more vulnerable to instability in domestic demand than agricultural products (especially foods) with relatively low-income elasticity.

Based on the above points, we speculate (i) a bi-directional contemporaneous causation between real agricultural export revenues and real GDP, (ii) a bi-directional contemporaneous causation between real mining export revenues and real GDP, and a unidirectional
contemporaneous causation from real GDP to real manufacturing export revenues. Policymakers cannot do much about the external shocks that affect the swings in agricultural and mining export revenues (see points 1 and 2). However, they may implement policies to stabilize the swings in real output. If successful, the reinforcing feedback effects from real output to agricultural, manufacturing, and mining sectors will be reduced, lessening the swings in real export revenues (see points 4 and 5).

Ideally, government stabilization policies should aim at prolonging periods of economic expansion, shortening periods of recession, and reducing the gap between actual and potential output. Inspection of the time plot of real GDP (not shown here) indicates the absence of these characteristics in Bolivian business cycles. This may be due to the absence of effective fiscal and monetary policies. For instance, the ability of the government to implement expansionary fiscal policies is limited. Currently, government debt is above 60% of GDP. Around 40% of this is foreign debt at 1% annual rate of interest, and the rest, borrowed from domestic sources, is at 5.5% annual rate of interest. The annual interest payment on the government debt comes to 1.5% of GDP. This situation is not expected to get better, since government debt continues to run at over 60% of GDP in 2006. A major part of the Bolivian economy is informal and operates outside the legal system. This makes it difficult for the government to collect taxes. Accordingly, the establishment of a more equitable and transparent tax system to increase tax revenue and, thus, to reduce the budget deficit, is essential.

Bolivian monetary policies have also been ineffective in recent years. Expansionary monetary policies aimed at reducing interest rates, for instance, have failed to stimulate capital spending. As reported by the SBEF (2003), loans to businesses actually declined over 1998-2003, despite the significant drop in interest rates on loans (i) in US dollars from 14.6% in 1998
to 9.9% in 2003, and (ii) in Bolivian pesos from 31.3% in 1998 to 20.2% in 2003.\textsuperscript{11} Obstacles such as excessive collateral requirements are partly to blame. One reason for the excessive collateral requirements is the large size of the informal sector. Another reason is the domestic currency devaluation risk. The Bolivian peso, for instance, has depreciated year after year by an average annual rate of 6.8% since 1988. Accordingly, more than 60% of loans are made in US dollars, mainly because domestic companies (which conduct business in pesos) are perceived as risky by lenders. Therefore, in order to increase the effectiveness of monetary policies, Bolivia should aim at reducing the size of the informal sector, restructuring the financial system, and promoting the use of domestic currency.

Other policies should include diversification by improving factor mobility. For instance, adverse world economic conditions for mineral exports may be dealt with by diversifying into other sectors of the economy. If export revenues are falling because of lower foreign demand or lower world prices, factor mobility will allow diversification to take place by moving into more stable sectors of the economy. This is especially important for Bolivia, since her trading partners are mostly neighboring countries. For instance, Bolivia’s exports to the Andean Market (with Colombia, Ecuador, Peru, and Venezuela as trading partners) account for 17% of her total exports. In addition, her exports enjoy preferential treatment, since they are subject to very limited, if any, tariff or non-tariff barriers. Heavy dependence on such regional markets, however, makes Bolivia vulnerable to the economic and political conditions of its neighboring countries. As an example, Colombia’s Free Trade Agreement (FTA) with the US is in its last stages of negotiation and has announced that in line with this agreement, it will soon end preferential treatment for Bolivia’s exports. Also, Peru, which recently signed an FTA with the

\textsuperscript{11} SBEF stands for Superintendencia de Bancos y Entidades Financieras (Superintendent of Banks and Financial Entities).
US (and is also in the process of signing an FTA with the EU), will have to end preferential treatment for Bolivia’s exports. The loss of preferential treatment will make Bolivia’s exports more expensive and thus less competitive in both Colombia and Peru; Bolivia’s exports to Colombia and Peru currently represent 75% of her exports to the Andean Market.

It is thus essential for Bolivia to search for more diversified and more stable foreign outlets in order to expand beyond regional markets. The Andean Pact Trade and Drug Enforcement Agreement (APTDEA) provides such an opportunity. This agreement between the US and the Andean countries allows over 6,000 Bolivian products to enter the US with preferential treatment (Bolivia’s exports to the US currently account for 13.5% of her total exports). The primary aim of APTDEA has been to discourage Bolivia from producing coca, the raw material for cocaine. The previously-coca-producing regions now cultivate legal crops for export that enjoy preferential treatment under APTDEA. In addition, due to the large portion of textiles exported to the US with preferential treatment, some coca-producing peasants have relocated to other industrial areas to work in the textile sector for higher pay. Accordingly, the renewal of APTDEA by the US government is essential for encouraging further exports. In particular, losing APTDEA makes Bolivia’s exports more expensive and less competitive, and thus can potentially bring back the years when Bolivia was ranked number one in coca and cocaine production. To minimize the impact of losing APTDEA, policies should include labor retraining programs to improve labor mobility within the country, short-term unemployment benefits, and even export subsidies in order to keep Bolivian exports competitive in foreign markets.

Finally, our sample (1988.1-2002.4) in this study closely corresponds to the period of free market and trade policies. As we have shown for this period (which ended with the sacking of
President Sanchez de Lozada), export revenues play an important role in determining current aggregate activity but fail to be an engine for future growth. One should not, however, associate this failure with the implementation of free trade policies, for such policies require a minimum level of development before they become effective in advancing future economic growth. Accordingly, the current as well as future governments in Bolivia must aim at pursuing outward-oriented policies (coupled with the aforementioned stabilization policies) to meet the required threshold for exports to become an engine of future growth in real output and employment.
References


Appendix A.

In testing the causal relationship between real GDP and the alternative measures of real export revenues, we have utilized (in the text) the procedure developed by Haugh (1976) and Feige and Pearce (1976). This test has the advantage of detecting contemporaneous causation which is important to our analysis. The purpose of this Appendix is to emphasize that our conclusion of no Granger-causality in either direction remains unchanged when we use other test procedures. For instance, consider the following test equation,

\[ \log(Y)_t = a_0 + \sum_{i=1}^{m} a_i \log(Y)_{t-i} + \sum_{i=1}^{n} b_i \log(X)_{t-i} + \varepsilon_t \]

where, again, \( Y \) is real output and \( X \) is real (total) export revenues. Failure to reject the null hypotheses that \( b_1 = 0, b_2 = 0, \ldots, b_n = 0 \) individually and jointly indicates that real export revenues do not cause real output. As the first step, we have determined the optimal lag structure \((m)\) for \( \log(Y) \) using the Akaike information criterion (AIC) or the criterion of minimum final prediction error (Akaike, 1969). As the second step, we have added the first four lagged values of \( \log(X) \) and tested the null hypotheses that \( b_1 = 0, b_2 = 0, b_3 = 0, b_4 = 0 \) individually and jointly. The results from these tests lead to the conclusions that \( X \) does not Granger-cause \( Y \), since we cannot reject the aforementioned null hypotheses either individually or jointly (the test \( p \)-values are all above 0.10). We reach the same conclusions when we replace \( X \) in the test equation by other measures of export revenues (\( A, Mf, \) and \( Mg \)).

Similarly, using the test equation

\[ \log(X)_t = c_0 + \sum_{i=1}^{m} c_i \log(X)_{t-i} + \sum_{i=1}^{n} d_i \log(Y)_{t-i} + \varepsilon_t \]

we have examined the null hypothesis that real output \((Y)\) does not cause real export revenues \((X)\). Our test results indicate that \( Y \) does not Granger-cause \( X \), since we cannot reject the null
hypotheses that $d_1 = 0$, $d_2 = 0$, $d_3 = 0$, and $d_4 = 0$ either individually or jointly (the test $p$-values are all above 0.10). Again, we reach the same conclusions when we replace $X$ in the second test equation by other measures of export revenues ($A$, $Mf$, and $Mg$).
Table 1. Augmented Dickey-Fuller (ADF) tests of unit roots: 1988.1-2002.4

<table>
<thead>
<tr>
<th></th>
<th>log(Y)</th>
<th>log(A)</th>
<th>log(Mf)</th>
<th>log(Mg)</th>
<th>log(X)</th>
</tr>
</thead>
<tbody>
<tr>
<td>ADF</td>
<td>-4.09(^a)</td>
<td>-3.57(^a)</td>
<td>-3.48(^a)</td>
<td>-3.99(^a)</td>
<td>-3.15(^a)</td>
</tr>
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<td>Q[16]</td>
<td>15.8</td>
<td>5.83</td>
<td>11.2</td>
<td>10.6</td>
<td>8.38</td>
</tr>
</tbody>
</table>

Notes: Y is real GDP; A is real agricultural export revenues, Mf is real manufacturing export revenues; Mg is real mining export revenues; and X is the sum of real agricultural, manufacturing, and mining export revenues. The calculated ADF statistics, compared with Mackinnon's (1991) unit root critical values, test the null hypothesis of a unit root. The number of augmented lags included in the ADF test equation is six for log(Y), and three for log(A), log(Mf), log(Mg), and log(X). The adequacy of the lag length is checked with tests for serial correlation using the Ljung-Box Q-statistics with 16 degrees of freedom, Q[16]. Superscript \(^a\) indicates significance at the less than 10% level.
Table 2. Estimates of ARMA models: 1988.1-2002.4

1. Real GDP ($Y$):

\[
(1 - 0.510 B) \log(Y_t) = 7.184 + 0.008 T + (1 - 0.295 B^5 + 0.280 B^6) \hat{u}_{Y,t}
\]

\[
(0.113) \quad (0.074) \quad (0.002) \quad (0.136) \quad (0.135)
\]

\[R^2 = 0.62, \quad Q[16] = 8.33\]

2. Real agricultural export revenues ($A$):

\[
\log(A_t) = 5.473 + 0.005 T + (1+ 0.528 B) \hat{u}_{A,t}
\]

\[
(0.080) \quad (0.002) \quad (0.112)
\]

\[R^2 = 0.33, \quad Q[16] = 8.39\]

3. Real manufacturing export revenues ($Mf$):

\[
\log(Mf_t) = 3.551 + 0.026 T + (1+ 0.636 B) \hat{u}_{Mf,t}
\]

\[
(0.073) \quad (0.002) \quad (0.095)
\]

\[R^2 = 0.88, \quad Q[16] = 15.37\]

4. Real mining export revenues ($Mg$):

\[
(1-0.520 B) \log(Mg_t) = 5.374 + 0.0013 T + (1+ 0.562 B^4 + 0.365 B^5) \hat{u}_{Mg,t}
\]

\[
(0.118) \quad (0.035) \quad (0.001) \quad (0.083) \quad (0.086)
\]

\[R^2 = 0.88, \quad Q[16] = 10.11\]

5. Sum of real agricultural, manufacturing, and mining export revenues ($X$):

\[
\log(X_t) = 6.188 + 0.007 T + (1+ 0.556 B)\hat{u}_{X,t}
\]

\[
(0.066) \quad (0.002) \quad (0.110)
\]

\[R^2 = 0.88, \quad Q[16] = 11.87\]

Notes: See the notes in Table 1 for the definitions of the series; $T$ is a time trend. Values in parentheses are standard errors. $R^2$ is the coefficient of determination. The Ljung-Box $Q$-statistic, $Q[16]$, is for testing the absence of serial correlation up to the 16th order.
Table 3. Calculated $S^*$-statistics: 1988.1-2002.4

<table>
<thead>
<tr>
<th>$k &gt; 0$</th>
<th>$k &lt; 0$</th>
<th>$k \neq 0$</th>
<th>$k = 0$</th>
<th>$k \geq 0 &lt;$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$df = 16$</td>
<td>$df = 16$</td>
<td>$df = 32$</td>
<td>$df = 1$</td>
<td>$df = 33$</td>
</tr>
</tbody>
</table>

$\hat{u}_{A,t} \& \hat{u}_{Y,t+k}$  | 9.49   | 20.45  | 29.94  | 27.22$^a$  | 57.16$^a$  |
$\hat{u}_{Mf,t} \& \hat{u}_{Y,t+k}$ | 7.81   | 17.82  | 25.63  | 30.30$^a$  | 55.93$^a$  |
$\hat{u}_{Mg,t} \& \hat{u}_{Y,t+k}$ | 9.75   | 10.24  | 19.99  | 36.57$^a$  | 56.56$^a$  |
$\hat{u}_{X,t} \& \hat{u}_{Y,t+k}$  | 9.48   | 20.95  | 30.43  | 30.95$^a$  | 61.38$^a$  |

Notes: The maximum number of cross-correlations ($M$) in calculating $S^*$-statistics for both $K > 0$ and $K < 0$ is 16. Numbers for $K > 0$ relate to the null hypothesis that the respective export revenues do not Granger-cause real GDP; numbers for $K < 0$ relate to the null hypothesis that real GDP does not Granger-cause the respective export revenues; numbers for $K \neq 0$ relate to the null hypothesis of no Granger-causality in either direction; numbers for $K = 0$ relate to the null hypothesis of no contemporaneous effect between the respective export revenues and real GDP; numbers for $K \geq 0 <$ relate to the null hypothesis of independence. $df$ stands for the degrees of freedom. Superscript $^a$ indicates significance at the less than 10% level.
Figure 1. Cross-correlation function estimates for pre-whitened real GDP and alternative real export revenues: 1988.1-2002.4

1: A and Y
\[ r(\hat{u}_{A,t}, \hat{u}_{Y,t+k}) = 0.68 \]

2: Mf and Y
\[ r(\hat{u}_{Mf,t}, \hat{u}_{Y,t+k}) = 0.71 \]

3: Mg and Y
\[ r(\hat{u}_{Mg,t}, \hat{u}_{Y,t+k}) = 0.76 \]

4: X and Y
\[ r(\hat{u}_{X,t}, \hat{u}_{Y,t+k}) = 0.72 \]