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Policy Implications for the Application of Countercyclical Capital Buffers When the Government Borrowing Crowds Out Private Sector Credit: The Case of Jamaica

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Temporary Resource Booms and Manufacturing Output: A Global Perspective

Abstract

This paper analyzes the effect of temporary resource booms on manufacturing industry at a global level, but emphasizing the South-American case. The main conclusions are the following: first, the world is facing a boom of booms since 2002, in which South-America plays a prominent role; second, fuel and minerals booms are more likely to be larger and longer, and to generate more Dutch disease symptoms than capital flows or agricultural products booms, and third, the negative impact over the industry tends to last two and three years after the boom has ended.

Keywords: Resource booms, deindustrialization, Dutch disease, capital flights.

JEL classification: 013, 014, 016.

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1. INTRODUCTION

In the last decade South America has benefitted from significant capital flows on account of exports of natural resources and greater access to international financial markets, which has produced significant economic growth. Nevertheless, many of the concerns analysts have been voicing for some time now regarding the sustainability of this driver of growth in an environment of reduced international liquidity and lower commodity prices have begun to materialize. One of the main questions is the role of manufacturing industry in this new environment and its potential for offsetting lower revenues from natural resources and capital.

The main question addressed by this paper is therefore whether the end of booms will be accompanied by a readjustment in relative prices (or depreciation) that might contribute to a fast recovery in manufacturing output, or in other potential export sectors, that partly offsets the fall in revenues generated by booms. Another question is whether the characteristics and consequences of booms vary according to the type of boom (agricultural products, fuel and minerals, or capital) countries have enjoyed. To answer these questions we identify the main natural resource and capital boom and post-boom periods that have occurred at a global level, and particularly in South America; describing them and establishing the effects they have had on manufacturing industry according to the sector they occurred in.

The impact of revenues associated to natural resources on manufacturing and the overall behavior of economies has been widely analyzed in economic literature. The corresponding studies can be divided into three main groups. The first group revolves around the idea of a secular decline in the terms-of-trade for commodities originally proposed by Prebisch (1959) and Singer (1950). This idea was severely questioned by later studies (e.g., Cuddington, 1992) but has been taken up again recently by Ocampo and Parra (2010) and Erten and Ocampo (2013), who not only study trends of price series, but also their cyclical components.

The second group of studies deals with the effect of so-called Dutch disease, where the works of Corden and Neary (1982), and Ismail (2010) stand out. The latter find important relations between commodity booms, the real exchange rate and poor performance in the manufacturing sector. In the same way, Spatafora and Warner (1995) identify a very strong relation between the effects of terms-of-trade and the real exchange rate. Another version of this hypothesis is that put forward by Krugman (1987), in which he highlights the long-term effects that can stem from a temporary overvaluation of the exchange rate on models with dynamic scale economies and endogenous learning processes (*learning by doing*).

The third group of works, in many ways complementary to the previous one, is based around the theory of "the curse of natural resources" proposed by Sachs and Werner (1995, 1997), in which the opportunity for technical advances in the production of primary products is limited as compared to those generated by the manufacturing industry. These works also emphasize the negative impact that revenues associated to the production of primary products normally have on the institutions and economic policy of countries that are overly reliant on them (Besley et al., 2013). This group would also include the recent *Industrial Development Report* of the UNIDO (2013), which shows that countries rich in natural resources (minerals and hydrocarbons) exhibit lower industrial development (especially in industries that are key for growth in medium-developed countries, such as electronic products, automobiles and chemicals).

Several of the abovementioned approaches highlighting the potentially negative impact on countries of revenues associated to natural resources have been challenged by works including a report by the World Bank from 2001 (De Ferranti et al., 2001) and the recent work of Cieplan (Meller et al., 2013), which emphasize instead the enormous possibilities offered by the availability of such resources. In any case, although there is no complete agreement on the long-term implications of natural resource booms on economies, there is some agreement on the fact that, if the necessary measures are not adopted, flows of extraordinary revenues to a country will cause an appreciation in the exchange rate that affects tradable goods production, including those produced by the manufacturing industry (World Bank, 2010).[1]

It is also worth mentioning that, in line with the viewpoint of Corden and Neary (1982), revenues stemming from capital flows can have a revaluation effect that has a negative impact on manufacturing output over the long-term. In this vein, Lartey (2008) uses a model of business cycles to study the effect of capital flows on resource allocation and real exchange rate movements in emerging economies, finding that an increase in capital flows causes an increase in the demand for non-tradable goods, which translates into an appreciation of the exchange rate and a loss of international competitiveness. Thus, Athukorala and Rajapatirana (2003) also find that capital flows other than from foreign direct investment (FDI) are related to an appreciation of the exchange rate. However, the literature recognizes a certain ambiguity regarding this result because capital flows also allow for financing investment and current account deficits, favoring manufacturing output. In this regard, Kamar et al. (2010) find that FDI flows have a neutral impact on competitiveness, which in some cases can even be positive.

The approach proposed in this paper differs from the traditional Dutch disease discussion for at least three reasons. First, it does not limit itself to the problems that might be generated by revenues from natural resources and encompasses revenues associated to capital flows. Second, it not only includes price booms, but also those of quantity.[2] Third, it does not concern itself with the advantages or disadvantages of natural resources booms but with their temporary dimension; i.e., the fact that they constitute substantial temporary revenues, but leave permanent negative effects on the rest of the economy.

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1 The debate does not revolve around whether Dutch disease exists, but whether it should be considered a disease.

2 Literature on the natural resources curse also generally refers to prices and quantities.

In line with the above, this paper is organized as follows: The first part defines and identifies temporary natural resource and capital booms at a global level and makes a comparison between the different types of booms. The second estimates the impact of different types of temporary booms on manufacturing output, and the last section sets out some conclusions and questions for further research.

I. TEMPORARY RESOURCE BOOMS AT A GLOBAL LEVEL: IDENTIFICATION AND CHARACTERIZATION

A. Natural Resource Exports and Private Capital Flows: Trends and Cycles

During the last 50 years, global exports of natural resources have amounted to between 3.5% and 7% of world GDP.[3] As Figure 1, panel A, shows, in said period there have been two major peaks: the first between 1974 and 1985, and the second, slightly larger than the former, from 2003 onwards. This paper attempts to focus more on episodes of this nature than on the behavior of the series as a whole.

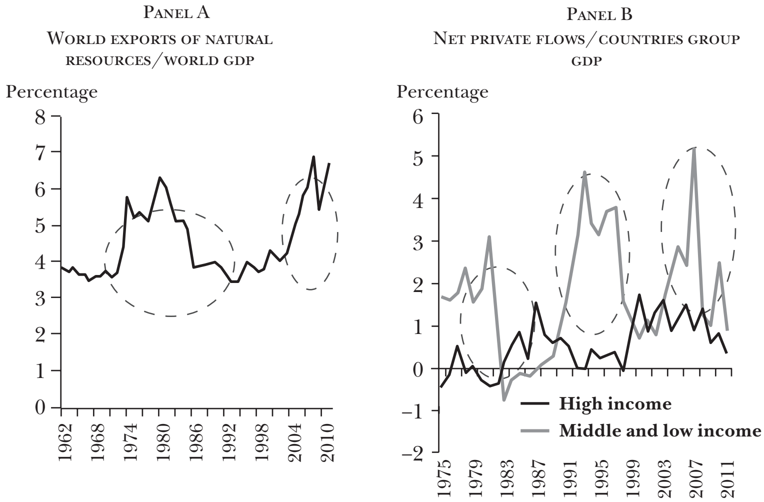
Private capital flows have also performed an increasingly important role in the global economy. According to the database of Bluedorn et al. (2013), between 1975 and 2011, gross capital flows as a percentage of GDP shifted from 5% to 25% in developed countries, and from 2.5% to 12% in developing ones. Nevertheless, as can be seen in Figure 1, panel B, the participation of net capital flows, the ones that can really have a revaluation effect on manufactured products, is relatively more stable for high-income economies than for middle and low-income countries. Three peak periods can also be identified for such flows, which, just like those of natural resources, are the main subject of this paper.

In the case of natural resources, as well as that of capital flows, these episodes tend to have a greater impact on middle and

[3] WDI World Bank.

Figure 1

NATURAL RESOURCES EXPORTS AND PRIVATE CAPITAL FLOWS



Sources: World Bank and own calculations.
Note: Global net flows are not presented because the aggregate of the total flows are netted.

low-income countries. Table 1 shows that, although middle and low-income countries do not receive the majority of the global revenues from commodity exports and net capital flows, they have been the most vulnerable to the fluctuations in those markets: The share of such revenues (exports and capital flows) in GDP is much higher and they are more volatile. In the case of South America, the share of GDP and volatility duplicate the values observed in high-income countries throughout the period studied. With respect to the evolution of this vulnerability, it is possible to conclude that the share of natural resource exports in GDP and their volatility have increased, while the volatility of net capital flows has tended to decline across all country aggregates. Nevertheless, it should be mentioned that the decline in volatility in South America is very low when taking into account

Table 1

NATURAL RESOURCE EXPORTS AND PRIVATE CAPITAL FLOWS

	Natural resource exports					*Net private capital flows*				
	Percentage of global exports	*Share of GDP (%)*				*Percentage of global flows*	*Share of GDP (%)*			
		Average		*Deviation*			*Average*		*Deviation*	
Region	1962-2011	1962-2011	2002-2011	1962-2011	2002-2011	1980-2011[a]	1980-2011	2002-2011	1980-2011	2002-2011
High income	64	4	6	1.1	1.6	86	0.6	1.0	0.6	0.4
Middle and low income	36	8	11	1.8	2.7	14	1.7	2.1	1.5	1.3
South America	7	9	12	2.1	3.3	3	1.5	0.8	2.5	2.1

Sources: Banco de México and own calculations.
[a] Share of total gross flows during the period. The share of net flows is not used because in principle it tends to be zero.

the fact that the size of flows as a percentage of GDP has decreased significantly.[4]

The following section presents a methodology for identifying resource booms at a global level, emphasizing the South American case, and the subsequent sections analyze the results at a regional and sectoral level.

B. Methodology for Identifying Booms

To identify natural resources booms the World Bank database of World Development Indicators (WDI, 1964-2012) for 144 countries was used.[5] Export series over long-term GDP[6] were

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[4] As Bluedorn et al. (2013) show, greater volatility could be explained by the size of the flows (or exports). In fact, when calculating the coefficient of variation(deviation/average) for natural resources (1962-2011) the results are similar among high-income countries (0.3), middle and low-income countries (0.2) and South America (0.2). Moreover, no changes are observed in the volatility coefficient in the last period (2002-2011), except for a small increase from 0.2 to 0.3 in South America. In the case of capital flows (1980-2011), the coefficient of variationis lower for middle and low-income countries (0.9) than for high-income economies (1), and it declines for both country aggregates during the last period (2002-2011) to 0.4 and 0.6, respectively. However, in the case of South America, the coefficient of variation is higher and has tended to increase (1.7, throughout the sample vs. 2.6 in the last period).

[5] The sample excludes countries such as Hong Kong, Panama, Singapore, Luxembourg, Kiribati, the Gaza Strip, Oman, Equatorial Guinea, Democratic Republic of the Congo and Bahamas, that are centers for re-exporting natural resources and whose inclusion would therefore distort the results or present statistics that do not provide logical results. Countries for which there was not sufficient information were also excluded according to the criteria that they should have at least 75% of the 25 data items (13 at the ends, increasing progressively up to 25) to be used for obtaining moving averages of the 25 order series. This means that it is necessary to have 75% of 13 data items for the ends and 75% of 25 data items for the middle of the series.

[6] Calculated for each year as trend GDP based on the Hodrick Prescott filter (1997), with parameter $\Lambda = 400$.

employed for agricultural products (foodstuffs and other commodities) and fuels and minerals, applying the criteria[7] summarized in Diagram 1, which must be met for three consecutive years[8] in order to define a boom:

1) The value of natural resource exports of a given group must be greater than four percentage points of long-term GDP (see Sachs and Warner, 1999). This criteria ensures that the booms selected are important for the economy of the country in question.

2) The value of exports over GDP of a given group must be at least one standard median deviation above the series

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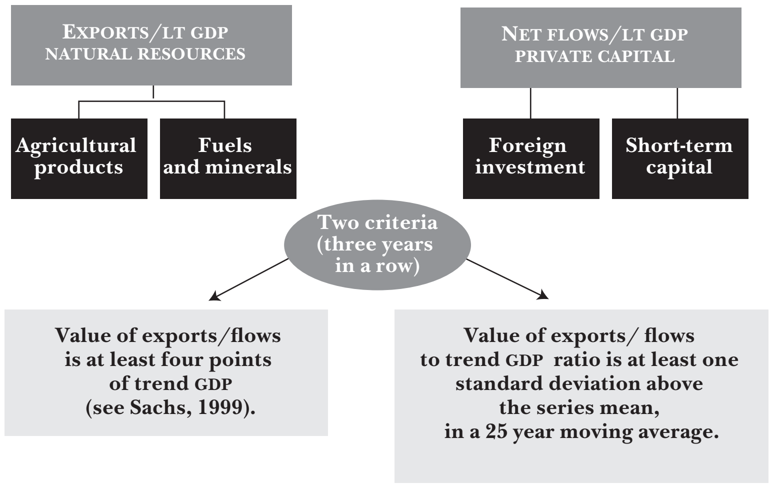
To avoid the problem presented by the filter with the first and ending observations, data from between 1960 and 1963 was eliminated from the filtered series. On the opposite end, the series was completed with IMF projections before proceeding to filter the series and the last four obervations were also eliminated from the filtered series. Parameter $\Lambda = 400$ was employed. This value is suggested for annual data by Correia et al. (1992) and Cogley and Ohanian (1991). Other authors suggest different values depending on the objectives sought (Backus and Kehoe, 1992, suggest a value of 100, and Ravn and Uhlig, 2002, a parameter of 6). However, for this exercise a parameter of 400 was chosen because it is desirable for the trend to be as linear as possible and ensure sustained falls (increases) in GDP are not interpreted as booms (end of booms).

[7] Additionally, exercises were also performed in which a third criterion was included: in the boom years the value of exports (or flows) was higher than the moving average of the series of order 25. We found that only 6% of the data did not meet this criterion, and several of these cases could accommodate the exceptions provided for bonanzas over four years (see note 8). It was decided to privilege the simplicity of the methodology and apply only the two criteria mentioned.

[8] In order to allow temporary and modest deviations, it is not necessary for intermediate year to met one out of the two established criteria or data available, as long as the data is above the median and the bonanza lasted at least four years. Large two-year booms (higher than the mean of all the sector's booms) are also included.

Diagram 1

METHODOLOGY FOR IDENTIFYING TEMPORARY BOOMS



Note: Non-fulfillment of one criteria is allowed in the year as long as the boom lasts for at least four years. Countries with at least 75% of potential data are included in order to obtain a 25-year moving average. 10 countries from the World Bank sample are excluded.

median,[9] on a 25-year moving average. This criteria excludes countries that are structurally producers of natural resources but have not undergone significant changes in the revenues they receive from that item. The use of a moving average prevents structural changes in the series, such as the so-called green revolution (*revolución verde*) in Bolivia, being captured as booms.

This exercise is also applied to the series of net private capital flows consisting of foreign direct investment and other short-term flows.[10] The database employed was that of Bluedorn,

[9] The median is used instead of the average in order to eliminate the bias created by extreme observations and the effect booms have on sample period averages.

[10] Portfolio held in bonds and stocks –less than 10% of the value of the firm–; derivatives and other private investments, including

Duttagupta, Guajardo and Topalova (2013), for the period 1980-2011.[11]

Annex 1 presents a full list of the temporary booms (natural resources and capital) found.

This methodology is comparable with other exercises in the literature for identifying natural resource booms: Sachs and Warner (1999) establish a selection criteria where exports of a given product must be at least 4% of GDP; Céspedes and Velasco (2011) apply a criteria based on an index of external prices[12] and Adler and Magud (2013) one based on the terms-of-trade.[13] A comparison between the results obtained for South America is presented in Annex 2. In general, all three methodologies tend to find booms around the peaks which Erten and Ocampo (2013) call super-cycles of commodity prices. Nevertheless, one advantage of the procedure employed in this paper as compared to other recent works is that it not only identifies booms stemming from price increases, but also from quantity booms. Although quantity booms generate greater added value, this added value is very limited in the case of natural resources. Of more importance is the fact that such booms are also temporary, while their negative effects on other sectors can be long-lasting. Leaving quantity booms out of the analysis could result in important omissions.

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loans, deposits, bank capital and foreign trade credits, aimed at the private sector.

11 Some countries have information since 1970.

12 Velasco and Céspedes define a boom as an episode during which the standardized and deflated price index of a primary product reaches a level of at least 25% above its trend (centered moving average with a 50 year window). The price index was constructed for 33 countries and is weighted using the share in exports or, alternatively, the share in output.

13 Adler and Magud (2013) define a boom as an episode in which commodity prices record an annual average increase of at least 3% and increase at least 15% from start to peak. A total of 270 episodes were identified. The boom ends when 33% of the upswing has reverted.

Table 2

PRICE AND QUANTITY BOOMS IN LATIN AMERICA

			Change in the index (2002-2011)			
Country	*Group*	*Product*	*Value*	*Price*	*Quantity*	*Type of boom*
Argentina	Foods	Soy, corn and wheat	2.9	0.8	1.2	Quantities
Paraguay	Foods	Soy	5.2	0.8	2.6	Quantities
Uruguay	Foods	Meat and cereals	3.1	0.1	2.6	Quantities
Chile	Foods	Copper	4.6	2.6	0.5	Price
Peru	Foods	Copper and precious metals	6.8	2.7	0.1	Price
Bolivia	Fuels	Natural gas and zinc	14.9	1.3	5.8	Quantities
Colombia	Fuels	Oil and coal	7.6	1.6	2.3	Quantities
Ecuador	Fuels	Oil	5.4	1.5	1.6	Both
Venezuela	Fuels	Oil	2.1	1.1	0.5	Price

Sources: World Bank, Comtrade and own calculations.

In the case of South America, recent agricultural product booms in Argentina, Paraguay and Uruguay, and those of fuels in Bolivia and Colombia, have consisted more of quantities than prices (see Table 2). Moreover, methodologies that only include price indicators might lead to identifying booms in times of crisis. One example of this could be Colombia's coffee boom at the end of the seventies. The procedure described here finds a boom between 1977 and 1980, while that employed by Adler and Magud (2011) identifies this boom between 1981 and 1985, right in the middle of the coffee crisis; and that of Céspedes and Velasco (2011) identify it between 1974 and 1985, a complete coffee cycle. Moreover, according to the price criteria, Venezuela could still be said to be in the oil boom in 2013, as found by Adler and Magud (2013), while our estimates find that the boom ended in 2008. In any case, and in order to make the results more robust, alternative exercises were carried out that change some of the methodology's discretional criteria, such as the minimum size that natural resource exports should have as a percentage of GDP.

C. Characteristics of Temporary Booms in a Global Context

The results from applying this methodology at a global level are shown in Table 3. In the case of natural resources, out of the 144 countries included in the sample,[14] 101 experienced booms, i.e., 67% of the countries have registered a natural resource boom at some time since 1964. In Latin America, 11 out of the 12 countries studied have enjoyed at least one boom episode. Meanwhile, the total number of natural resource booms found with the procedure employed is 231, meaning that on average each country has experienced 1.6 booms during the last 50 years. South America is the region that has had the largest share of booms per country (2.9). This is in contrast to China,

[14] At least one piece of data in a sector has sufficient information (see criteria) for calculating the median in a moving window of 25.

Table 3

INCIDENCE OF TEMPORARY BOOMS BY REGION

	Agricultural product or fuel and mineral booms (1965-2012)			*Foreign investment or short-term capital booms (1980-2011)*[1]		
	Number of countries included in the sample	*Incidence of countries with booms (%)*[1]	*Incidence of booms*[1]	*Number of countries included in the sample*	*Incidence of countries with booms*	*Incidence of booms*[1]
South America	12	92	2.9	12	92	1.4
Central America	14	79	1.5	16	88	1.6
Sub-Saharan Africa	34	68	1.4	32	63	0.8
South Asia	6	33	0.7	6	50	0.7
East Asia and the Pacific	11	73	1.9	12	92	1.5
Europe and Central Asia	17	71	1.2	17	76	1.2
Middle East and North Africa	8	63	2.1	9	67	1.1
High-income countries	42	71	1.7	38	82	1.3
Total	**144**	**71**	**1.6**	**142**	**77**	**1.2**

Sources: World Bank and own calculations.
Note: The number of booms for Sub-Saharan Africa and Asia and Central Europe might have been underestimated because in the majority of cases no information is available before 2000. The same occurs with private capital booms.
[1] Countries with boom or booms / Number of countries included in the sample.

Table 4

TIMELINE OF BOOMS

	Agricultural + fuel and minerals sector				*Foreign investment + short-term flows*			
	Number of boom years		*Boom years/ available data*		*Number of boom years*		*Boom years/ available data*	
	1964-2001	*2002-2011*	*1964-2001*	*2002-2011*	*1982-2001*	*2002-2011*	*1982-2001*	*2002-2011*
South America	68	64	10	27	43	8	10	4
Central America	22	24	4	9	35	32	7	11
Sub-Saharan Africa	48	64	8	15	34	43	3	9
South Asia	–	–	–	–	7	11	3	10
East Asia and the Pacific	30	16	8	12	48	17	13	8
Europe and Central Asia	0	66	0	23	0	59	0	19
Middle East and North Africa	62	26	19	19	8	21	4	15
High-income countries	120	134	5	17	67	73	5	10
Total	**350**	**394**	**7**	**17**	**242**	**264**	**6**	**11**

Sources: World Bank and own calculations.

India and South Korea, which have not experienced any natural resource booms during the last 45 years.

In the case of capital flow booms, the region with the highest boom indicator is Central America (1.6), followed by East Asia and the Pacific (1.5). One might initially think that the number of capital booms is lower than that of natural resources. However, it is important to take into account that the study period for capital flows is much smaller.

The results for the duration of booms in each region during the recent period as compared to previous years are presented in Table 4. The most interesting result is that the years of natural resource booms the last decade have been more numerous than in the previous 38 years and, in the case of capital flows, slightly numerous than during the two previous decades. It could be argued that the aforementioned is due to the amount of available data. Nonetheless, if the number of years in boom is divided by the available information, it is found that the probability of a country experiencing a natural resource boom in any given year during the last decade is 17% as compared to 7% in previous decades, and 11% as compared to 6% in previous years. The Middle East was the great protagonist of natural resource booms until 2001, but since then South America has become the region where it is most likely for a country to have a boom in any given year. In the case of capital flows, the region with the highest number of booms according to the information available between 1982 and 2001 was East Asia and the Pacific, while in the recent decade, Europe and Central Asia took the lead in this indicator.

As for magnitude (defined as the ratio of exports to long-term GDP minus the series mean in an average year of the boom), the largest agricultural product booms take place in Central America and the Caribbean, and in sub-Saharan Africa. For instance, the coffee boom of 1976 lasted around five years and generated 13 additional points of GDP for El Salvador, 7.5 for Nicaragua, and 5 for Costa Rica. In Colombia that boom generated four points of GDP for four years. In the mining sector, the recent copper boom generated substantial additional revenues for some Latin American countries and in sub-Saharan Africa. Said mineral produced 15 additional points over four years in Zambia; ten additional points over three years in Chile, and six additional points over eight years in Peru. With regard to fuels, as would be expected, booms have been most intense in oil producing countries. In Brunei, for instance, oil exports reached 169% of long-term GDP in 1980 and the size of the boom, as we measured it, was 100% of GDP. The country in Latin America that has faced the largest oil shocks, taking into account the size of its economy, is Trinidad and Tobago. As for short-term capital flows, the greatest shocks have been experienced by high-income countries such as Iceland (which received additional revenues amounting to 46 points of long-term GDP over five years) and Ireland (which received additional revenues totaling 24 points of long-term GDP over three years). In foreign investment, besides tax havens, the case of Bolivia, which received 7.5 additional points of long-term GDP for eight years, stands out.

However, even more interesting than examples of countries that have experienced booms, are those of countries that have never had them. Countries traditionally used as examples of development such as Japan, India, China and Korea, have not experienced a natural resource boom in the last 45 years. On the other extreme are countries such as Malaysia, which in the last 50 years has had eight natural resource booms, and Belgium and Bolivia that faced five booms during the same period. Meanwhile, countries like Germany have never received a natural resource boom, while Jordan and Malaysia have had four, and Chile and Argentina, three.

D. Natural Resource Booms in South America

As mentioned previously, the methodology employed in this paper provides very intuitive results for South America (Table 5). It also correctly identifies the mineral booms of Chile and Peru, the oil booms of Ecuador, Colombia and Venezuela, and the sixties and seventies coffee booms of Colombia, as well as the cereal booms of Argentina, Uruguay and Paraguay. As for capital flows, the only recent booms identified are those of foreign investment flows to Uruguay and Costa Rica.

If both natural resources and capital are taken into account, the country that has had most booms is Chile. The latter suggests a priori that well-managed booms can generate good macroeconomic results. At the other extreme of the results is Brazil, which stands out for the small number of booms identified. This is explained by its high level of diversification and limited economic openness, meaning natural resource shocks in Brazil are not as important for its economy as in other countries of the region.

A comparison of the size of booms shows that Bolivia experienced the largest ones out of the whole group of countries. In particular, with the recent fuel and minerals boom, it has been receiving 11 additional points of GDP since 2005. Although in Venezuela oil exports account for around a quarter of GDP, such share is relatively stable (the median is 22%) and therefore in terms of size the boom only occupies fourth place in South America.

E. Comparison of Booms by Sector

The results from applying the methodology can be analyzed by sector of specialization: agricultural products, fuels and minerals, short-term capital flows, and investment flows. Among natural resources, instinct indicates that this differentiation could be crucial when analyzing the effects of booms on industry. According to the World Bank (2010), the different effects of booms can be explained by the fact that the characteristics distinguishing commodities from other kinds of goods are more pronounced in the case of minerals and fuels than for agricultural products. Some of these specific characteristics mentioned in the report are: *i)* their highly volatile prices; *ii)* high initial investment requirements, discouraging private investment and meaning a large amount of the companies are state owned[15] and, in the case of mining, in foreign hands; *iii)* the fact they are not renewable, and *iv)* their production often takes place in specific geographical enclaves. Among capital flows, foreign direct investment tends to be more stable and more actively involves purchasing national assets, which can create different effects when analyzing the impact on the value-added in manufacturing.

Some of these differences become evident when carrying out a simple characterization of booms. As can be seen in Table 6, in general, the fuel and minerals sector has been characterized by longer and larger booms, while the agricultural products sector has exhibited smaller-sized booms (in terms of the exports indicator minus the median of the series of exports/GDP) and their duration has been shorter. The latter can be partly explained by the so-called cobweb theory[16] (Kaldor, 1934). Furthermore, mineral booms in South America have also been long and large.

Figure 2 shows the number of booms for each type of good over the last 50 years. According to the Figure, there is currently a kind of boom of booms in which the fuel, mineral sector and short-term capital have played an important role. Upon analyzing these results in terms of the size of booms to world GDP (Figure 3), the cycles observed become more pronounced and it becomes evident that fuel and minerals sector and short-term capital flow booms are the largest. In addition, capital flows are frequently received by larger economies, and a higher number of countries, and therefore become more important when they are seen in terms of size as compared to how they appear in terms of the number of booms.

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[15] Céspedes and Velasco (2012) provide the theoretical framework for analyzing how natural resources shocks affect the economy and mention that the results are sensitive to whoever is the owner of the resources: the workers (in the case of some agricultural products) or the government (mainly in the case of fuels).

[16] In a world of perfect competition and elastic supply (such as that of agricultural products), the quantities self-regulate in line with price signals from the preceding period, and the path followed by price and quantity take the shape of a cobweb.

Table 5

CHARACTERISTICS OF BOOMS IN LATIN AMERICA

	Agricultural products			*Fuels and minerals*			*Short-term capital flows*			*Investment flows*		
	Start	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*
Argentina	1977	7	2.7				1993	2	10.7	1999	2	5.6
	2007	6	2.5				1997	2	6.1			
Bolivia	1994	5	1.6	1974	10	4.4				1998	8	7.5
				2005	8	11.1						
Brazil	1964	2	2.1				1994	3	7.0			
Chile	1994	5	1.8	1979	3	4.9	1992	3	5.9	1996	4	2.9
				1988	2	3.0	1996	2	4.5			
				2006	3	10.2						
Colombia	1964	2	2.5	2008	5	4.5	1981	2	3.8			
	1977	4	4.1				1994	3	5.2			

Ecuador	1964	2	5.1	1980	6	4.4	1990	3	7.7			
	1994	5	4.7	2008	4	6.9						
	2011	2	1.8	2011	2	4.9						
Paraguay	1989	2	7.1	2010	3	9.6	1981	2	5.1			
	2001	3	4.1									
	2007	2	9.2									
Peru	1964	3	2.6	1979	7	5.3				1994	4	5.1
	1994	4	1.3	2005	8	5.8						
	2008	5	1.0									
Uruguay	1980	4	3.4							2006	5	3.2
	1996	3	1.5									
	2008	5	3.1									
Venezuela				1979	4	11.9				1997	2	5.3
				2005	4	8.9						

Table 6

SECTORAL DIFFERENCES BETWEEN BOOMS BY TYPE OF RESOURCE

	Total			*South America*		
	Numer of booms	*Duration of boom (years)*	*Size of boom (percentage of GDP)*	*Numer of booms*	*Duration of boom (years)*	*Size of boom (percentage of GDP)*
Foods and materials	133	3.5	4.1	20	3.8	3.2
Minerals and fuels	101	4.0	8.5	15	4.7	7.2
Short-term capital flows	80	2.7	8.8	9	2.4	6.3
Investment flows	88	3.4	6.9	8	3.9	6.4
Weighted average	402	3.5	6.7	52	3.7	5.3

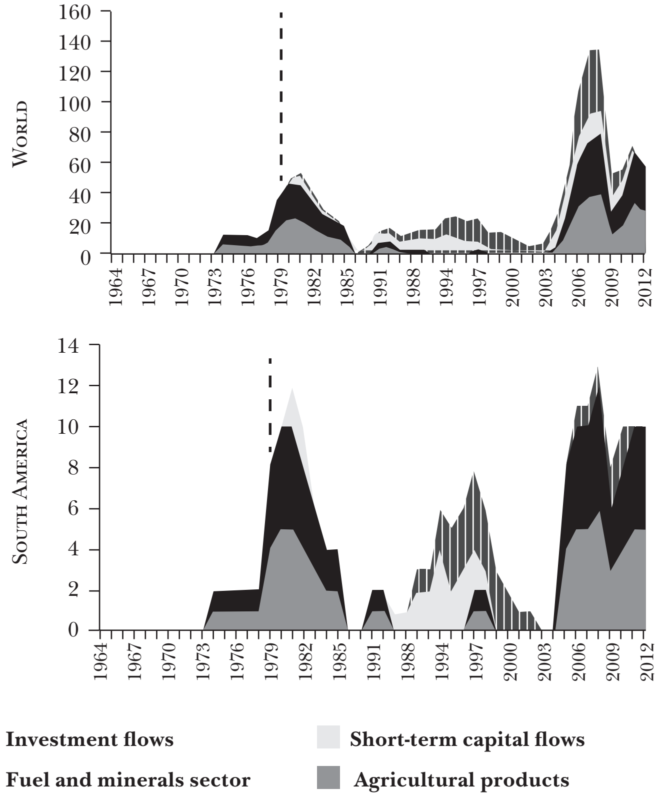
Source: World Bank and own calculations.

Furthermore, the group of figures above shows how South America is also currently undergoing a real natural resources boom of booms with minerals and fuels playing a prominent role.[17] Once again, the results in terms of size intensify the cycles and illustrate the size of capital flows that the region experienced during the mid-nineties.

[17] These results are not significantly affected when they are divided by the number of countries included in the sample due to the fact data series for South America are sufficiently long and the number of countries included in the sample does not change significantly over time.

Figure 2

NUMBER OF BOOMS, 1964-2012 AND PRIVATE CAPITAL FLOWS



Sources: World Bank and own calculations.
Dotted line: start of capital flow data.

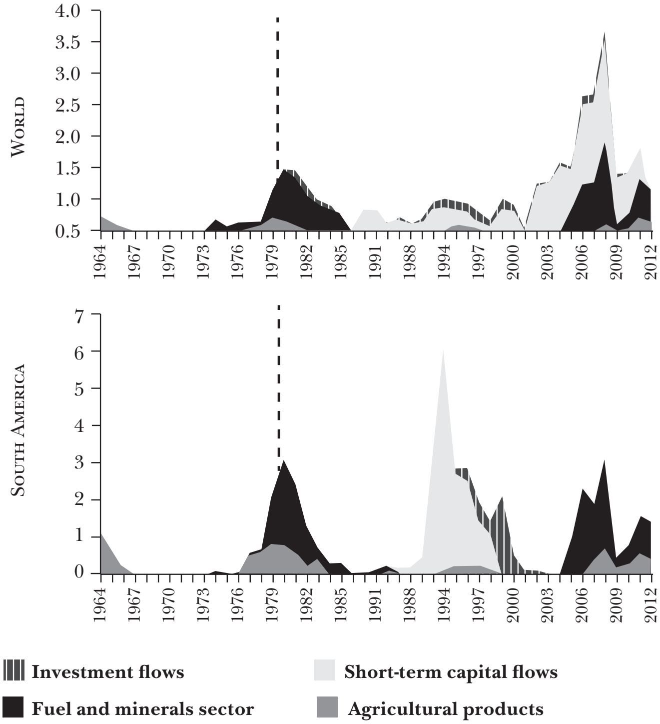
It can be concluded that:

- Natural resource booms are very important for South America, especially in recent times.
- Although capital booms have been relatively less frequent in the region, they were very important in the mid-nineties. These booms have generally played a procyclical role with respect to natural resource booms.

Figure 3

BOOMS MAGNITUDE

(as countries group GDP percentage)



Sources: World Bank and own calculations.
Dotted line: start of capital flow data.

- There are reasons for thinking that the type of product an economy specializes in explains the differences in the characteristics of booms and their expected impact on the economy.
- In general, fuel and minerals booms (as opposed to those of agricultural products) have tended to be long and large. Capital booms are also large, but short.

IMPACT OF BOOMS ON MANUFACTURING'S SHARE OF GDP: DETAILS OF THE ESTIMATION

The econometric estimations aim to examine the effects of booms on the performance of manufacturing using information from all the countries and taking advantage of the structure of panel data. After carrying out the statistical tests, the estimator of Driscoll and Kray (1998) of fixed effects with standard errors that are robust to the heteroskedasticity, contemporaneous and serial correlation of this type of data, is used (Hoechle, 2007). According with that suggested by the latter, it is desirable to have relatively long panels in order for the estimator to be more robust, given its asymptotic properties. The database was therefore restricted to countries for which there would be at least 30 pieces of available data for making the corresponding regressions. In general terms, the equation is estimated is as follows:

$$\begin{aligned} y_{i,t} &= cte + tamalimat_{i,t} + tammincom_{i,t} + tamfdk_{i,t} \\ &+ tamfdi_{i,t} + postalimat_{i,t} + postmincom_{i,t} \\ &+ postfdkcp_{i,t} + postfdi_{i,t} + controls_{i,t} + e_{i,t}, \end{aligned}$$

where $y_{i,t}$ is the value added in manufacturing as a percentage of GDP. *cte* is the constant; $tamalimmat_{i,t}$, $tammincom_{i,t}$, $tamfdkcp_{i,t}$, and $tamfdi_{i,t}$ are variables that take a value of 0 if country *i* is not in boom during year *t* or the value of the boom in that year (measured as the value of the series minus the mean / long-term GDP, in the case of agricultural products, fuel and minerals, short-term capital flows and investment flows) if country *i* experiences a boom. Variables with the prefix *post* correspond to the post-boom periods that take a value of 0 if country *i* is not in a post-boom period during year *t* or the average value of the boom. Post-boom periods are calculated as the three subsequent years after the boom ends for all sectors except for short-term capital flows, where the results two years after the boom were found to be most significant. The variables $controls_{i,t}$ include GDP per capita in constant terms, the same variable squared (to capture the effect on manufacturing of the level of development, which is assumed to be decreasing) and the value of exports and capital flows to verify whether it is booms or regular flows of resources that are having an impact on the value added in manufacturing. $e_{i,t}$ is the random error component.

Two groups of regressions are presented. The first group is made for 1980-2011 and includes variables for all capital booms. The second is for the period 1965-2012 and only uses variables for natural resources (those for capital flows are not available for the whole period). The Federal Reserve funds rate is added to the regressions to control for capital flows, while this variable is in turn controlled by US economic growth to prevent the equation capturing the effect of GDP growth in that country as a result of its counter-cyclical monetary policy.

II. EFFECTS OF BOOMS ON THE ADDED-VALUE IN MANUFACTURING

To analyze the effect of booms on manufacturing output, an equation was estimated that uses the ratio of value added in manufacturing to long-term GDP[18] as a dependent variable and the size of booms and corresponding post-boom periods multiplied by the size of the respective booms, and an indicator for the countries' level of development as independent variables (see Box 1).

Table 7 shows the estimates for a group of 20 countries in the period between 1980 and 2011. One of the most interesting results obtained is the different effects of the booms: the contemporary impact of fuel and mineral booms is negative, while the effect of agricultural product booms tend to be positive and those of capital flows is not significant. The aforementioned might be explained by the characteristics mentioned in the previous section. *Dutch disease* effects tend to be greater for the fuel and minerals sector due to the inelasticity of supply, the greater discretion governments usually exercise with regards to revenues associated with the booms, and the few links the sector has with manufacturing industry. In the case of capital flows, the potentially negative effects of a revaluation are offset by the positive impact of financing on the industry.

However, the most outstanding effect obtained by the exercise is that related to post-boom periods. During the three years following the boom (two years in the case of capital booms) there is still a significant negative impact on manufacturing, highlighting how difficult it is for industry to recovery from the shocks it suffers during boom periods, especially those that will probably be generated by the appreciation of the local currency.

[18] This ratio is calculated in constant local currency, preventing exchange rate movements from affecting the value of the variable. Nigeria and the Democratic Republic of the Congo, which presented non-intuitive values in WDI data series, were excluded from the analysis.

Table 7

ESTIMATE OF THE VALUE-ADDED IN MANUFACTURING / LONG-TERM GDP (1980-2011)

		Total of the sample				*Middle and low-income countries*[1]			
		(1)		*(2)*		*(3)*		*(4)*	
Boom magnitude	Foods	0.23[a]	(1.83)			0.30	(1.66)	0.22[b]	(2.44)
	Fuels and minerals	–0.39[c]	(–3.57)	–0.46[c]	(–4.19)	–0.41[c]	(–3.80)	–0.50[c]	(–4.75)
	Short-term capital flows	0.01	(0.17)			–0.02	(–0.18)		
	Investment flows	0.03	(0.43)			0.44[b]	2.49	0.42[c]	(3.00)
Post-booms	Foods	–0.17	(–1.61)	–0.25[b]	(–2.22)	–0.01	(–0.07)		
	Fuels and minerals	–0.39[c]	(–3.80)	–0.42[c]	(–3.86)	–0.34[c]	(–3.72)	–0.36[c]	(–4.82)
	Short-term capital flows	–0.09	(–1.74)			–0.22[b]	(–2.43)	–0.22[b]	(–2.83)
	Investment flows	–0.14[b]	(–2.17)	–0.12[b]	(–2.40)	–0.07	(–0.49)		

Controls	GDP per capita	-0.70^{c}	(−4.01)	-0.61^{c}	(−5.39)	4.44^{c}	(4.22)	4.78^{c}	(4.58)
	GDP per capita^{2}	0.01^{c}	(5.23)	0.01^{c}	(6.03)	$--0.30^{c}$	(−3.16)	-0.33^{c}	(−3.52)
	NR exports/LT GDP	−0.06	(−1.00)			−0.07	(−0.83)		
	Capital flows/ LT GDP	−0.04	(−1.53)			-0.07^{a}	(1.87)	-0.08^{b}	(−2.76)
	Trend	0.01	(0.50)			-0.08^{b}	(−2.64)	-0.09^{b}	(−2.76)
	Constant	−6.43	(−0.12)	20.72^{c}	(19.97)	167.97^{c}	2.87	183.36^{c}	(2.96)
Indicators	Observations	606		606		242		242	
	Groups	20.00		20.00		8.00		8.00	
	R within	0.23		0.22		0.38		0.38	
	F	111.94		23.67		102.84		29.24	

Sources: World Bank and own calculations.
Note: Driscoll and Kray, fixed effects (1980-2011).
[1]Excludes Middle Eastern and North African countries.
[a] $p < 0.1$, [b] $p < 0.05$, [c] $p < 0.01$.

In fact, if the economies were totally flexible, a boom would imply a simple reallocation of productive sectors associated to the appreciation of the currency, which would revert once the boom ended. However, the results found here indicate that once the boom ends the revenues derived from it revert rapidly (and the currency probably depreciates again), but the process of recovery in manufacturing industry is much slower.

The real exchange rate is one of the variables that might explain the limited capacity of industry to recover rapidly. An exercise which analyzes the average performance of the real exchange rate two years before a boom, during a boom, and two years after booms, finds that currencies appreciate during booms, but during the two years after they do not adjust rapidly to their new equilibrium level, and can even continue to appreciate (Table 8). More important is the fact that exchange rate effects, and those related to prices in general, tend to have a considerable lag and cause substantial inertia in the production of different types of goods.

The above does not mean to say that there are no other factors limiting the ability of industry to recover. Among such factors it is worth mentioning: the loss of position on the learning curve (Krugman, 1987), the difficulty of reallocating factors across sectors and the problems that emerge while attempting to recover markets for manufacturing products. In the case of capital flows, the impact can also be understood as the end of the financing effect.

As can be seen in Table 7, among the post-boom impacts, that of the fuel and mineral sector is the largest, followed by investment flows. The effect is not significant for foods. It is essential to keep in mind that these coefficients refer to each point of the annual average size of the boom, i.e., a boom that generates five additional points of annual GDP would on average cause around two points less in the value of manufactures as a percentage of long-term GDP during the boom and in the three years following it.

Another aspect worth pointing out involves the impact that exports of natural resources have on GDP, besides that which

Table 8

BEHAVIOR OF THE REAL EXCHANGE RATE DURING BOOM CYCLES

Averages	*Change in the growth rate of the real exchange rate during the boom*	*Change in the growth rate of the real exchange rate during the post-boom*	*Change in the growth rate of the real exchange rate between pre and post-boom*
Agricultural products	6.0[b]	−1.5	3.1
Fuels and minerals	5.3[a]	1.8	7.8
Aggregate natural resources	6.1[b]	2.0	8.5
Short-term flows	7.3[c]	3.1	11.7[b]
Investment flows	4.5[b]	1.0	6.6
Aggregate capital flows	6.5[c]	2.3	9.2[a]

Sources: World Bank, Bluedorn et al. (2013) and own calculations.
Levels of significance obtained with *t*-statistic: [a] $p < 0.1$, [b] $p < 0.05$, [c] $p < 0.01$.

takes place through booms. The regressions include this control variable but it was not statistically significant, indicating that booms, rather than the stable flow of resources, tend to be associated with an impact on the value added in manufacturing. Moreover, the fact that this variable is not significant ensures that the effect captured from the booms is not the result of a simple reallocation of shares in GDP. In the case of capital flows, the variable expressed as a percentage of long-term GDP is significant, but its coefficient is modest, and much smaller in size than the other coefficients in the equation.

The above exercise was repeated, excluding high-income countries, and Middle Eastern and North African countries, most of which are oil producers. The results are shown in estimates 3 and 4 of Table 7 and are very similar to those obtained with the whole sample. Nevertheless, the coefficients for the post-boom periods tend to be higher for capital flows.

To support the above exercise, and include the cumulative booms from the seventies, an exercise was carried out that made the same estimation since 1965. The results of the latter are presented in Table 9. The effects of capital flows are not included there because the corresponding data only starts to be published consistently after 1980. To address the absence of these variables, the series are controlled by the Federal Reserve funds rate and US real economic growth, ensuring that the Federal Reserve rate captures the effect of capital flows and not the impact of us anticyclical policy.

As seen by comparing Table 9 with Table 7, exercises on a longer period of analysis (1965-2012 vs. 1980-2012) result in significant changes in the results: the incorporation of the value of exports/GDP as a control variable leads to statistically significant results and the contemporary impact of natural resource booms is no longer significant. However, the persistence of the negative impact in the post-boom period is seen once again, although less pronounced, in the cases of fuel and mineral exports. The aforementioned might suggest that the negative effect of these booms on manufacturing has tended to increase during the last 30 years. Once again, the exercise is repeated excluding high-income, North African and Middle Eastern countries from the sample. Said exercise shows how the negative effect of post mineral and fuel booms on manufacturing industry is stronger for developing countries.

III. CONCLUSIONS AND NEXT STEPS

The main conclusions that can be made from the above analysis are:

- The world is undergoing a boom of booms at a global level, in which South America is playing a prominent role.

- Booms, more than stable income derived from natural

Table 9

ESTIMATE OF THE VALUE-ADDED IN MANUFACTURING / LONG-TERM GDP (1965-2012)

		Total of the sample				*Middle and low-income countries*[1]			
		(1)		*(2)*		*(3)*		*(4)*	
Boom magnitude	Foods	0.36^{c}	(4.15)	0.37^{c}	(4.75)	0.26^{c}	(3.40)	0.31^{c}	(5.18)
	Fuels and minerals	0.04	(0.70)			−0.12	(−1.49)		
Post-booms	Foods	−0.093	(−1.10)			0.04	(0.73)		
	Fuels and minerals	-0.06^{a}	(−1.82)	-0.06^{b}	(−2.20)	-0.16^{c}	(−6.10)	-0.13^{c}	(−3.46)
Controls	GDP per capita	0.82^{c}	(8.65)	0.81^{c}	(8.35)	8.65^{c}	(7.88)	8.52^{c}	(7.41)
	GDP per capita[2]	-0.01^{c}	(−6.94)	-0.01^{c}	(−6.86)	-0.71^{c}	(−6.58)	-0.70^{c}	(−6.22)
	Federal Reserve	0.12^{c}	(3.77)	0.13^{c}	(5.29)	0.04^{a}	(1.75)		
	US GDP growth	0.04	(1.64)			0.03	(0.93)		
	NR exports/LT GDP	-0.16^{c}	(−5.05)	-0.15^{c}	(−6.82)	-0.09^{c}	(−2.73)	-0.12^{c}	(−4.23)
	Trend	−0.01	(−0.39)			-0.10^{c}	(−6.10)	-0.11^{c}	(−7.08)
	Constant	23.78	(0.92)	13.79^{c}	(17.25)	203.13^{c}	(6.50)	219.27^{c}	(7.59)
Indicators	Observations	1,625		1,625		1,001		1,001	
	Groups	40.00		40.00		24.00		24.00	
	R within	0.14		0.14		0.31		0.30	
	F	120.52		95.63		44.68		27.91	

Sources: World Bank and own calculations.
[1] Excludes Middle Eastern and North African countries.
[a] $p < 0.1$, [b] $p < 0.05$, [c] $p < 0.01$.

resource exports or capital flows, tend to generate negative impacts on the share of manufacturing industry in long-term GDP. Such effects persist after the booms have ended.

- Fuel and mineral booms are likely to be longer and larger, generate more Dutch disease symptoms and have more persistent effects on manufacturing industry.

- Capital flow booms tend to be large but short. The contemporary effects of these booms on manufacturing are likely to be neutral, which possibly explains why the revaluation effect is offset by greater financing in favor of industry. Nevertheless, the end of these booms also brings a period where manufacturing industry's share in long-term GDP is low.

- Agricultural products booms are likely to have a positive contemporary effect on industry, which might be explained by the elasticity of supply, the lower discretion governments usually exercise with regards revenues associated to the booms and the greater links the sector has with manufacturing industry. The foods post-boom is not significant.

APPENDIX 1

Appendix 1

BOOMS BY REGION

	Agricultural products			Fuels and Minerals			Short-term flows			Investment flows		
	Start	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*
South Asia												
Cambodia										2007	2.0	3.0
China										1993	6.0	3.9
Fiji										2004	5	7.5
Indonesia	1977	4.0	2.4	1980	6.0	9.9						
	1994	4.0	1.5									
	2008	5.0	1.4									
Laos										1994	4.0	7.0
Malaysia	1964	2.0	7.2	1964	2.0	5.9	1984	2.0	6.0	1991	7.0	4.6
	1979	3.0	8.1	1979	7.0	7.9	1992	2.0	8.9			
	2011	2.0	4.3	2006	3.0	5.3	1995	2.0	7.7			
				2011	2.0	5.4						
Papua New Guinea	1992	2.0	8.0				2006	2.0	8.1			
							2009	2.0	11.1			

Phillipines	1974	8.0	1.6				1994	4.0	9.7			
Solomon Islands							1988	3.0	3.2	2008	3.0	17.0
Thailand	1964	2.0	5.4				1991	6.0	9.9	1998	2.0	2.7
	1979	4.0	2.9									
	1995	2.0	2.6									
	2008	5.0	2.7									
Tonga	1991	5.0	2.9									
Vanuatu							2003	3.0	9.6	1991	7.0	7.3
	2011	2.0	3.3	2005	4.0	4.3				1995	3.0	6.8
Europe and Central Asia												
Albania				2011	2.0	3.2				2007	4.0	5.0
Armenia				2010	3.0	3.5				2006	4.0	4.9
Azerbaijan				2007	5.0	28.4				2003	2.0	11.3
Belarus	2010	3.0	1.8	2006	3.0	9.5				2007	2.0	3.0
				2011	2.0	10.1						
Bulgaria	2008	5.0	3.3	2006	3.0	9.4	2006	3.0	13.4	2004	5.0	16.6
				2011	2.0	7.7						
Georgia							2007	2.0	7.9	2006	3.0	9.5

	Agricultural products			*Fuels and Minerals*			*Short-term flows*			*Investment flows*		
	Start	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*
Kazakhstan				2006	3.0	16.0						
Kyrgyz Republic	2010	2.0	2.6	2011	2.0	2.9						
Lavia	2007	2.0	2.6	2011	2.0	2.9	2005	3.0	15.8	2006	2.0	4.6
Lithuania	2007	3.0	6.7	2004	5.0	5.0	2006	2.0	9.8	2006	3.0	2.3
				2011	2.0	6.8						
Macedonia FYR							2008	2.0	3.5	2006	3.0	4.6
Moldova							2007	2.0	11.0	2007	2.0	8.8
Romania				2006	3.0	1.8	2005	4.0	6.8	2004	5.0	5.8
Russia Federation				2006	3.0	8.3						
Turkey							2010	2.0	3.9			
Ukranie	2008	5.0	3.7							2005	4.0	5.8
Central America												
Antigua and Barbuda							1996	2.0	8.3	1987	5.0	6.5
										2003	5.0	15.2
Belize	1984	2.0	26.0	2007	5.0	3.7	2002	3.0	12.1	2004	5.0	4.8

Costa Rica	1976	6.0	5.2							6.0	3.0	2.1
	1993	6.0	6.9									
Dominica	1991	4.0	3.8				1999	2.0	7.3	1994	2.0	10.4
										2008	2.0	4.7
Dominican Republic										1999	3.0	2.1
El Salvador	1976	5.0	12.9							2007	2.0	4.2
	1995	5.0	4.6									
	2011	2.0	1.6									
Grenada										2007	2.0	9.8
Guatemala	1965	2.0	2.9				2000	2.0	3.3			
	1977	5.0	5.4									
Honduras	1965	2.0	4.9				1980	2.0	5.8	2004	5.0	2.3
	1978	3.0	5.1									
	1995	7.0	5.7									
Jamaica				1978	5.0	2.2						
				2006	3.0	3.3						
Mexico				1980	6.0	7.6	1991	3.0	7.7			
				2005	4.0	2.0						
Nicaragua	1976	6.0	7.4							1997	4.0	3.1
	2010	3.0	8.8									

	Agricultural products			Fuels and Minerals			Short-term flows			Investment flows		
	Start	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*
St. Kitts and Nevis										1989	2.0	21.6
										2000	2.0	8.2
St. Lucia	1990	3.0	6.2				2007	2.0	8.6	1990	2.0	5.7
										2006	3.0	13.2
St. Vincent and the Granadines	1999	3.0	2.9				1991	2.0	6.1	1993	2.0	8.9
										1997	2.0	15.9
South America												
Argentina	1977	7.0	2.7				1993	2.0	10.7	1999	2.0	5.6
	2007	6.0	2.5				1997	2.0	6.1			
Bolivia	1994	5.0	1.6	1974	10.0	4.4				1995	8.0	7.5
				2005	8.0	11.1						
Brazil	1964	2.0	2.1				1994	3.0	7.0			
Chile	1994	5.0	1.8	1979	3.0	4.9	1992	3.0	5.9	1996	4.0	2.9
				1988	2.0	3.0	1996	2.0	4.5			

				2006	3.0	10.2						
Colombia	1964	2.0	2.5	2008	5.0	4.5	1981	2.0	3.8			
	1977	4.0	4.0				1994	3.0	5.2			
Ecuador	1964	2.0	5.1	1980	6.0	4.4	1990	3.0	7.7			
	1994	5.0	4.7	2005	4.0	6.9						
	2011	2.0	1.8	2011	2.0	4.9						
Guyana										1992	4.0	14.8
										2009	2.0	4.0
Paraguay	1989	2.0	7.1	2010	3.0	9.6	1981	2.0	5.1			
	2001	3.0	4.1									
	2007	2.0	9.2									
Peru	1964	3.0	2.6	1979	7.0	5.3				1994	4.0	5.1
	1994	4.0	1.3	2005	8.0	5.8						
	2008	5.0	1.0									
Suriname	1995	5.0	6.6	1997	2.0	22.2						
Uruguay	1980	4.0	3.4							2006	5.0	3.2
	1996	3.0	1.5									
	2008	5.0	3.1									
Venezuela				1979	4.0	11.9				1997	2.0	5.3
				2005	4.0	8.9						

	Agricultural products			*Fuels and Minerals*			*Short-term flows*			*Investment flows*		
	Start	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*
East Asia and the Pacific												
Maldives							2005	4.0	5.4	2004	7.0	2.8
Nepal							1991	5.0	4.6			
Pakistan	1964	4.0	2.6									
	1979	3.0	1.6									
Sri Lanka	1964	3.0	8.0				1993	2.0	4.9			
	1977	5.0	3.4									
Sub-Saharan Africa												
Angola										1998	4.0	7.6
Benin							1981	2.0	10.8	1989	4.0	4.6
Botswana										2002	5.0	5.0
Burkina Faso	2010	2.0	10.4									
Cabo Verde										1995	2.0	4.4
										2006	3.0	8.5
Cameroon	1964	2.0	4.5	2006	2.0	5.5						
	1978	2.0	2.8									
	2008	4.0	6.2									

Central Africanan Republic	1998	2.0	1.5									
	2007	2.0	1.8									
Cote d’Ivoire	1964	2.0	13.4	2005	5.0	7.5						
	2010	2.0	4.1									
Ethiopia	2010	3.0	1.8									
Gabon	1964	2.0	20.0	2007	2.0	22.4						
Ghana	1974	5.0	6.3	1976	3.0	1.5				2007	4.0	4.9
				2011	2.0	13.7						
Guinea										2007	2.0	7.1
Kenya	1994	5.0	2.6									
Lesotho										1995	5.0	30.7
Madagascar	1975	6.0	2.7									
	1994	3.0	3.1									
Malawi	1977	5.0	5.0									
	1990	2.0	7.4									
	1996	2.0	4.7									
Mauritania										2007	2.0	3.3
Mauritius	1995	3.0	3.2									

	Agricultural products			Fuels and Minerals			Short-term flows			Investment flows		
	Start	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*
Mozambique	2011	2.0	1.7	2004	4.0	9.0				1998	2.0	6.2
										2001	3.0	2.9
Namibia	2010	3.0	1.2							2007	4.0	2.5
Niger	1965	3.0	1.5	2010	3.0	7.7				2008	2.0	9.6
	2008	2.0	2.5									
Nigeria	1964	4.0	4.1	1974	6.0	9.5		2.0	9.9		6.0	4.1
Senegal				1974	8.0	3.3	2007	2.0	5.0			
				1996	2.0	2.9						
Seychelles	1978	3.0	6.0							2006	5.0	4.9
	2000	5.0	7.9									
Sierra Leone										2004	4.0	3.9
South Africa				2006	7.0	3.6	1976	4.0	6.3			
							2006	2.0	5.8			
Middle East and North Africa												
Algeria				1979	6.0	9.8						
				2005	4.0	17.9						
Djibouti										2006	4.0	15.1

Egypt	1973	5.0	2.3	1980	6.0	5.2				2005	4.0	4.3
				1988	2.0	3.0						
				2005	4.0	3.2						
Jordan	1979	4.0	1.2	1988	5.0	2.0	1991	2.0	30.2	2005	4.0	9.5
	1993	5.0	2.7				1994	2.0	7.6			
							2007	3.0	6.9			
Morocco	1964	3.0	2.8	1974	2.0	7.0						
	1994	4.0	2.7	1979	3.0	3.9						
Syrian Arab Republic							1980	4.0	8.2			
							1993	2.0	5.2			
Tunisia	1972	3.0	1.9				1979	7.0	7.5	2006	3.0	4.0
	2006	3.0	1.0				2007	5.0	3.1			
High-income countries												
Australia	1964	3.0	3.0	1989	4.0	1.2	1986	4.0	1.4			
	1965	3.0	1.0	2008	5.0	5.2						
Bahrain				2007	2.0	17.5	1993	2.0	25.7	1991	2.0	12.9
										1995	2.0	16.4
Barbados	1967	2.0	13.3				1980	2.0	6.3	2005	6.0	9.6
	1974	2.0	6.7				1999	3.0	7.2			
	1995	3.0	1.3				2004	2.0	6.7			

	Agricultural products			*Fuels and Minerals*			*Short-term flows*			*Investment flows*		
	Start	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*
Belgium	1978	4.0	1.4	1979	5.0	2.9						
	1994	3.0	1.0	2004	5.0	4.4						
	2007	2.0	2.0	2011	2.0	5.3						
Brunei Darussalam				1979	6.0	55.1						
Canada	1979	3.0	0.8	1974	3.0	1.3	2009	3.0	4.4			
	1994	4.0	0.7	1979	3.0	1.6						
				2005	4.0	3.9						
Croatia				2006	3.0	1.7				2006	3.0	4.9
Cyprus	1990	2.0	1.7				1989	2.0	3.9	1999	3.0	3.0
	1995	2.0	2.1									
Czech Republic	2011	2.0	1.6	2010	3.0	1.9						
Denmark	1964	3.0	4.0				1987	2.0	4.0			
	1978	4.0	1.7				2009	2.0	11.1			
	1990	3.0	1.0									
Estonia				2006	6.0	5.3						
Finland	1964	3.0	3.4				1987	4.0	7.7			
	1979	3.0	2.1				2008	4.0	14.6			

Francia							2009	2.0	5.6			
Greece				2011	2.0	3.6	1998	2.0	4.2			
Greenland	1987	4.0	12.3	1988	2.0	9.5						
Hungría	1979	5.0	3.3	1980	5.0	1.8	1993	3.0	8.2	1995	4.0	4.3
	2008	5.0	2.1				2004	3.0	5.6			
Islandia	1964	3.0	15.2	2007	6.0	7.6	2004	5.0	46.1			
	1978	4.0	7.0									
	1987	6.0	3.4									
Irlanda	1975	7.0	4.0				2004	3.0	24.4	1999	5.0	9.0
Israel	1976	6.0	1.0				1995	3.0	6.4			
Italia							2007	3.0	4.3			
Korea, Rep.	1965	2.0	3.7									
Kuwait				1979	3.0	47.2				1994	2.0	6.2
				2006	3.0	24.5						
Malta				2010	3.0	16.8				1999	2.0	10.1
										2006	2.0	13.5
Monaco										2002	3.0	2.9
Netherlands	1978	2.0	2.9	1976	10.0	3.9	2010	2.0	10.6			
	1994	3.0	2.2	2006	7.0	4.3						
	2011	2.0	4.7									

	Agricultural products			Fuels and Minerals			Short-term flows			Investment flows		
	Start	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*	*Start*	*Duration*	*Magnitude*
New Zealand	1979	4.0	3.4				2004	4.0	6.2	1995	3.0	3.6
	1995	2.0	2.3									
	2011	2.0	2.9									
Norway	1964	3.0	1.8	1979	4.0	5.5	1986	3.0	5.4			
				2005	4.0	8.2						
				2011	2.0	4.6						
Poland	2008	5.0	1.4				2008	3.0	4.6			
Portugal							1981	2.0	7.5			
							2006	2.0	6.5			
Slovak Republic	2011	2.0	1.9	2006	3.0	2.1						
				2011	2.0	3.1						
Saudi Arab				2005	4.0	16.5				1981	4.0	6.1
				2011	2.0	15.3				2006	4.0	4.9
Slovenia				2006	3.0	2.8						
				2010	3.0	3.2						

Spain				2004	4.0	8.7			
Sweden				1988	3.0	10.2			
Switzerland				2008	4.0	13.7			
Trinidad and Tobago	1974	9.0	26.6				1993	2.0	3.7
	2005	4.0	32.6				1997	2.0	7.1
United Arab Emirates				1993	3.0	9.9			
United Kingdom	1980	6.0	2.7	1987	2.0	5.6			
				1999	2.0	7.2			
United States				2002	7.0	2.9			

APPENDIX 2

Appendix 2

THE RESULTS COMPARED WITH OTHER METHODOLOGIES

	Fernández-Villar (2013)						*Velasco-Céspedes (2012)*				*Adler and Magud*	
	Agricultural products		*Fuels and minerals*		*Natural resources aggregate*		*Output*		*Exports*			
	Start	*End*	*Start*	*End*	*Start*	*End*	*Start*	*End*	*Start*	*End*	*Start*	*End*
Argentina	1977	1983			1977	1981	1973	1985	1974	1985	1971	1974
	2007	2012			2005	2012	2003	2009	2005	2009	1990	1998
											2003	2012
Bolivia	1994	1998	1974	1983	1974	1981	1973	1985	1973	1992	1973	1974
			2005	2012	2006	2012	2003	2009	2005	2008	1979	1980
											1984	1985
											2003	2012
Brazil	1965	1965			2006	2012	1973	1981			2006	2012
							2007	2009				
Chile	1994	1998	1979	1981	1979	1980	1970	1983	1979	1980	1979	1980
			1988	1989	1988	1989	2006	2009	2006	2009	1987	1991
			2006	2008	2006	2008					1994	1995
											1962	2012
Colombia	1964	1965	2008	2012	1964	1965	1973	1985	1974	1987	1976	1977
	1977	1980			1977	1980	2005	2009	2005	2009	1981	1986
					1995	1997					1995	2003
					2008	2012					2004	2012

Ecuador	1964	1965	1980	1985	1980	1981	1974	1985	1974	1985	1973	1976
	1994	1998	2005	2008	1994	1997	2005	2009	2006	2009	1977	1985
	2011	2012	2011	2012	2006	2008					2002	2012
					2011	2012						
Paraguay	1989	1990	2010	2012	2007	2011	1973	1974	1973	1974	1988	1990
	2001	2003					1979	1981			2002	2004
	2007	2008									2008	2012
Peru	1964	1966	1979	1983			1974	1985	1973	1984	1973	1974
	1994	1997	2005	2011			2005	2009	2006	2009	1979	1989
	2008	2012									1990	1992
											1994	1998
											2003	2012
Uruguay	1980	1983			1980	1983	1977	1980	1977	1981	1971	1973
	1996	1998			2008	2012					1986	1989
	2008	2012										
Venezuela			1979	1980	1979	1982	1973	1985	1974	1985	1989	1990
			2005	2008	2005	2008	1990	1992	1990	1992	1995	1996
							2004	2009	2004	2009	1999	2000
											2003	2012

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Kirsten Roach

A Structural Analysis of Oil Price Shocks on the Jamaican Macroeconomy

Abstract

This paper utilizes structural vector autoregression models to examine the impact of oil price shocks on key Jamaican macroeconomic variables over the period 1997:01-2012:06. The results indicate that oil price shocks largely do not have a permanent effect on the Jamaican economy. Furthermore, the findings suggest that an oil shock emanating from an increase in global aggregate demand generally precedes an improvement in the domestic economy while demand shocks associated with precautionary holdings of oil (oil-specific demand shocks) and oil supply shocks generally result in a deterioration in domestic macroeconomic variables.

Keywords: Oil price, vector autoregressions, oil demand shocks, oil supply shocks.

JEL classification: E31, E32, Q43.

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1. INTRODUCTION

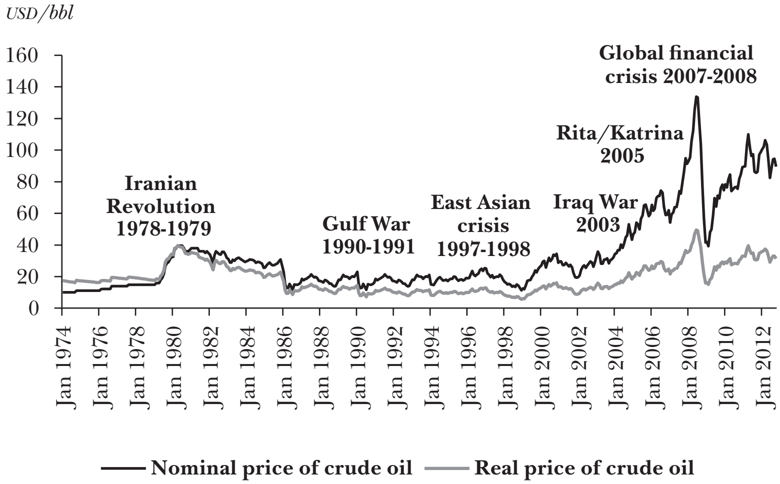
Researchers and policymakers have invariably had an intrinsic interest in commodity price movements owing to their correlation with major macroeconomic events. This interest has emerged since the 1970s when significant fluctuations in crude oil prices triggered an ongoing examination of the impact of oil price shocks on macroeconomic variables. Arguably, global macroeconomic volatility and stagflation during the 1970s and 1980s have been largely attributed to oil supply shocks (Baumeister et al., 2010). These shocks were triggered by major political and economic events such as the Iranian Revolution in 1979 and the collapse of the Organization for the Petroleum Exporting Countries (OPEC) in 1986. Since then, other shocks such as the invasion of Kuwait in 1990-1991, the Asian crisis in 1997-2000, and the global financial crisis in late 2008 have preceded increases in oil prices (see Figure 1). While much of the early literature suggested that spikes in fuel prices primarily resulted from oil supply disruptions, more recent studies indicate that the demand for oil has significantly fomented a large portion of the uptick in oil prices since the 1970s (Kilian, 2009).

Research has revealed that sharp increases in the real price of oil have had an impact on the global business cycle by affecting productivity levels and the level of real interest rates in the economy. For Jamaica, oil remains the most important raw material in various production processes. As a result, the oil bill has accounted for approximately a third of the total value of imports over the past ten years. Given the importance of oil in the production process, volatility in oil prices has major implications for domestic price stability and other macroeconomic variables. Against this background, an assessment of the relation between these shocks and the macroeconomic variables in the Jamaican economy is warranted.

This paper therefore seeks to examine the impact of oil shocks on key Jamaican macroeconomic variables, including real GDP, inflation, the nominal exchange rate, the current

Figure 1

NOMINAL VS REAL PRICE OF WTI CRUDE OIL



Source: Bloomberg L.P.

account balance, and interest rates. It is anticipated that a disaggregation of the oil price shocks would help inform policy by providing a better understanding of exactly how specific spikes in oil prices influence Jamaica's key macroeconomic variables. As aggregate demand shocks are typically associated with global economic expansion, these shocks are expected to have a positive albeit lagged impact on the Jamaican economy whereas oil-specific demand shocks emanating from speculative behavior should have adverse implications for Jamaica. While previous studies such as Burger et al. (2009) have explored the effects of oil shocks on Jamaica's external capital structures, this paper seeks to broaden the scope to include the impact on domestic macroeconomic variables. The shocks explored in this paper registered varied outcomes based on the type of disturbance. In particular, the results suggest that an oil shock emanating from an increase in aggregate demand is likely to contribute to an improvement in the domestic economy, reflecting the favorable impact of this shock on Jamaica's real output in response to gains in overall global trade. Conversely, oil-specific demand shocks and oil supply shocks would likely result in a deterioration in domestic macroeconomic variables, particularly inflation in the case of the former, largely due to increased speculation associated with this type of shock. The remainder of the paper is organized as follows. Section 2 presents stylized facts. Section 3 reviews the literature on oil price shocks and the macroeconomy. Section 4 presents the data considerations and methodology, while empirical results are discussed in Section 5. Concluding remarks and policy recommendations are presented in Section 6.

2. STYLIZED FACTS

As previously outlined in Section 1, oil plays an integral role in the Jamaican economy. In effect, fuel imports represented the largest contributor to total imports during the period 2004-2013 (see Figure 2), averaging 33% of imports. Jamaica's heightened demand for crude oil can be attributed to its use as an input in the domestic production process and electricity generation[1].

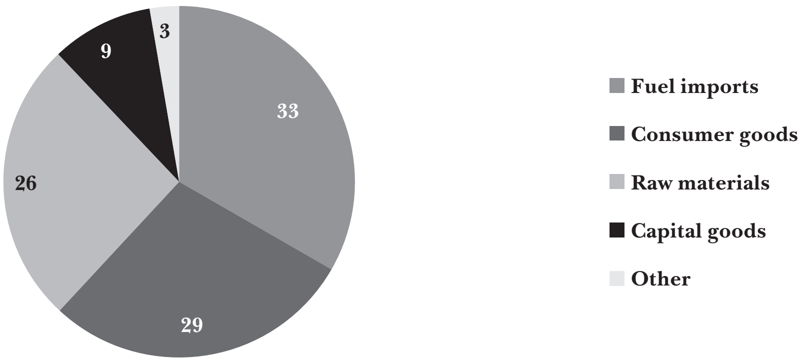
The Petroleum Corporation of Jamaica (PCJ) and bauxite companies are the primary importers of fuel in Jamaica. The PCJ purchases crude oil in accordance with the PetroCaribe Energy Accord and imports and distributes oil derivatives such as liquid petroleum gasoline (LPG), automotive diesel oil, and kerosene[2]. Notwithstanding the agreement, the West Texas Intermediate (WTI) oil price represents the relevant international benchmark for Jamaica. Thus, changes in the WTI

[1] In terms of the remaining categories, 29%, 26% and 9% of imports for that period accounted for imports of consumer goods, raw materials (excluding fuel), and capital goods, respectively.

[2] The PetroCaribe agreement is a preferential arrangement between Venezuela and 13 Caribbean islands for the purchase of oil. Jamaica has been purchasing oil under this facility since 2005.

Figure 2

JAMAICA'S MAJOR IMPORTS BY END USE
(2004-2013 average)
Percentage



Source: Bank of Jamaica.

oil price result in similar adjustments to domestic fuel prices (see Figure 3). Given the strong co-movement between WTI oil prices and Jamaica's current account deficit, an increase in WTI oil prices in 2008, for example, led to a widening of the trade deficit due to the impact of higher prices on the country's fuel bill (see Figure 4).

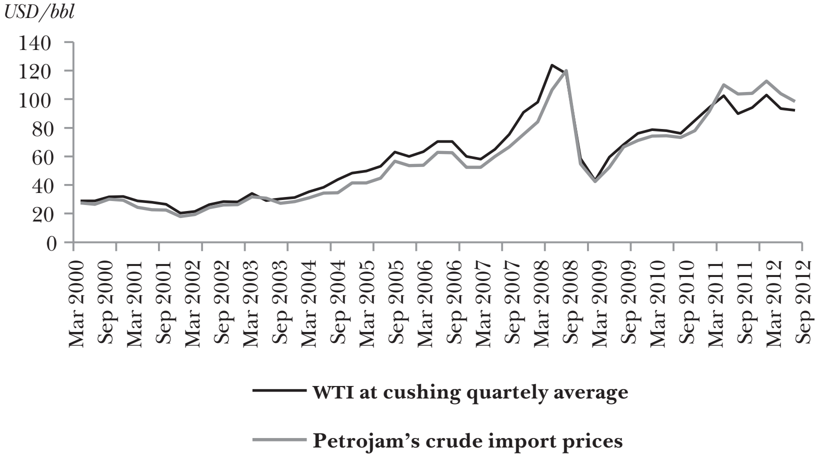
3. LITERATURE REVIEW

Studies on the relation between oil price shocks and macroeconomic variables have been widespread[3]. Hamilton (1983), in his seminal paper, highlighted that a sharp increase in crude oil prices was a precursor to seven of the eight post-war US recessions, particularly during the 1948-1972 period, based on the statistical significance of the correlation between

[3] See Barsky and Kilian (2002, 2004) and Kilian (2008, 2009, 2010).

Figure 3

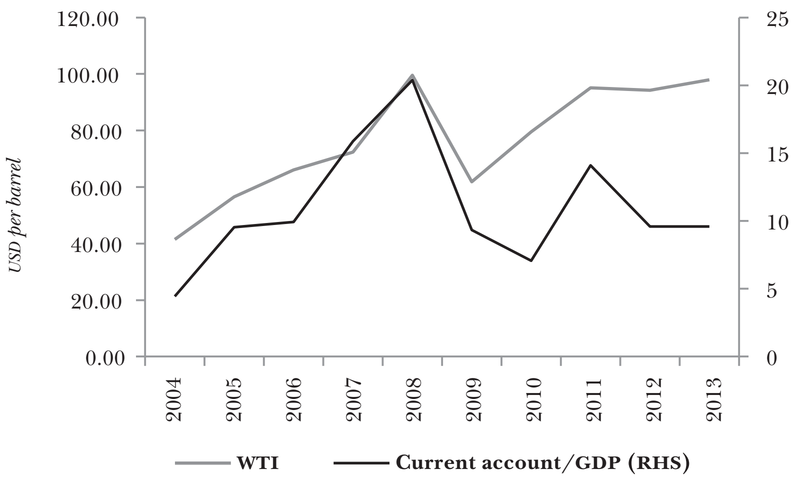
WTI CRUDE OIL PRICES AND PETROJAM'S CRUDE IMPORT PRICES



Source: Bloomberg L.P. and Bank of Jamaica

Figure 4

RELATION BETWEEN WTI OIL PRICES AND JAMAICA'S CURRENT ACCOUNT DEFICIT/GDP RATIO



Source: Bloomberg L.P. and Bank of Jamaica oil shocks and real GDP growth. He proposed three possible hypotheses: *1)* recessions coinciding with oil price increases occurred by a mere coincidence, *2)* the correlation resulted from an endogenous explanatory variable that generated both the oil price increases and the recessions, and *3)* an exogenous increase in the price of crude petroleum prompted some of the recessions in the United States before 1973. The paper concluded that the third hypothesis can be substantiated. That is, the timing, magnitude, and duration of a portion of the recessions predating 1973 would have been more severe in the absence of the oil price increase or fuel supply shortfalls.

While Hamilton (1983, 1996) and Bernanke et al. (1997) support the exogeneity of the major increases in the price of oil, research has demonstrated that there is insufficient evidence to give credence to this school of thought (see Kilian, 2008, 2009, 2010; Peersman and Van Robays, 2009; and Baumeister et al., 2010). In particular, Kilian (2008) focused on the exogeneity of oil shocks since 1973 in order to ascertain how shortfalls in oil production resulting from wars and other exogenous political events in OPEC countries affect oil prices, US real GDP growth, and US CPI inflation. He determined that increases in oil prices generally resulted in a significant contraction in US GDP five quarters subsequent to the shock and that only a miniscule proportion of the observed oil price shock resulted from exogenous disruptions to oil supplies during crisis periods. In addition, the results indicated that a sharp rise in the US CPI occurred three quarters after the exogenous oil supply shock, in contrast with the commonly held view that a sustained increase in inflation would occur.

Against this background, Kilian highlighted in 2009 that the impact of oil price shocks on the real price of oil depended on the origin of the shock. In particular, oil price shocks were decomposed under the assumption of the endogeneity of the price of oil. Kilian's approach entailed a structural decomposition of the shocks to the real price of crude oil into three categories, namely *1)* crude oil supply shocks, representing sharp increases in oil prices emanating from disruptions to crude oil production; *2)* aggregate demand shocks, reflecting increases in oil prices driven by an expansion in global economic activity; and *3)* oil-specific demand shocks, resulting from higher precautionary demand primarily due to concerns regarding near-term shortages in oil supply during periods of political unrest. In his analysis, Kilian asserted that a rise in oil prices was largely caused by positive global aggregate demand shocks as well as increased precautionary demand for oil in lieu of the actual supply disruptions. The paper estimated the relation between these shocks and the real price of oil and concluded that the type of oil shock determined the impact of higher oil prices on US real GDP and CPI inflation, a finding that also had implications for the design of national energy policy frameworks.

Baumeister et al. (2010) examined a set of industrialized economies to determine the economic consequences of oil shocks as defined by Kilian (2009) and Peersman and Van Robays (2009). Their main findings indicated that oil demand shocks associated with increased global aggregate demand resulted in a temporary increase in real GDP for all economies subsequent to an increase in oil prices. Conversely, oil-specific demand shocks were revealed to contribute to a temporary decline in real GDP[4]. Furthermore, their findings suggested that in the context of an adverse oil supply shock, net oil-importing economies all encountered a permanent contraction in real GDP, while the impact was insignificant or positive for net oil-exporting economies. The results for the pass-through to inflation were varied among oil-importing economies. Notwithstanding this variation, the results indicated that the pass-through to inflation in an oil-importing economy was contingent on second-round effects largely reflected in upward movements in wages, while the pass-through in an oil-exporting economy

[4] Aggregate demand shocks are associated with an expansion in global economic activity, while oil-specific demand shocks represent a demand shock specific to the oil market whereby growth in precautionary demand for fuel results from increased fears of future fuel supply shortages.

was limited largely in the context of the appreciation of the effective exchange rates following an oil supply shock. The paper also revealed reduced vulnerability to oil shocks in the case of economies with a favorable net energy position.

Other studies have sought to examine the relation between oil shocks and the current account balance in oil-importing and exporting countries. In the case of Turkey, an oil-importing economy, Ozlale and Pekkurnaz (2010) used a structural vector autoregression (SVAR) model to assess the impact of oil price shocks on the current account deficit. The results showed that the current account deficit to GDP ratio increased gradually in response to an oil price shock within the first three months before declining, which indicated that oil price shocks have a significant effect in the short run. Similarly, the discussion in Chuku et al. (2011) utilized a SVAR over the period 1970 to 2008 to assess the relation between oil price shocks and current account dynamics in Nigeria, an oil exporter and importer. Oil price shocks had a significant positive effect on current account deficits for Nigeria in the short run. As such, the policy implications for garnering of the benefits associated with oil price shocks on the Nigerian economy included increased emphasis on reserve-augmenting strategies, lax monetary policy, and heightened international financial integration.

In relation to the Caribbean, Burger et al. (2009) examined the possibility that a country's external capital structure could dampen the impact of oil price shocks on external accounts[5]. The economies analyzed were highly vulnerable to oil price shocks, particularly an oil-importer such as Jamaica and an oil-exporter, Trinidad and Tobago. The findings demonstrated that Jamaica's external capital structure is highly vulnerable given the country's high debt-to-GDP ratio and substantial negative foreign exchange exposure. Against this background, Burger et al. (2009) recommended that Jamaica should adjust

[5] External capital structure can be defined as the composition of foreign assets and liabilities according to instrument, currency, and maturity.

the composition of its net international reserves (NIR) portfolio with a view to stimulating capital gains in the event of adverse oil market shocks[6]. In this regard, the paper suggested the adoption of an official reserves portfolio that is positively correlated with oil prices[7]. Conversely, Burger et al. (2009) indicated that although Trinidad and Tobago's capital structure was not vulnerable to currency fluctuations, there was still room to mitigate the impact of oil shocks on the country's external accounts by hedging against the macroeconomic effects of such shocks. Thus, Trinidad and Tobago could augment capital gains amid oil shocks by modifying the structure of its NIR portfolio to incorporate an increased exposure to foreign assets that have a negative correlation with movements in oil prices.

4. METHODOLOGY AND DATA CONSIDERATIONS

Using the methodology of Kilian (2009), the impact of oil price shocks on the Jamaican economy was estimated via two main steps during the period from January 1997 to June 2012. The first step involved the examination of movements in the real price of crude oil in order to determine the underlying demand and supply shocks that affect the crude oil market. This step will be outlined in Section 4.1. The second step encompassed the estimation of the response of key Jamaican macroeconomic variables to the identified structural shocks in Section 4.2. In this context, individual SVAR models were estimated in order to assess the response of the respective macroeconomic variables under study to the shocks.

[6] Capital gains are the differences between changes in the net foreign asset position and the current account balance.

[7] For example, the official reserves portfolio could be positively correlated with the currencies of oil exporting countries such as Norway and Canada in order to increase capital gains from oil price shocks.

4.1 Determining the Underlying Demand and Supply Shocks that Affect the Crude Oil Market

In undertaking the first step highlighted above, a multivariate SVAR model was estimated utilizing monthly data over the sample period January 1997 to June 2012 for the vector time series, $z_t = (\Delta prod_t, rea_t, rpo_t)'$ where $\Delta prod_t$ represents the percent change in the production of crude oil globally, rea_t is a measure of global real economic activity in industrial commodity markets, and rpo_t is the real price of crude oil using the WTI benchmark, with rea_t and rpo_t being expressed in logs. The period of study was chosen to encompass the various oil shocks both before and after the 2008 global financial crisis. The assessment period was also determined by the availability of data.

The term global *real economic activity* refers to an index of real economic activity that measures industrial commodity markets and is used in lieu of the broadly understood concept of real economic activity associated with world real GDP or industrial output. Borrowing from Kilian (2009), this study employs a measure of global real economic activity in commodity markets. This global index comprises dry cargo single voyage freight rates for bulk dry cargoes including grain, oilseeds, coal, iron ore, fertilizer, and scrap metal, compiled by Drewry Shipping Consultants Ltd. The subsequent steps for constructing the index involve deflating the series with the US CPI. The real index was in turn detrended in order to capture cyclical variation in ocean freight rates. This measure was adopted largely due to the availability of data at a monthly frequency as well as the failure of measures of value added to capture demand in commodity markets[8]. The oil data was garnered from the US Energy Information Administration (EIA) and the International Energy Agency (IEA). The real price of oil is

[8] Of note, this measure of crude oil prices represents the best proxy for the free market global price of imported crude oil in the literature. See Kilian (2009) for a full discussion of the rationale and construction of this index.

measured using WTI oil prices deflated by the US CPI. Data on Jamaican macroeconomic variables were obtained from the Bank of Jamaica's database.

The model utilized a lag length of two months based on the criteria selection [sequential modified LR test statistic (LR), final prediction error (FPE), Akaike information criterion (AIC), and Hannan-Quinn information criterion (HQ)], for which the SVAR representation of the model consisting of a vector of serially and mutually uncorrelated structural innovations, ε_t may be seen below:

$$A_0 z_t = \propto + \sum_{i=1}^{2} A_i z_{t-i} + \varepsilon_t \,. \tag{1}$$

The structural innovations were generated by imposing exclusion restrictions on A_0^{-1}. Fluctuations in the real price of oil were underpinned by three structural shocks: ε_{1t}, which captures crude oil supply shocks; ε_{2t}, which denotes aggregate demand shocks; and ε_{3t}, which represents a demand shock specific to the oil market. The last of the three was geared toward capturing shifts in precautionary demand for fuel that coincided with increased concerns regarding the availability of future oil supplies.

Under the assumption that z_t will respond to shocks to each variable in the vector, additional restrictions were imposed. In terms of the restrictions on A_0^{-1}, it was assumed that:

1. $a_{12} = 0$ and $a_{13} = 0$, an assumption that imposes the restriction of no response in crude oil production to aggregate demand shocks and oil-specific demand shocks, respectively, within the same month. This restriction is imposed on the premise that there are high costs associated with increasing oil production and as such that only a persistent rise in demand is expected to significantly increase the supply of crude oil.

2. $a_{23} = 0$, which assumes that an increase in the real price of oil emanating from oil-specific demand shocks will

not reduce global real economic activity in industrial commodity markets within the month.

Notably, innovations to the real price of oil that cannot be explained by oil supply shocks or aggregate demand shocks must be the result of demand shocks that are specific to the oil market.

The foregoing assumptions yielded a recursively identified model with reduced form errors, $e_t = A_0^{-1}\ \varepsilon_t$ of the form:

$$e_t = \begin{pmatrix} e_t^{\Delta prod} \\ e_t^{rea} \\ e_t^{rpo} \end{pmatrix} = \begin{bmatrix} a_{11} & 0 & 0 \\ a_{21} & a_{22} & 0 \\ a_{31} & a_{32} & a_{33} \end{bmatrix} \begin{pmatrix} \varepsilon_t^{oil\ supply\ shock} \\ \varepsilon_t^{aggregate\ demand\ shock} \\ \varepsilon_t^{oil\text{-}specific\ demand\ shock} \end{pmatrix}. \tag{2}$$

4.2 Estimating the Response of Jamaican Macroeconomic Variables to Oil Price Shocks

An examination of the impact of crude oil demand and supply shocks on the Jamaican economy necessitated estimations of the relation between the structural innovations in Equation 1 and selected Jamaican macroeconomic variables. This study builds on the work done by Kilian (2009), which only focused on the impact of oil shocks on GDP and inflation, by including additional macroeconomic variables to provide a more holistic analysis of the impact of oil shocks on the Jamaican economy in individual SVAR models aimed at ascertaining the response of the respective macroeconomic variables to each oil price shock. As a result, the variables under analysis include real GDP ($\Delta\, y_t$), the quarterly point-to-point inflation rate (π_t), the quarterly end of period (e.o.p.) nominal exchange rate between the US dollar and the local currency (XR_t), the quarterly e.o.p 180-day Treasury Bill yield (IR_t) represented in differences, as well as a measure of Jamaica's external accounts, the current account balance (CA_t), expressed in log differences[9]. In order to facilitate the inclusion of quarterly variables such as real GDP in this analysis as well as maintain the identifying assumptions, quarterly shocks were constructed by averaging the monthly structural innovations implied by the VAR model in Equation 1 for each quarter:

$$\hat{\zeta}_{jt} = \frac{1}{3}\sum_{i=1}^{3}\hat{\varepsilon}_{j,t,i}, \quad j = 1,\ldots,3, \tag{3}$$

where $\hat{\varepsilon}_{j,t,i}$ is the estimated residual for the jth structural shock in the ith month of the tth quarter of the sample.

These shocks were treated as exogenous based on the identifying assumption of no feedback from Δy_t, π_t, XR_t, IR_t, and CA_t to $\hat{\zeta}_{jt}$, j = 1,..., 3 within a given quarter. In this context, the dynamic effects of the shocks on Jamaica's real gdp, inflation, exchange rate, interest rate, and current account deficit, respectively, were examined based on five individual quarterly regressions of the form and lag length selection criteria in Equations 4-8, respectively:

$$\Delta y_t = \alpha + \sum_{i=0}^{1}\phi_i \zeta_{jt-i} + u_t, \text{ j = 1,..., 3 (real GDP SVAR)} \tag{4}$$

$$\pi_t = \delta + \sum_{i=0}^{1}\psi_i \zeta_{jt-i} + v_t, \text{ j = 1,..., 3 (inflation SVAR)} \tag{5}$$

[9] The 180-day Treasury Bills (T-Bills) yield was utilized in this study, as BOJ does not have a policy rate that consistently captures monetary policy actions. For example, in September 2000, BOJ introduced 270 and 360-day tenors with higher margins but did not increase rates. Similarly, in November 2008, BOJ tightened policy by introducing a special 180-day certificate of deposit at 20.5% but did not increase rates on its other instruments. Rates on 180-day omo instruments remained at 15.35%, while there was an increase in yields on 180-day T-Bills. There have also been several instances when the longer-term rates were increased but the shorter-term rates were unchanged. In all instances, yields on T-Bills responded to the policy actions. T-bills also capture market sentiment.

6 $$XR_t = \beta + \sum_{i=0}^{1} \varphi_i \zeta_{jt-i} + w_t, \text{j} = 1,\ldots, 3 \text{ (exchange SVAR)}$$

7 $$IR_t = \gamma + \sum_{i=0}^{1} \omega_i \zeta_{jt-i} + z_t, \text{j} = 1,\ldots, 3 \text{ (interest rate SVAR)}$$

8 $$CA_t = \theta + \sum_{i=0}^{1} \rho_i \zeta_{jt-i} + x_t, \text{j} = 1,\ldots, 3 \text{ (current account SVAR)},$$

where u_t, v_t,w_t, x_t, z_t were potentially serially correlated errors while ζ_{jt} was a serially uncorrelated shock. The respective impulse response coefficients were denoted as ϕ_i, ψ_i, φ_i, ω_i and ρ_i.

The equation-by-equation approach shown in Equations 4-8 is consistent with the premise that the quarterly shocks $\hat{\zeta}_{jt}$, j = 1,...,3, are mutually uncorrelated. In essence, despite the potential existence of some omitted variable bias, the particularly low contemporaneous correlations between the quarterly shocks and autoregressive residuals of the selected macroeconomic variables permitted the quarterly shocks to be treated as orthogonal or uncorrelated. Notably, low correlations in turn gave credence to the estimation of separate equations for each shock (see Table 1). The equation-by-equation approach was deemed the most parsimonious in assessing the impact of oil shocks on macroeconomic variables. This conclusion is based on an examination of additional investigations by Kilian et al. (2009) of alternative methodologies comprising the estimation of equivalent Equations 4-8, which included current and lagged values of all shocks. To the extent that there was a lack of data availability given the need for five lags for each shock, this alternative approach was found to be unsuitable. Another alternative entailed the addition of lagged dependent variables as regressors in Equations 4-8. Since strict exogeneity of $\hat{\zeta}_{jt}$ with respect to each macroeconomic variable was a necessary condition for this alternative, it was found to be infeasible for the purposes of the study as such a condition would eliminate the effects of shocks on the macroeconomic variable (Kilian, 2009). In this regard, the equation-by-equation approach was found to be the most viable methodology.

Table 1

CONTEMPORANEOUS CORRELATION OF QUARTERLY SHOCKS WITH AUTOREGRESSIVE RESIDUALS FOR SELECTED JAMAICA MACROECONOMIC VARIABLES

	Oil supply shock	*Aggregate demand shock*	*Oil-specific demand shock*
Real GDP	0.009	0.395	0.135
Inflation	–0.320	0.176	–0.161
Exchange rate	–0.218	0.273	0.307
Interest rate	–0.118	0.095	0.056
Current account	0.150	0.082	0.204

5. DISCUSSION OF RESULTS

With the incorporation of the quarterly structural innovations into the five quarterly VAR models as shown in Equations 4-8, the results of the impact of the three oil price shocks on macroeconomic variables could be analyzed. These shocks were generated by aggregating the monthly disturbances from Equation 1 for each quarter over the sample period from the first quarter of 1997 to the second quarter of 2012. The augmented Dickey-Fuller test was employed to verify the existence of a unit root in the variables. The results indicated that all variables, excluding the inflation rate and the interest rates, possessed a unit root (see Table 2). Notwithstanding, the results of the stability tests for all variables revealed that no root lies outside of the unit circle, reflecting the satisfaction of the VARs' stability conditions (see Figure 5). Further robustness checks on the VARs based on the portmanteau tests for autocorrelations revealed

Table 2

UNIT ROOT TESTS
(Augmented Dickey-Fuller *t*-statistic)

	Level		*1st difference*		*Degree of Integration*
	t-statistic	*P-value*	*t-statistic*	*P-value*	
Real GDP	−2.5622	0.1068	−19.2779	0.0000	I(1)
Inflation rate	−5.5254	0.0000	–	–	I(0)
Exchange rate	−1.0604	0.7258	−4.8191	0.0002	I(1)
Interest rate	−8.0892	0.0000	–	–	I(0)
Current account	−2.6428	0.0902	−13.1600	0.0000	I(1)

Notes: Lag lengths in the ADF regressions were chosen using the Bayesian information criterion. Asymptotic critical values are: 1 percent, −3.51; 5 percent, −2.89; 10 percent, −2.58.

that the residuals were serially uncorrelated (see Tables 3-7). The impulse response functions are reported in Figures 6 to 10 using both the 95% and 68% confidence intervals. Of note, the responses of Jamaica's macroeconomic variables under study to all three shocks were identical irrespective of the confidence bands utilized. Nevertheless, while the majority of the responses were statistically significant based on the 68% confidence interval, most were not for the 95% confidence interval[10].

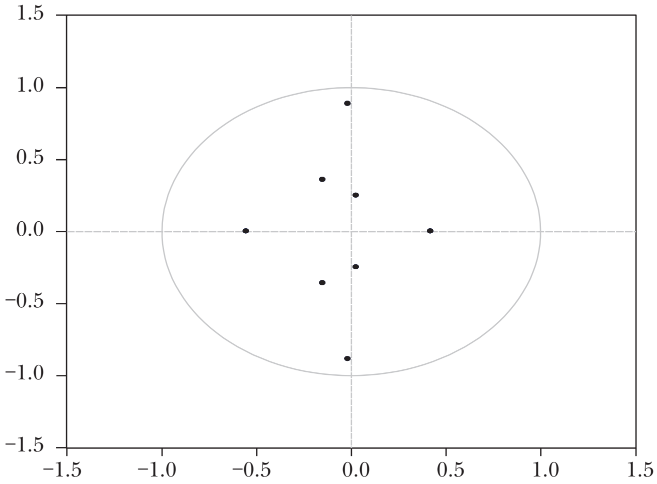
The impact of both oil demand and supply shocks on real GDP failed to dissipate in the short term, albeit having a marginal impact on domestic output (see Figure 6). The initial response of real GDP was a contraction under an oil supply shock

[10] Sims and Zha (1999) endorse the use of 68% confidence intervals for the purposes of impulse responses and argue that "there is no scientific justification for reporting hypotheses at the 5% significance level in every application."

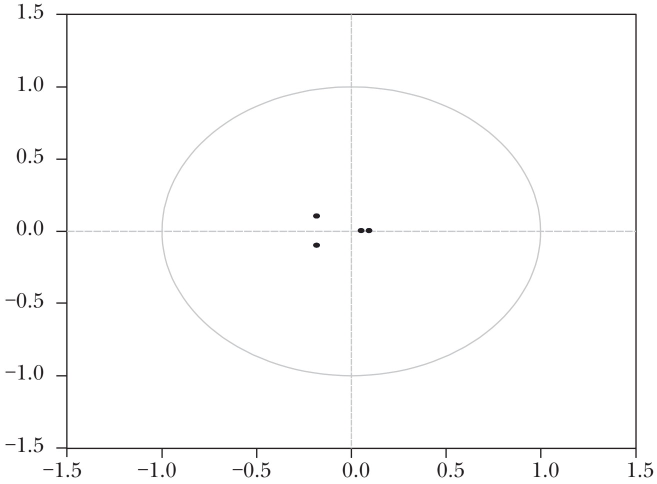
Figure 5

STABILITY CONDITION TESTS

A. REAL GDP[1]



B. INFLATION[1]

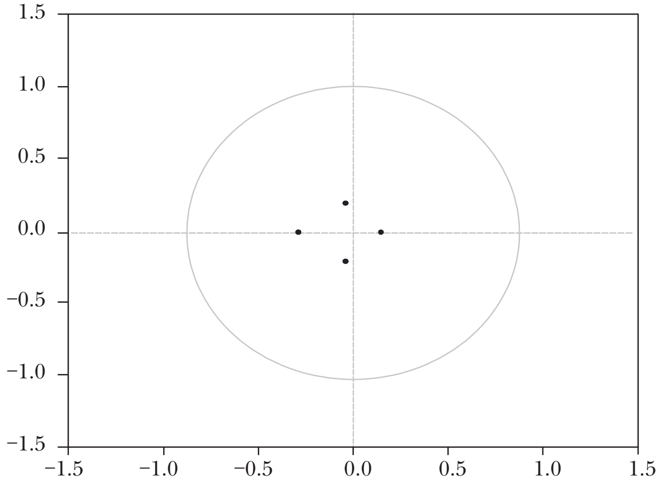


[1] Inverse roots of AR characteristic polynomial.
Sources: Bloomberg L.P. and Bank of Jamaica.

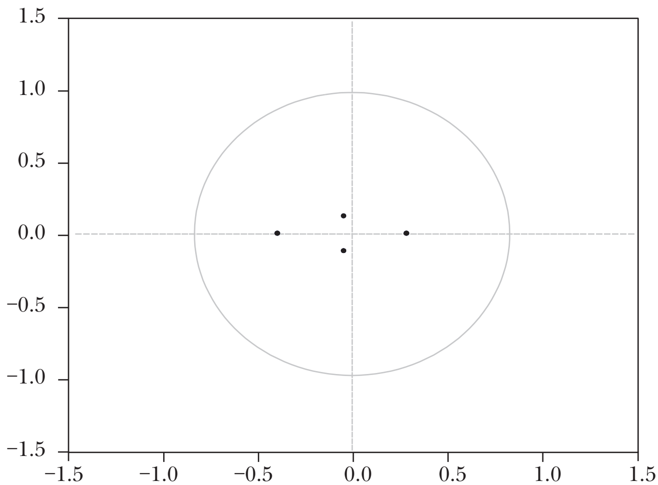
Figure 5 (cont.)

STABILITY CONDITION TESTS

C. EXCHANGE RATE[1]



D. INTEREST RATE[1]

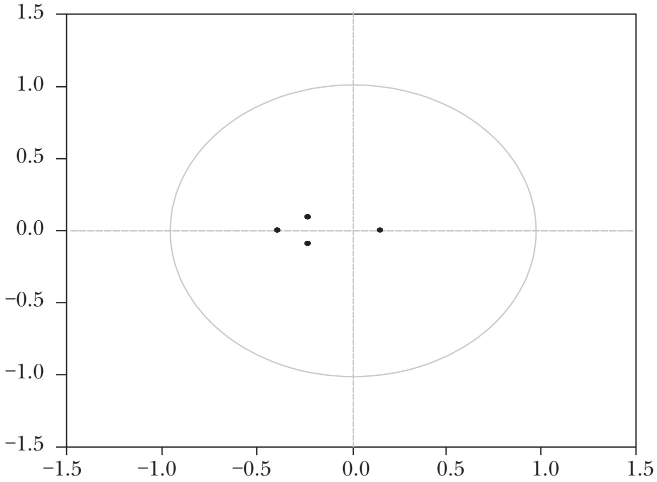


[1] Inverse roots of AR characteristic polynomial.
Sources: Bloomberg L.P. and Bank of Jamaica.

Figure 5 (cont.)

STABILITY CONDITION TESTS

E. CURRENT ACCOUNT[1]



[1] Inverse roots of AR characteristic polynomial.
Sources: Bloomberg L.P. and Bank of Jamaica.

Table 3

REAL GDP AUTOCORRELATION TEST
VAR residual portmanteau tests for autocorrelations

Lags	*Q-stat*	*Prob.*	*Adj. Q-stat*	*Prob.*	*df*
1	8.525442	NA[1]	8.672433	NA[1]	NA[1]
2	17.32332	NA[1]	17.77901	NA[1]	NA[1]
3	37.74929	0.1280	39.29923	0.0961	29
4	52.56783	0.2043	55.19548	0.1419	45

Notes: [1]The test is valid only for lags larger than the VAR lag order.
df is degrees of freedom for (approximate) chi-square distribution.
df and Prob. may not be valid for models with exogenous variables.

Table 4

INFLATION AUTOCORRELATION TEST

VAR Residual Portmanteau Tests for Autocorrelation

Lags	*Q-stat*	*Prob.*	*Adj. Q-stat*	*Prob.*	*df*
1	11.86208	NA[1]	12.06313	NA[1]	NA[1]
2	26.13026	0.6185	26.82332	0.5812	29
3	44.25690	0.5033	45.90399	0.4345	45
4	62.17170	0.4342	65.09842	0.3361	61

Notes: [1]The test is valid only for lags larger than the VAR lag order.
df is degrees of freedom for (approximate) chi-square distribution.
df and Prob. may not be valid for models with exogenous variables.

Table 5

EXCHANGE RATE PORMANTEAU AUTOCORRELATION TEST

Lags	*Q-stat*	*Prob.*	*Adj. Q-stat*	*Prob.*	*df*
1	10.94135	NA[1]	11.12680	NA[1]	NA[1]
2	30.41066	0.3937	31.26746	0.3529	29
3	48.29284	0.3413	50.09081	0.2785	45
4	64.10392	0.3682	67.03125	0.2780	61

Notes: [1]The test is valid only for lags larger than the VAR lag order.
df is degrees of freedom for (approximate) chi-square distribution.
df and Prob. may not be valid for models with exogenous variables.

Table 6

INTEREST RATE PORMANTEAU AUTOCORRELATION TEST

Lags	*Q-stat*	*Prob.*	*Adj. Q-stat*	*Prob.*	*df*
1	9.720715	NA[1]	9.885473	NA[1]	NA[1]
2	34.05432	0.2373	35.05817	0.2026	29
3	48.72004	0.3257	50.49576	0.2654	45
4	61.47620	0.4588	64.16308	0.3663	61

Notes: [1]The test is valid only for lags larger than the VAR lag order.
df is degrees of freedom for (approximate) chi-square distribution.
df and Prob. may not be valid for models with exogenous variables.

Table 7

CURRENT ACCOUNT PORMANTEAU AUTOCORRELATION TEST
VAR residual portmanteau tests for autocorrelations

Lags	*Q-stat*	*Prob.*	*Adj. Q-stat*	*Prob.*	*df*
1	9.425405	NA[1]	9.585158	NA[1]	NA[1]
2	29.09564	0.4601	29.93367	0.4173	29
3	45.95350	0.4325	47.67879	0.3643	45
4	62.11750	0.4361	64.99737	0.3393	61

Notes: [1]The test is valid only for lags larger than the VAR lag order.
df is degrees of freedom for (approximate) chi-square distribution.
df and Prob. may not be valid for models with exogenous variables.

and an oil-specific demand shock. However, both shocks were mostly statistically insignificant at the 5% level. In contrast, an aggregate demand shock resulted in an initial expansion in domestic output that was statistically significant at the 5% level. Notably, the responses of real GDP to all three shocks are significant using the 68% confidence interval. Though higher oil prices emanate from an aggregate demand shock, other factors including gains from international trade arising from increased global demand can influence the response of real GDP to the oil price shift[11]. Additional statistical analysis has shown that over the period 1997-2012, crude oil prices had a weak linear relation with output in Jamaica, as evidenced by the low positive correlation of 0.1. While most research findings indicate at least a negative correlation between the two variables, the low positive correlation could, however, be attributed to particular factors affecting the local economy. Some of these factors include Jamaica's high inelastic fuel demand, which indicates that irrespective of the directional movement in oil prices, Jamaica's dependence on the commodity is necessary for domestic production.

Regarding the response of inflation to an oil supply shock, inflation declined temporarily during the first two quarters with no impact observed thereafter. The result was statistically insignificant at the 5% level but significant using the 68% confidence bands (see Figure 7). As a result, policymakers need not be concerned about the impact of supply disruptions in major oil producing countries on Jamaica's inflation in the short term. This outcome can be ascribed to the fact that supply disruptions in one area typically result in increased oil production in other regions to compensate for the shortfall. In contrast, the impact of an aggregate demand shock led to an acceleration in inflation by the third quarter, albeit statistically insignificant at both the 95% and 68% levels. Oil-specific demand shocks resulted in an initial acceleration in inflation within the first two quarters prior to decelerating by the fourth quarter.

[11] See Baumeister et al. (2010).

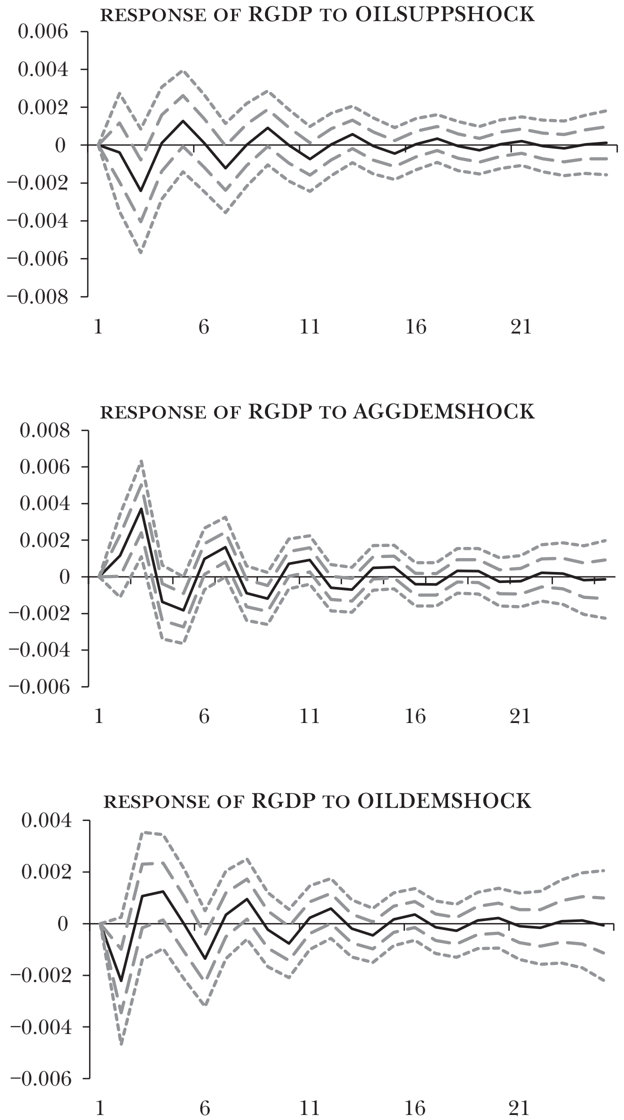
This result was statistically significant at both the 95% and 68% confidence intervals. A temporary spike in inflation indicates the need for the possible implementation of short-term policy measures to stem an increase in other prices such as wages.

In terms of the nominal exchange rate, there was a marginal depreciation following an oil supply shock, although statistically insignificant at both confidence levels under analysis (see Figure 8). Similarly, an aggregate demand shock engendered a depreciation of the domestic currency, particularly within the first two quarters, which was statistically significant at both confidence levels. Some investors, based on ignorance of the source of the shock, may initially respond by increasing the demand for foreign currency for portfolio rebalancing. In addition, there could be an expansion in demand for foreign currency for current account transactions as investors increase the input in the production process to meet the growth in external demand. This depreciation, however, dissipated by the third quarter, possibly reflecting the impact of the improvements in Jamaica's major trading partners on foreign currency earnings in the domestic economy. Similarly, an oil-specific demand shock led to depreciation in the exchange rate within the first two quarters. This result is in keeping with the notion that uncertainty in the oil market leads to possible hoarding or speculative behavior by local investors. This impact was, however, statistically insignificant at the 5% level but was found to be significant using the 68% confidence interval.

Regarding interest rates, impulse responses indicated an increase in market interest rates within the first four quarters following an oil supply and oil-specific demand shock (see Figure 9). While the impact was statistically significant in the case of the oil-specific demand shock, the converse holds as it relates to the oil supply shock at each level of significance under study. In response to an aggregate demand shock, interest rates fell initially but increased by the third quarter. The effect of this shock on interest rates was not significant at the 5% level. Of note, however, the 68% error bands yielded a significant response in the second quarter.

Figure 6

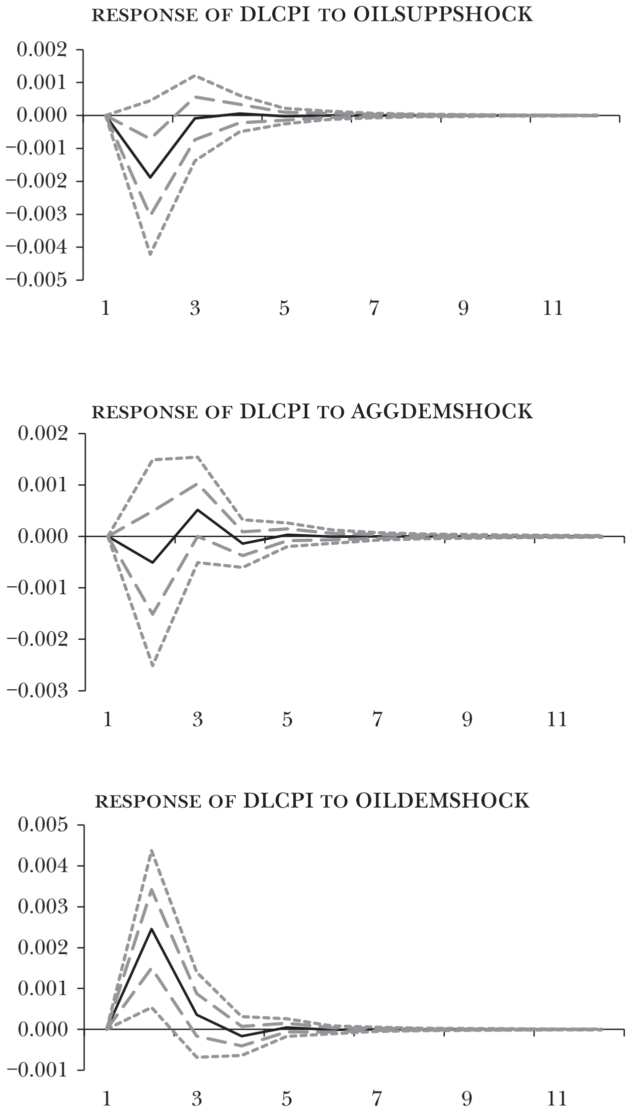
RESPONSE OF REAL GDP TO ONE-STANDARD DEVIATION OIL SHOCKS



Notes: Estimates based on a quarterly VAR (2) system in Equation 3. OILSUPPSHOCK, AGGDEMSHOCK, OILDEMSHOCK and DRGDP represent oil supply shocks, aggregate demand shocks, oil-specific demand shocks, and real GDP growth. Dotted lines are 95% confidence intervals while dashed lines are 68% confidence intervals.

Figure 7

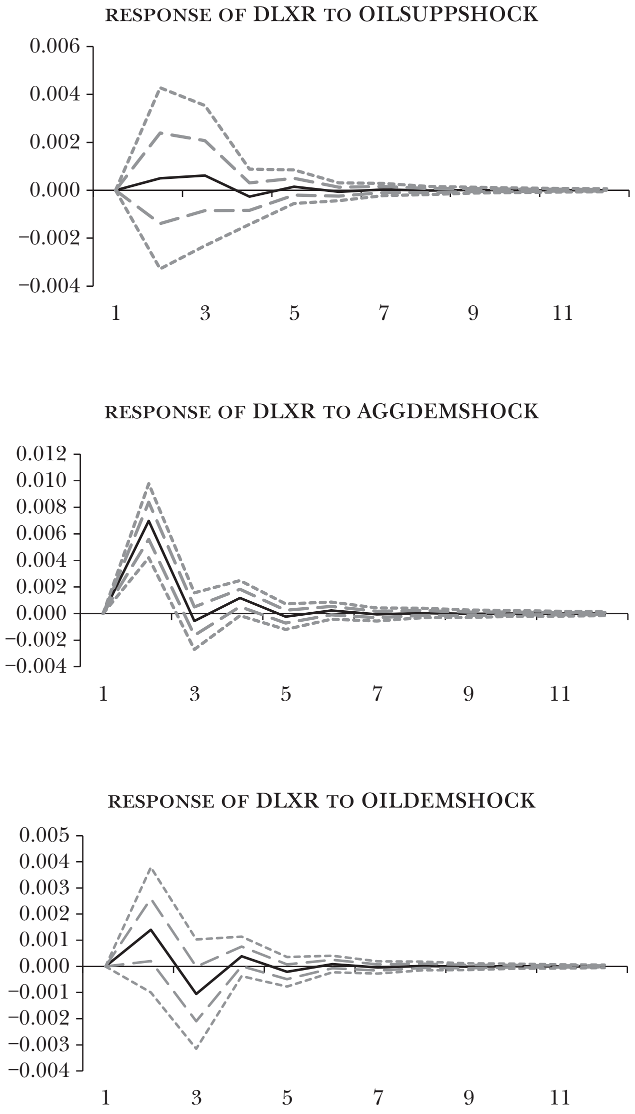
RESPONSE OF INFLATION TO ONE-STANDARD DEVIATION OIL SHOCKS



Notes: Estimates based on a quarterly VAR (1) system in Equation 3. OILSUPPSHOCK, AGGDEMSHOCK, OILDEMSHOCK and DLCPI represent oil supply shocks, aggregate demand shocks, oil-specific demand shocks, and inflation. Dotted lines are 95% confidence intervals while dashed lines are 68% confidence intervals.

Figure 8

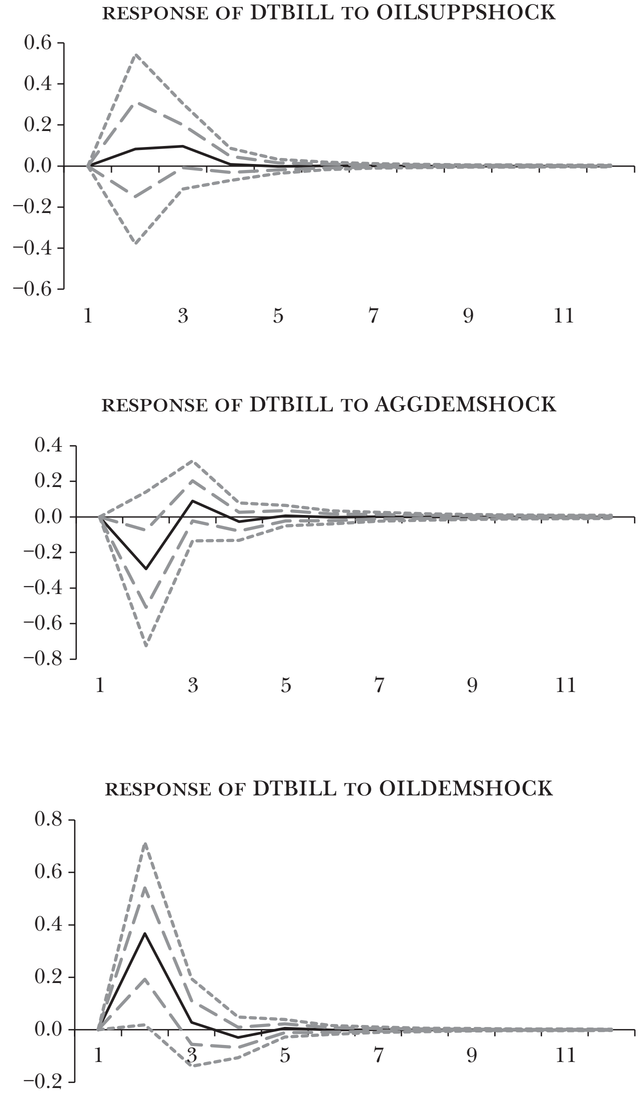
RESPONSE OF EXCHANGE RATE TO ONE-STANDARD DEVIATION OIL SHOCKS



Notes: Estimates based on a quarterly VAR (1) system in Equation 3. OILSUPPSHOCK, AGGDEMSHOCK, OILDEMSHOCK and DLXR represent oil supply shocks, aggregate demand shocks, oil-specific demand shocks, and the nominal exchange rate. Dotted lines are 95% confidence intervals while dashed lines are 68% confidence intervals.

Figure 9

RESPONSE OF EXCHANGE RATE TO ONE-STANDARD DEVIATION OIL SHOCKS

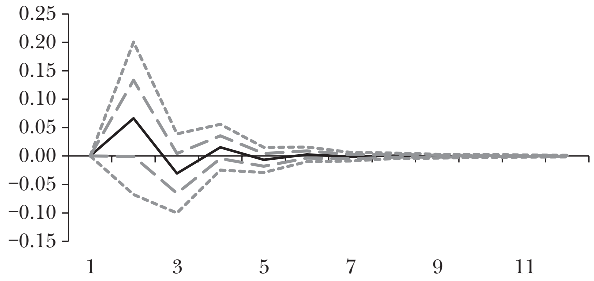


Notes: Estimates based on a quarterly VAR (1) system in Equation 3. OILSUPPSHOCK, AGGDEMSHOCK, OILDEMSHOCK, and DTBILL represent oil supply shocks, aggregate demand shocks, oil-specific demand shocks, and the 180-day Treasury bill interest rate. Dotted lines are 95% confidence intervals while dashed lines are 68% confidence intervals.

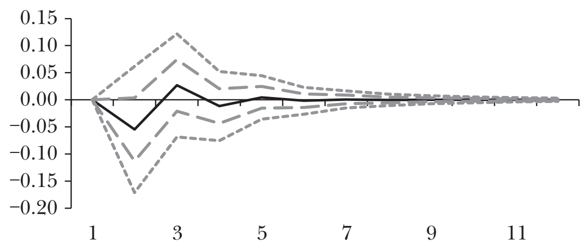
Figure 10

RESPONSE OF CURRENT ACCOUNT TO ONE-STANDARD DEVIATION OIL SHOCKS

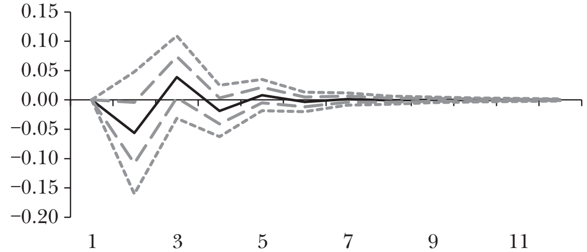
RESPONSE OF DLABSCASA TO OILSUPPSHOCK



RESPONSE OF DLABSCASA TO AGGDEMSHOCK



RESPONSE OF DLABSCASA TO OILDEMSHOCK



Notes: Estimates based on a quarterly VAR (1) system in Equation 3. OILSUPPSHOCK, AGGDEMSHOCK, OILDEMSHOCK and DLABSCASA represent oil supply shocks, aggregate demand shocks, oil-specific demand shocks, and the seasonally adjusted current account deficit. Dotted lines are 95% confidence intervals while dashed lines are 68% confidence intervals.

As for the response of Jamaica's external accounts to an oil supply shock, the current account deficit increased within the first two quarters (see Figure 10). This result could be associated with the initial high fuel prices generally stemmed from the prospect of reduced oil supplies, which in turn leads to an increase in the value of imports and hence an overall deterioration in the trade balance. As other oil producers augment supplies and some countries cut demand, fuel prices fall, which then leads to a reduction in the deficit by the third quarter. In contrast, aggregate demand and oil-specific demand shocks resulted in lower current account deficits within the first two quarters, but this impact was reversed by the third quarter. The initial reduction in the deficit may be attributed to the impact of the gains from global economic activity, which offset the impact of the higher prices of oil. The responses of the current account deficit to the oil supply and aggregate demand shocks were statistically insignificant at the 95% and 68% confidence intervals. However, the response of the current account deficit to an oil-specific demand shock was significant at the 68% confidence interval (see Table 8).

In an effort to delve more deeply into the extent to which each shock contributed to the responses by the respective macroeconomic variables, variance decompositions were conducted (see Tables 9-13)[12]. With respect to the effect of the oil supply shock on real GDP, inflation, the exchange rate, the interest rate, and the current account deficit, variance decompositions indicated that this shock accounted for 4.2%, 4.9%, 0.4%, 0.7%, and 2%, respectively, of the movements in each variable by the third quarter. Overall, this shock is show to have the smallest impact since it accounts for only a small percentage of the variation in the different macroeconomic variables.

[12] While impulse response functions trace the effects of a shock to one endogenous variable on the other variables in the VAR, variance decomposition separates the variation in an endogenous variable into the component shocks to the VAR. Thus, the variance decomposition provides information about the relative importance of each random innovation in affecting the variables in the VAR.

Table 8

SUMMARY OF IMPULSE RESPONSES

	Real GDP	*Inflation*	*Exchange rate*	*Interest rate*	*Current account deficit*
Oil supply shock	↓	↓	↑	↑	↑
Aggregate demand shock	↑[a]	↑	↑[a]	↑	↓
Oil-specific demand shock	↓	↑[a]	↑	↑[a]	↓

[a] Denotes rejection using the 95% confidence bands.

Regarding the effect of the aggregate demand shock on real GDP, inflation, the exchange rate, the interest rate, and the current account deficit, the respective variance decompositions highlighted that this shock contributed to 10.5%, 0.7%, 26%, 4%, and 1.5%, respectively, of movements by the third quarter. Despite the results from the impulse response, which suggest an eventual acceleration in inflation, the variance decomposition indicates the negligible importance of the shock to inflation and the current account deficit.

As for the oil-specific demand shock, variance decompositions demonstrated that 5.2%, 8.5%, 1.6%, 6% and 2% of movements in real GDP, inflation, the exchange rate, the interest rate, and the current account deficit, respectively, can be attributed to this shock within the first three quarters. The results suggest the relatively high importance of this shock to inflation in the context of Jamaica's economy.

Table 9

VARIANCE DECOMPOSITION OF REAL GDP

Period	*S.E.*	*Real GDP*	*Oil supply shock*	*Aggregate demand shock*	*Oil-specific demand shock*
1	0.009369	100.0000	0.000000	0.000000	0.000000
2	0.009705	93.20463	0.156556	1.419289	5.219524
3	0.011984	81.10716	4.147979	10.52971	4.215156
4	0.012127	79.22231	4.058926	11.56134	5.157421
5	0.013240	79.74182	4.338183	11.59270	4.327297
6	0.013353	78.51046	4.271593	11.93259	5.285358
7	0.013983	78.24095	4.646024	12.23443	4.878598
8	0.014050	77.59154	4.602327	12.51957	5.286565
9	0.014418	77.62504	4.768557	12.56058	5.045823
10	0.014463	77.23827	4.739356	12.72456	5.297810

Notes: Cholesky ordering- real GDP, oil supply shock, aggregate demand shock, oil-specific demand shock. Standard errors: Monte Carlo (10,000 repetitions).

Table 10

VARIANCE DECOMPOSITION OF INFLATION

Period	*S.E.*	*Inflation*	*Oil supply shock*	*Aggregate demand shock*	*Oil-specific demand shock*
1	0.007709	100.0000	0.000000	0.000000	0.000000
2	0.008434	86.18130	4.960546	0.368265	8.489888
3	0.008458	85.70174	4.943022	0.736347	8.618890
4	0.008461	85.64169	4.944264	0.763648	8.650395
5	0.008461	85.63743	4.944630	0.764876	8.653061
6	0.008461	85.63727	4.944652	0.764905	8.653176
7	0.008461	85.63726	4.944653	0.764906	8.653179
8	0.008461	85.63726	4.944653	0.764906	8.653179
9	0.008461	85.63726	4.944653	0.764906	8.653179
10	0.008461	85.63726	4.944653	0.764906	8.653179

Notes: Cholesky ordering- inflation, oil supply shock, aggregate demand shock, oil-specific demand shock. Standard errors: Monte Carlo (10,000 repetitions).

Table 11

VARIANCE DECOMPOSITION OF EXCHANGE RATE

Period	*S.E.*	*Exchange rate*	*Oil supply shock*	*Aggregate demand shock*	*Oil-specific demand shock*
1	0.010342	100.0000	0.000000	0.000000	0.000000
2	0.013682	72.77623	0.134163	26.04880	1.040805
3	0.013769	72.15912	0.330073	25.89110	1.619711
4	0.013852	71.65637	0.363733	26.30258	1.677315
5	0.013856	71.61108	0.374977	26.31525	1.698693
6	0.013859	71.59109	0.376627	26.33037	1.701910
7	0.013859	71.58838	0.377111	26.33163	1.702883
8	0.013859	71.58751	0.377199	26.33224	1.703057
9	0.013859	71.58736	0.377222	26.33232	1.703103
10	0.013859	71.58732	0.377226	26.33234	1.703112

Notes: Cholesky ordering- exchange rate, oil supply shock, aggregate demand shock, oil-specific demand shock. Standard errors: Monte Carlo (10,000 repetitions).

Table 12

VARIANCE DECOMPOSITION OF INTEREST RATE

Period	*S.E.*	*Interest rate*	*Oil supply shock*	*Aggregate demand shock*	*Oil-specific demand shock*
1	1.431336	100.0000	0.000000	0.000000	0.000000
2	1.508329	90.05159	0.299485	3.741150	5.907780
3	1.515037	89.34945	0.700266	4.060880	5.889403
4	1.515576	89.28714	0.702346	4.088828	5.921682
5	1.515612	89.28372	0.702437	4.090737	5.923102
6	1.515615	89.28345	0.702443	4.090998	5.923106
7	1.515615	89.28342	0.702443	4.091028	5.923107
8	1.515615	89.28342	0.702443	4.091032	5.923108
9	1.515615	89.28342	0.702443	4.091032	5.923108
10	1.515615	89.28342	0.702443	4.091032	5.923108

Notes: Cholesky ordering- interest rate, oil supply shock, aggregate demand shock, oil-specific demand shock. Standard errors: Monte Carlo (10,000 repetitions).

Table 13

VARIANCE DECOMPOSITION OF CURRENT ACCOUNT DEFICIT

Period	*S.E.*	*Current account deficit*	*Oil supply shock*	*Aggregate demand shock*	*Oil-specific demand shock*
1	0.436645	100.0000	0.000000	0.000000	0.000000
2	0.484645	95.49759	1.885092	1.273898	1.343424
3	0.494444	94.38356	2.194169	1.514285	1.907980
4	0.496307	94.12907	2.275173	1.557974	2.037782
5	0.496634	94.08317	2.290263	1.564048	2.062518
6	0.496689	94.07540	2.293006	1.564866	2.066730
7	0.496698	94.07414	2.293472	1.564974	2.067418
8	0.496700	94.07393	2.293550	1.564989	2.067529
9	0.496700	94.07390	2.293563	1.564991	2.067547
10	0.496700	94.07389	2.293565	1.564992	2.067550

Notes: Cholesky ordering- current account deficit, oil supply shock, aggregate demand shock, oil-specific demand shock. Standard errors: Monte Carlo (10,000 repetitions).

6. CONCLUSION

Given the exposure of the Jamaican economy to oil price shocks, an analysis of the impact of these disturbances on the major macroeconomic indicators was deemed important. In addition, recognizing that increases in oil prices could stem from either demand or supply related developments, the shocks were decomposed in an effort to understand the impact of various oil shocks on the Jamaican economy.

The effects of the shocks on the macroeconomic variables of the Jamaican economy varied in accordance with the type of shock. Changes in oil prices stemming from increased global aggregate demand generally led to an improvement in domestic macroeconomic variables, particularly real GDP. However, higher oil prices emanating from a shock to global crude oil supplies or from a perceived threat to future oil supplies (leading to speculative demand) largely resulted in an overall deterioration in Jamaica's economy, contributing to an acceleration in inflation and a potentially higher current account deficit. Of note, the impact of oil price shocks on the Jamaican macroeconomy largely failed to exhibit permanent effects. This finding could be associated with the relative dependence on oil, reflected in Jamaica's fairly inelastic demand for the product. Given the conclusions, it would be useful to study the impact of price shocks to agricultural raw materials on the domestic macroeconomic variables to determine if the results hold for all imported raw materials.

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Economic Growth and Convergence in Latin America, 1950-2010

Abstract

Latin America's long-run economic growth is dealt with to prove conditional convergence in per capita GDP for two types of leading economies. Mixed empirical evidence in favor of economic convergence is found for the period 1950-1990, while conditional convergence toward both a region's average and the US economy is shown to exist in the period 1990-2010. The possibility for units to exhibit cross-section dependency in heterogeneous panels is taken into account by the second generation tests here applied.

Keywords: Econometric models, economic growth, Latin American economies, United States, per capita GDP.
JEL classification: C13, F44, C54.

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1. INTRODUCTION

Latin America's (LA) long-run economic growth can be divided into at least two clearly identifiable subperiods. The first of these corresponds to the years from 1950 to 1980, known as the golden years, when it is generally considered that LA was one of the most developed regions outside the industrial world (Elson, 2005), with economic potential very similar to that of Spain, Italy and South Korea (Barboni and Treibich, 2010). Nonetheless, everything points to the fact that this potential could not be consolidated due to political, religious and quality of human capital, factors which led to a process of divergence from the referred economies (Barboni and Treibich, 2010).

The crisis of 1981-1982 started the so-called *lost decade* in LA, characterized by slow growth. In the nineties average growth was modest, while over the following ten years (2000-2010) higher growth combined with greater variability was observed (Solimano and Soto, 2003).

For this last phase of long-run growth in LA, discussion on the processes of convergence or divergence draws different conclusions. The works of Astorga, Bergés and Fitzgerald (2005), and Astorga (2010) conclude that if the behavior of six LA countries (Argentina, Brazil, Chile, Colombia, Mexico and Venezuela) is analyzed in the last century (1900-2000), it is found that they tend toward economic and social convergence mainly due to the similarity in their patterns of industrialization, urbanization and public provision. These authors also state that the remaining countries of the region did not experience a process of convergence and that the main sources of economic growth are concentrated in the accumulation of investment and human capital. Moreover, Martín-Mayoral (2010) studies the disparities across South American, Central American (excluding Belize) and North American (excluding the United States and Canada) countries during the period 1950-2008. The results show a slow convergence process up to 1985, subsequently a process of accelerated conditional convergence with different steady states is observed, which is mainly explained by the rate of saving to investment, public spending and trade liberalization.

For a specific period, from 1980 to 2010, characterized by moments of low growth, debt crises, structural reforms, paradigm shifts and globalization, Barrientos (2007) suggests it is much more appropriate to talk of at least three groups of countries: the group of countries with *good institutions* (composed of Argentina, Brazil, Chile, Colombia, Costa Rica, Mexico and Uruguay), that suffered serious consequences of the debt crises but then tend toward higher growth rates; the *painful* group with weak institutions (composed of Bolivia, Ecuador, Guatemala, Honduras, Nicaragua, Paraguay, Peru, El Salvador and Venezuela), which exhibit bad economic and social results and, finally, the *vulnerable* (composed of the Bahamas, Barbados, Belize, Cuba, Dominica, Dominican Republic, Guyana, Haiti, Jamaica, Panama, Puerto Rico, St. Kitts and Nevis, Santa Lucia, Saint Vincent and the Grenadines, and Trinidad and Tobago). From the point of view of sigma convergence, there are no data to conclude convergence or divergence across all the countries, although for the *good institutions* group a process of convergence was found until 1990 and divergence after that year. Several countries from the *painful* group exhibit high per capita GDP dispersion levels and negative economic growth rates. The *vulnerable* group, which is more homogenous and has a very low GDP dispersion, maintained periods of convergence during 1970-1995, divergence in 1995-1999, and convergence again after that. Barrientos's (2007) results for the *good institutions* group show absolute and conditional convergence of 2% and 3.6%, respectively, for the period 1980-2010; the *painful* group exhibited absolute convergence of 0.7% and conditional one of 5.7%, while the *vulnerable* group converged in absolute terms at 6% over the same period. The conclusion is that external factors were determinant of the path of convergence among the countries in each group.

Holmes (2006), Cermeño and Llamosa (2007), Escobari (2011), Rodríguez et al. (2012) utilize the concept and methods of stochastic convergence, unit roots or cointegration to study convergence processes comparing leading economies inside and outside LA. Holmes (2006) evaluates the convergence hypothesis for eight Latin American countries in the period 1900-2003 using the Markov methodology of regime switching and defines for it the concepts of partial convergence (change from a steady regime to another non-stationary) and varied convergence (degree of persistence). By applying this methodology, he found the existence of a switching from a stationary or convergence process to another non-stationary or divergent process, which can also be identified as the existence of two different stationary regimes. Cermeño and Llamosa (2007) use the approach of Bernard and Durlauf (1995) to analyze possible convergence processes for Argentina, Brazil, Canada, Chile, USA and Mexico for the period 1950 to 2000. Neither the restricted nor the unrestricted versions (or absolute and conditional convergence, respectively) of the cointegration analysis for the comparison between LA countries and the USA show strong evidence of convergence, although in the cases comparing Argentina-USA, Chile-USA and Brazil-Argentina the results show weak evidence.

The work of Escobari (2011) for 19 countries and the period 1945 to 2000 applies unit root analysis and compares pairs of countries using the same methodology employed by Bernard and Durlauf (1995). Thus, it finds a process of convergence between the Dominican Republic and Paraguay. When groups of countries were considered it found more evidence of convergence across the economies of Central America and the Caribbean than across the economies of South America. Finally, the study of Rodríguez et al. (2012) on the hypothesis of convergence toward the economy of the USA for 17 Latin American countries during the period 1970 to 2010 using unit root tests and panel cointegration finds no evidence of absolute convergence, but does see conditional convergence.

This paper presents an analysis of the path of long-run economic growth of Latin American countries in accordance with the hypotheses of absolute and conditional convergence in per capita GDP with respect to two types of leading economies: a region's average and the USA. To test the convergence hypothesis first and second-generation cointegration and unit root panel tests were applied for the period 1950-2010. The second-generation tests, such as those of Maddala and Wu (1999) and Pesaran (2007) have the advantage of eliminating root homogeneity assumptions and independence between the cross-section units, assumptions upheld by the majority of first generation tests, e.g. those of Pesaran and Smith (1995), Pesaran (1997) and Pesaran et al. (1999). The results obtained show mixed and inconclusive evidence for economic convergence in the period 1950-1990 and of conditional convergence toward the region's average and to the USA during the 1990-2010 period of trade liberalization.

The paper is organized as follows. Section 2 broadly outlines the tests employed and presents a brief review of the empiric literature. Section 3 describes the econometric methodology employed and the data. Section 4 gives the results of the econometric tests carried out and, finally, Section 5 states the conclusions.

2. REVIEW OF THE LITERATURE

2.1 Specification of Absolute and Conditional Convergence Tests

Baumol, Nelson and Wolff (1994) make a classification of the different interpretations of convergence most used in the specialized literature: homogenous, catch-up, gap, absolute convergence, explained convergence, asymptotic convergence and limit convergence. All these interpretations can be linked to the conclusions of the neoclassical growth model for closed economies (Ramsey, 1928; Solow, 1956; Cass, 1965; and Koopmans, 1965), which predict that the growth rate trend of the capital-labor ratio (K/L) is inversely related to its initial level (Galindo and Malgesini, 1994).

In response to the many criticisms of the endogenous growth theory, Barro and Sala-i-Martin (2004), and Mankiw, Romer and Weil (1992) substituted the concept of Baumol's absolute convergence with that of conditional convergence, taking into account the international economic consistency of the nineties. The first interpretation of this concept is that the existence of convergence does not only depend on the capital-labor ratio, but also on other economic conditions (human capital, social capital, technology, policies, etc.), which can drive the process of convergence across countries. For Sala-i-Martin (1997), the conditional convergence hypothesis also allows for understanding the conditions that economies should fulfill in order to be able to group them into convergence clubs.

The convergence concept commonly employed in most studies is that of $\beta-$ convergence. It is said that there is absolute $\beta-$ convergence across countries if there is a negative relation between the growth rate of per capita income and the initial value of per capita income, which implies that the poorest countries grow at a faster rate than rich countries in such way as to arrive at the same long-run equilibrium.

In the nineties, most studies concentrated on the relation between the growth rate of income per capita and different standards of living measures in cross section to investigate the growth process. These studies were based on the following model:

$$g_i = \alpha X_i + \beta y_{i0} + \varepsilon_i , \tag{1}$$

where g_i is the country's growth rate, y_{i0} is the value of the country variable at the start of the period studied, X_i includes variables by country to control for the specific effects of each of them and ε_i is the error term. The initial value of variable y_{i0} is included in order to test the convergence hypothesis (Durlauf, 2000). Thus, if the value of β is negative in Equation 1, there is $\beta-$ convergence. In terms of Equation 1, one way of testing the absolute, or unconditional, version is by excluding each country's specific control variables, verifying that β is negative, while a conditional convergence test is carried out by including the X_i control variables (Barro and Sala-i Martin, 2004).

Different studies have criticized the application of cross-section growth models to prove absolute or conditional convergence and have proposed panel methodologies compatible with the inferences of exogenous and endogenous models (Bond et al., 2010).[1] For instance, Bernard and Durlauf (1995) state that once this analysis is applied to a group of country data through an appropriately specified model with multiple steady-states a negative β coefficient for the total sample can be attributed to a subsample of those countries that converges to the specific steady-state group. In addition, Quah (1993, 1996a, 1996b, 1997) suggests that these tests for the convergence hypothesis suffer from Galton's fallacy, i.e., once we regress growth rates to their initial levels, a negative β coefficient is due to a regression toward the mean, which does not necessarily imply convergence.

The vast majority of studies that have used Equation 1 have tended to ignore underlying patterns of heterogeneity in the data by using an identical regression model for all countries in the sample. Some of them use dummies for Latin America or sub-Saharan Africa in order to take into account the differences in the growth process for those groups of countries. However, this is not enough to capture the statistical measures of the clubs in the group of data. In this regard, Bernard and Durlauf (1994 and 1995) evaluate the possibility of convergence using the following model:

$$y_{it} = \alpha_{ij} + \beta y_{jt} + \varepsilon_{ijt}, \tag{2}$$

where y_{it} is per capita income of the country in question, y_{jt} is per capita income of the leading or reference economy and a_{ij} is a

[1] In the same way as Bond et al. (2010), in this paper we use the estimators proposed by Pesaran and Smith (1995), Pesaran (2007), and Pesaran et al. (1999). The difference between the specification of Bond et al. (2010) and ours lies in the fact that he aims to analyze how capital accumulation affects growth, he does not carry out estimates for Latin America and does not use the referred estimators to test convergence, while the specification used here is applied to the convergence test for Latin American countries.

constant that denotes permanent differences between the two economies (Cermeño and Llamosas, 2007). If convergence exists, the differences between two countries will tend to decrease over time, i.e., it requires that $\alpha_{ij} = 0$ in order for the differences to be completely eliminated (absolute convergence). If the latter is not fulfilled, it will tend toward a different determined level (conditional convergence). Thus, fulfillment of the absolute convergence hypothesis requires that $\beta = 1$ and $\alpha_{ij} = 0$. Therefore, if $\alpha_{ij} \neq 0$ there is evidence of conditional convergence.

If absolute convergence is fulfilled, a simple and direct way of proving it would be to obtain the difference between per capita income of the country in question and per capita income of the leading or reference economy, both in natural logarithms:

3 $$y_{it} - y_{jt} = \varepsilon_t .$$

Based on this series, the null hypothesis of non-convergence can be written as:

4 $$H_0 : y_{it} - y_{jt} = I(1), \quad \forall i = 1, \ldots, N.$$

The above can be carried out through unit root tests. This version of the test is known as the restricted version. According to Cheung and García (2004), testing the null hypothesis set out in Equation 4 can bias the results toward acceptance of the non-convergence hypothesis due to the reduced power of the unit root tests. Cheung and García therefore propose evaluating the convergence hypothesis in the following way:

5 $$H \ : y_{it} - y_{jt} = I(0), \quad \forall i = 1, \ldots, N.$$

If it is not possible to reject Equations 4 and 5 at the same time, the data cannot provide evidence for accepting or rejecting the convergence hypothesis.

As for the unrestricted version of the test, it is not assumed a priori and the model of Equation 2 is employed for estimating parameters α_{ij} and β. In this version of the test, the non-convergence hypothesis is evaluated by applying the unit root test to the errors estimated in this model. With this approach, the null hypothesis states that there is no cointegration between income per capita of the country studied relative to the leading economy. This version of the test also has the advantage that it is possible to determine if the constant is significant and, therefore, can show evidence of conditional convergence as well as verify whether the vector $(1,-1)$ of the restricted model is fulfilled or not.

The test defined in Equation 3 for demonstrating the convergence hypothesis between two countries can be extended for a panel model that includes a group of countries in the following way:

$$D_1 y_{it} = y_{it} - y_{lt} \ , \tag{6}$$

where y_{it} is the income per capita of country i at time t, and y_{jt} is the income per capita of the leading country at time t, both in algorithms. Thus, the convergence hypothesis between two countries can be tested through panel integration and cointegration analysis when the income per capita of both countries are not stationary (Díaz et al., 2009), which can be carried out applying different panel unit root tests to the group of series resulting from Equation 6.

A less restrictive version of Equation 6 is an extension of Equation 2 to the panel model as follows:

$$\begin{aligned} y_{it} &= \alpha_i + \beta y_{jt} + v_{it} \\ D_2 y_{it} &= y_{it} - \alpha_i - \beta y_{jt} = v_{it} \end{aligned} . \tag{7}$$

Thus, Model 7 gives an estimate of the slope parameter for the panel as a whole, which allows for testing the convergence hypothesis for the group of countries included in the panel given that, as will be shown below, according to the estimation methodology of Pesaran, Shin and Smith (1999) for cointegrated, panels it is possible to estimate parameter β for the panel as a whole and a speed of adjustment coefficient for each of the units considered. If the GDP per capita of countries included in the sample and that of the leading economy are cointegrated it will therefore also be possible to allocate a homogeneous long-run relation for the whole panel and the way in which it responds to each of the units in such relation.

2.2 Literature on Convergence

Evans (1997) demonstrates that when control variables are introduced into Equation 1, although these control 90% of the variance of steady-state GDP per capita values, the probability limit of the least squares estimator of the initial income coefficient (which is the convergence indicator) is approximately equal to half its true value. For this reason, it is not advantageous to make inferences employing this type of regressions.

Among the studies that have employed time series techniques, the following stand out: Linden (2000) studies the OECD group of countries by applying multivariate augmented Dickey-Fuller (ADF) and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) unit root tests by pairs, finding convergence only for Norway, Sweden and UK. Amable and Juillard (2000) apply the same tests for a sample of 53 countries, finding that the ADF test almost never confirmed convergence except in the cases of Denmark and Germany. Camarero, Flôres and Tamarit (2002) study countries in the Mercosur through multivariate ADF tests and panel models, finding evidence of convergence for some countries. Easterly, Fiess and Lederman (2003) analyze the convergence hypothesis between Mexico and USA with Johansen's test and find evidence of conditional convergence. Finally, Cheung and Pascual (2004) analyze the case of the Group of Seven (G7) through multivariate ADF tests and panel studies, showing evidence that the multivariate ADF test does not confirm convergence.

Cermeño and Llamosas (2007) employ the restricted and unrestricted version of Model 2 to test the convergence hypothesis for GDP per capita across six emerging countries with respect to the USA. To do this they implement the Gregory and

Hansen (1996) approach of cointegration under possible structural change. Their results suggest that in most cases there is no evidence to support convergence under structural change, and that the gaps of income per capita between the countries considered relative to the USA are consistent with a non-convergence processes.

3. ECONOMETRIC METHODOLOGY AND DATA

3.1 Panel Unit Root Tests

Panel unit root tests are similar, but not identical, to the unit root tests carried out on any series in particular. This section briefly describes the two panel unit roots tests employed in this paper.

Maddala and Wu (1999, hereafter, MW), sustain that various difficulties emerge in the Im-Pesaran-Shin (IPS) test because it relaxes the homogeneity assumption through the unit roots.[2] MW suggest using a Fisher type test, which is constructed based on a combination of p values (denoted by π_i) of the unit root test statistic in each of the cross sections. The MW test statistic, λ, is given by:

$$\lambda = -2\sum_{i=1}^{N} \ln \pi_i \,, \tag{8}$$

which is distributed as an $\chi^2(2N)$ under the null hypothesis of cross-sectional independence. In the same way, Breitung (2000) argues that IPS tests lose power by including individual trends. One of the advantages of the Maddala y Wu (1999) test is that its value does not depend on the different lags included in the individual regressions for obtaining each of the ADF statistics.

[2] The homogeneity assumption implies that all the individual roots are equal, meaning it has to be assumed $(\alpha_i = \alpha = 0, \forall i)$, while the heterogeneity assumption indicates that all the roots are different, but $(\alpha_i = 0, \forall i)$ must be fulfilled for convergence to exist.

As in the case of most ADF tests, both IPS and MW tests rest on the assumption that cross section units are independent. The second generation panel unit root test we employ in this paper is that of Pesaran (2007),[3] who proposed the CIPS test, the test statistic of which is the individual cross section mean of the t statistics of individual ordinary least square coefficients of y_{it-1} in regression CADF (cross-sectionally ADF) for each unit in the panel. The CADF regressions correspond to the ADF test that incorporates the cross-section averages of lagged levels and first-differences of the individual series. Thus, the regressions are of the following type:

$$\Delta y_{it} = \alpha_i y_{it-1} + \lambda_i \overline{y}_{t-1} + \sum_{j=1} \eta_{ij} \Delta \overline{y}_{t-j} + \sum_{j=1}^{p} \delta_{ij} \Delta y_{i,t-j} + e_{it} \quad (9)$$

In this test, the null hypothesis $(\alpha_i = 0, \forall i)$ is that all units in the panel possess a unit root, as opposed to the variance stationarity alternative where at least some of them possess one.

3.2 Kao (1999) Panel Cointegration Tests

Kao (1999) proposed ADF type tests similar to the standard single equation approach adopted in Engle and Granger's two-step procedure. In the case dealt with here, the procedure consists of estimating the following panel regression model:

$$y_{it} = \alpha_i + \delta_i z_{it} + \beta y_{lt} + \varepsilon_{it}, \quad (10)$$

where it is assumed that y_{it} and y_{lt} are non-stationary and that z_{it} is a matrix of deterministic components. The residuals of this model are used to estimate the following model:

$$\hat{e}_{it} = \rho \hat{e}_{i,t-1} + \upsilon_{it}. \quad (11)$$

[3] This test takes into account the possibility that units in the panel could be dependent.

where $\widehat{e}_{it} = \left(y_{it} - \alpha_i - \delta_i z_{it} - \beta y_{lt}\right)$. This case attempts to test the null hypothesis of non-convergence, $H_0: \rho=1$, in Equation 11, against the alternative hypothesis where y_{it} and y_{lt} are cointegrated, i.e., that $H_1: \rho<1$. Kao developed four Dickey-Fuller (DF) type tests that only limit the case of fixed effects. Two of Kao's tests assume robust exogeneity of regressors and errors in Equation 10 denoted by DF_p and DF_t, while the other tests, which are non-parametric, make corrections for any endogenous relation and are denoted by DF_p and DF_t. The four tests include non-parametric corrections for the possibility of any serial correlation given that Equation 11 involves an ordinary least squares regression (MCO) of de $\widehat{e}_{it}$ over a single lagged value of $\widehat{e}_{it}$.

As an alternative, Kao also proposed a test that extends Equation 11 by including lagged differences at residuals. He therefore obtains an ADF version of his test on the existence of serial correlation as part of the regression procedure. All the tests are asymptotically distributed in accordance with standard normal distribution. It is important to point out that the versions of Kao's test impose homogeneity in the β slope coefficient, i.e., it is not allowed to vary across the individuals making up the panel.

3.3 Panel Estimation Methods for Cointegrated Variables

For panel cointegration models the asymptotic properties of regression model coefficient estimators and associated statistical tests are different from those estimated by cointegrated time series models (Baltagi, 2008).

Some of these differences have been revealed in recent works by Kao and Chiang (2000), Phillips and Moon (1999), Pedroni (1999, 2000, 2004), and Mark and Sul (2003). Panel cointegration models are designed for studying long-term relations typically found in macroeconomic and financial data. Such long-term relations are often cited by economic and financial theory, which is the main reason for estimating regression coefficients and testing whether or not they satisfy theoretical restrictions. Phillips and Moon (1999) and Pedroni (2000) propose a fully-modified (FM) estimator, which can be viewed as a generalization of the Phillips y Hansen (1990) estimator, while Kao and Chiang (2000) advance an alternative method based on the dynamic least squares estimator, taking the works of Saikkonen (2001) and Stock and Watson (1993) as a reference.

3.3.1 Group Mean Estimator

To test the convergence hypothesis for Latin American countries we employ the estimators proposed by Pesaran, Shin and Smith (1999), who suggest two different estimators in order to resolve the possible lag attributable to slope heterogeneity in dynamic panels. These are the mean group (MG) and pooled mean group (PMG) estimators.

The MG estimator allows long-term parameters to be obtained for the panel from an average of the long-term parameters in autoregressive distributed lag (ADRL) models for units or individuals (Asteriou and Hall, 2007). For instance, if the ADRL is as follows:

12 $$y_{i,t} = a_i + \gamma_i y_{i,t-1} + \beta_i x_{i,t} + e_{i,t}\,.$$

Therefore, the long-term parameter, θ_i, for the individual or unit i is:

13 $$\theta_i = \frac{\beta_i}{1-\gamma_i}\,.$$

The estimators for the whole panel would therefore be given by:

14 $$\hat{\theta} = \frac{1}{N}\sum_{i=1}^{N}\theta_i\,,\ \ \hat{a} = \frac{1}{N}\sum_{i=1}^{N}a_i. \quad .$$

It is possible to show how with a sufficiently large number of lags the MG estimator provides super consistent estimators for the long-term parameters even when the order of integration of the regressors is equal to one (Pesaran, Shin and Smith, 1999).

The MG estimators are consistent and have normal asymptotic distributions for sufficiently large N and T. Nevertheless, for samples where T is small, the MG estimator is lagged and can lead to erroneous inferences, meaning it should be used with caution in such cases.

3.3.2 Pooled Mean Group Estimator

Pesaran and Smith (1995) show that, unlike static models, pooled heterogeneous dynamic panels generate estimators that are inconsistent even in large samples. Baltagi and Griffin (1997) argue that the benefit in terms of efficient data aggregation outweighs the loss caused by the bias induced by heterogeneity. Pesaran and Smith (1995) observe how it is improbable that dynamic specification is common to all units, while it is at least conceivable that long-run parameters of the model may be common. They propose carrying out the estimate by averaging the estimated parameters individually or pooling the long-term parameters where the data allows it, and estimating the model as a system. Pesaran, Shin and Smith (1999) refer to this method as the pooled mean group (PMG) estimator, which combines the efficiency of the pooled estimate while avoiding the problems of inconsistency arising from pooling dynamic heterogeneous relations.

The PMG sits in between the MG (where both slopes and intercepts are allowed to vary across units) and the classic fixed effects model (where slopes are fixed and intercepts vary across units). Calculation of the PMG estimator only restricts long-term coefficients to be the same across units, while allowing short-term coefficients to vary across them.

More precisely, the unrestricted specification of the ADRL system of equations is as follows:

$$y_{it} = \mu_i + \sum_{j=1}^{p} \lambda_{ij} y_{i,t-1} + \sum_{j=0}^{p} \lambda \delta'_{ij} x_{i,t-j} + \varepsilon_{it} , \quad (15)$$

where $x_{i,t\text{-}j}$ is a vector of explanatory variables and μ_i represents the fixed effects. In principle the panel can be unbalanced and p and q may vary across units. This model can be reparametrized as a vector error correction model (VECM):

$$\Delta y_{it} = \theta_i \left(y_{i,t-1} - \beta' x_{i,t} \right) + \sum_{j=1}^{p-1} \gamma_{ij} \Delta y_{i,t-j} + \sum_{j=1}^{q-1} \phi'_{ij} \Delta x_{i,t-j} + \varepsilon_{it}, \quad (16)$$

where the θ_i are short-term parameters for each of the units, and β is the short-term parameter common to all of them. The estimate can be carried out by MCO, imposing and testing cross section restrictions on β. Nevertheless, this procedure could be inefficient as it ignores contemporary residual covariance. Given the latter, an estimator could be calculated with Zellner's SUR method, which is a type of feasible generalized least squares estimation. However, the SUR procedure is only possible if $N < T$, the reason why Pesaran, Shin and Smith (1999) suggest employing the maximum likelihood method.

4. RESULTS

First, we look into the possible presence of unit root in the difference between each country's income per capita relative to each of the two indicators considered as *leading economy:* GDP per capita of the USA and average GDP per capita of the region. The latter calculation includes GDP per capita of the USA. To this end, we apply the tests of Maddala and Wu (1999) and of Pesaran (2007), with different lags to Dy_{it}, as established in Equation 6. Tables 1 and 2 show the results of the root test for different periods and the sample as a whole.

In the case of the difference between the GDP per capita of each country as compared to that of USA, the MW and Pesaran tests carried out with and without trend (see Table 1) show that for both the total sample and the first subperiod it is not possible to reject the unit root null hypothesis in any case in the panel considered, meaning that in these cases there are

Table 1

PANEL UNIT ROOT TESTS FOR $D_l y_{it} = y_{it} - y_{lt}$, RELATIVE TO THE USA: TOTAL SAMPLE AND BY PERIODS

	Maddala and Wu (1999)				***Pesaran (2007)***			
	Without trend		*With trend*		*Without trend*		*With trend*	
Lags	χ^2	*p value*	χ^2	*p value*	χ^2	*p value*	χ^2	*p value*
				Total sample (1951-2010)				
0	19.85	[0.97]	10.87	[1.00]	0.27	[0.61]	1.61	[0.95]
1	22.96	[0.92]	18.27	[0.99]	0.04	[0.52]	0.48	[0.68]
2	25.15	[0.86]	23.72	[0.91]	0.49	[0.69]	0.78	[0.78]
3	32.02	[0.57]	19.90	[0.97]	0.74	[0.77]	0.99	[0.84]
4	28.77	[0.72]	24.60	[0.88]	1.67	[0.95]	2.22	[0.99]
				First period (1951-1990)				
0	13.56	[0.99]	10.86	[1.00]	2.16	[0.99]	1.31	[0.91]
1	16.03	[0.99]	17.91	[0.99]	2.22	[0.99]	0.71	[0.76]
2	22.32	[0.94]	14.44	[0.99]	2.90	[0.99]	1.84	[0.97]
3	22.10	[0.94]	23.75	[0.91]	3.57	[1.00]	2.01	[0.98]
4	22.25	[0.94]	20.57	[0.97]	4.23	[1.00]	3.16	[0.99]
				Second period (1990-2010)				
0	35.42	[0.40]	7.64	[1.00]	−1.34	[0.09]	−0.29	[0.39]
1	39.77	[0.23]	11.30	[1.00]	−2.58	[0.01]	−3.43	[0.00]
2	42.09	[0.16]	11.05	[1.00]	−1.78	[0.04]	−5.07	[0.00]
3	37.60	[0.31]	8.45	[1.00]	−1.33	[0.09]	−4.70	[0.00]
4	52.12	[0.02]	17.82	[0.99]	−0.93	[0.18]	−2.30	[0.01]

Note: Numbers in parenthesis are p values for the lags included in each test.
Source: Own elaboration.

no signs of convergence with respect to this indicator in the periods analyzed. For the second subperiod the MW test without trend and with four lags, and Pesaran's test without trend, one and two lags and trend for lags one to four, reject the unit root null hypothesis. This suggests some indications of stationarity in the difference between the GDP per capita of each

Table 2

PANEL UNIT ROOT TESTS FOR $D_l y_{it} = y_{it} - y_{lt}$ RELATIVE TO REGION'S AVERAGE: TOTAL SAMPLE AND BY PERIODS

	Maddala and Wu (1999)				*Pesaran (2007)*			
	Without trend		*With trend*		*Without trend*		*With trend*	
Lags	χ^2	*p value*	χ^2	*p value*	χ^2	*p value*	χ^2	*p value*
				Total sample (1951-2010)				
0	44.54	[0.16]	31.18	[0.70]	0.59	[0.72]	1.03	[0.85]
1	39.95	[0.30]	31.28	[0.69]	0.23	[0.59]	−0.23	[0.41]
2	31.34	[0.69]	22.08	[0.97]	1.02	[0.85]	0.40	[0.66]
3	28.82	[0.80]	17.95	[0.99]	1.52	[0.94]	0.43	[0.67]
4	26.40	[0.88]	17.32	[0.99]	2.23	[0.99]	1.40	[0.92]
				First period (1951-1990)				
0	35.44	[0.50]	30.15	[0.74]	3.03	[0.99]	2.14	[0.98]
1	28.07	[0.83]	33.82	[0.57]	3.04	[0.99]	1.49	[0.93]
2	25.91	[0.89]	23.16	[0.95]	4.03	[1.00]	2.58	[0.99]
3	19.71	[0.99]	22.19	[0.97]	4.89	[1.00]	2.97	[0.99]
4	15.33	[0.99]	26.87	[0.87]	5.42	[1.00]	4.33	[1.00]
				Second period (1990-2010)				
0	64.06	[0.00]	24.23	[0.93]	−0.19	[0.43]	−0.85	[0.20]
1	56.96	[0.02]	42.55	[0.21]	−1.10	[0.14]	−3.36	[0.00]
2	42.07	[0.23]	29.58	[0.77]	0.10	[0.54]	−1.47	[0.07]
3	44.47	[0.16]	27.99	[0.83]	−0.43	[0.33]	−0.17	[0.43]
4	45.43	[0.14]	36.20	[0.46]	−1.49	[0.07]	0.57	[0.72]

Note: Numbers in parenthesis are *p* values for the lags included in each test.
Source: Own elaboration.

Latin American country and that of the USA and, therefore, of convergence between both indicators for the subperiod corresponding to trade liberalization.

With respect to MW and Pesaran tests, with and without trend, on the differences between each country's GDP and average GDP per capita for the region, they show a similar result for the total sample and for the first subperiod given that it is not possible in any case to reject the unit root null hypothesis in the panel for that variable (see Table 2).

For the second subperiod the MW test without trend, without lags and with one lag, and the Pesaran test with trend, with one and two lags, allow for rejecting the unit root hypothesis, which suggests the presence of some indications of stationarity in the difference of each Latin American country's GDP per capita relative to the region's average and, therefore, of convergence between both indicators for the second subperiod 1990-2010. The same can be said for the tests implemented with the difference between GDP per capita of countries in the region and that of USA.

Thus, both indicators constructed for proving the restricted version of the test show evidence that there are indications of stationarity in said indicators only during the second subperiod. This implies that the process of convergence between the Latin American countries and the USA, and the region's average was only seen in the second subperiod corresponding to the phase of trade liberalization.

Once the possible presence of convergence was verified in the total sample and the subperiods according to the restricted version of the test, we applied panel unit root tests in order to examine the possible presence of unit root in the natural logarithm of GDP per capita for countries of the region. And, if it exists, proceed to carry out panel cointegration tests of this indicator with respect to per capita GDP of the USA and average GDP per capita of the region. The results of the panel unit root tests applied to the natural logarithm of GDP per capita of the countries of the region considered are shown in Table 3.

As can be seen in Table 3, MW unit root tests do not allow for rejecting the unit root null hypothesis in the natural logarithm of GDP per capita of any of the countries considered. However, Pesaran's test in some cases shows that said hypothesis is rejected, mainly for the total sample and the first subperiod, when the test is specified with few lags. Meanwhile, in the majority of cases, Pesaran's test with trend cannot reject the unit root hypotheses for this variable. Notwithstanding the aforementioned, in the following analysis we assume that per capita GDP

Table 3

PANEL UNIT ROOT TESTS OF MADDALA AND WU (1999), AND PESARAN (2007) FOR YIT

TOTAL SAMPLE AND BY PERIODS

	Maddala and Wu (1999)				*Pesaran (2007)*			
	Without trend		*With trend*		*Without trend*		*With trend*	
Lags	χ^2	*p value*	χ^2	*p value*	χ^2	*p value*	χ^2	*p value*
Total sample (1951-2010)								
0	21.54	[0.95]	12.52	[1.00]	−1.77	[0.04]	−0.50	[0.31]
1	16.12	[0.99]	17.16	[0.99]	−2.33	[0.01]	−1.31	[0.10]
2	18.43	[0.99]	15.95	[0.99]	−1.83	[0.03]	−0.83	[0.20]
3	19.81	[0.98]	17.41	[0.99]	−1.45	[0.07]	−0.14	[0.44]
4	19.55	[0.98]	17.78	[0.99]	−0.32	[0.37]	1.45	[0.93]
First period (1951-1990)								
0	31.04	[0.61]	10.72	[1.00]	−2.29	[0.01]	−0.43	[0.34]
1	27.17	[0.79]	16.00	[0.99]	−2.49	[0.01]	−0.97	[0.17]
2	24.41	[0.89]	15.24	[0.99]	−2.13	[0.02]	−0.47	[0.32]
3	29.01	[0.71]	18.71	[0.98]	−1.08	[0.14]	0.85	[0.80]
4	26.70	[0.81]	16.15	[0.99]	−0.08	[0.47]	2.03	[0.98]
Second period (1990-2010)								
0	16.57	[0.99]	24.73	[0.88]	−0.33	[0.37]	−0.18	[0.43]
1	16.88	[0.99]	35.51	[0.40]	−2.28	[0.01]	−3.32	[0.00]
2	8.30	[1.00]	37.72	[0.30]	−0.80	[0.21]	−2.12	[0.02]
3	8.42	[1.00]	25.15	[0.87]	−1.17	[0.12]	−1.50	[0.07]
4	9.96	[1.00]	34.65	[0.44]	−0.66	[0.25]	−0.43	[0.33]

Notes: Numbers in parenthesis are *p* values for the lags included in each test.
Source: Own elaboration.

Table 4

PANEL COINTEGRATION TESTS OF KAO (1999)
TOTAL SAMPLE AND SUBPERIODS
H0: NO COINTEGRATION

Total sample (1951-2010)		*First period (1951-1990)*		*Second period (1990-2010)*	
Statistics	*Prob.*	*Statistics*	*Prob.*	*Statistics*	*Prob.*
		Relative to USA			
−1.24	[0.11]	−0.29	[0.39]	−2.15	[0.02]
		Relative to region's average			
−1.30	[0.10]	−0.63	[0.27]	−3.42	[0.00]

Note: Testing conducted by incorporating individual intercepts.
Source: Own elaboration.

of the Latin American countries considered has an order of integration equal to 1.

The results of the panel cointegration test of Kao (1999) for GDP per capita of Latin American countries and that of the USA and a region's average, both in turn considered as *leading economy*, are presented in Table 4. As can be seen, evidence of cointegration between the two indicators only exists in the second subperiod, given that for both the total sample as well as for the first subperiod it is not possible to reject the null hypotheses of non-cointegration between GDP per capita of Latin American countries and GDP per capita of the *leading economy*.

Taking into account these results, we estimate the β convergence coefficient of the restricted version of the test between GDP per capita of countries of the region and the leading economy. The results are shown in Table 5.

Table 5

RESULTS OF PMG, MG AND DFE ESTIMATORS FROM PESARAN, SHIN AND SMITH (1999)

	Total sample (1951-2010)		*First period (1951-1990)*		*Second period (1990-2010)*	
			Relative to USA			
$\hat{\beta}_{PMG}$	0.66	[0.00]	0.10	[0.32]	0.90	[0.00]
$H_0: \hat{\beta}_{PMG} = 1$	29.07	[0.00]	85.89	[0.00]	1.17	[0.28]
$\hat{\beta}_{MG}$	0.61	[0.00]	0.21	[0.62]	1.00	[0.08]
$H_0: \hat{\beta}_{MG} = 1$	8.32	[0.00]	3.59	[0.06]	0.00	[0.99]
$\hat{\beta}_{DFE}$	0.60	[0.00]	0.14	[0.57]	1.24	[0.00]
$H_0: \hat{\beta}_{DFE} = 1$	10.52	[0.00]	11.71	[0.00]	0.92	[0.34]
Hausman tests						
PMG vs MG	0.17	[0.68]	0.06	[0.81]	0.03	[0.86]
MG vs DFE	0.00	[0.99]	0.00	[0.99]	0.00	[0.99]
			Relative to region's average			
$\hat{\beta}_{PMG}$	0.97	[0.00]	0.74	[0.00]	0.83	[0.00]
$H_0: \hat{\beta}_{PMG} = 1$	0.42	[0.52]	30.43	[0.00]	27.31	[0.00]
$\hat{\beta}_{MG}$	0.94	[0.00]	0.81	[0.00]	1.26	[0.00]
$H_0: \hat{\beta}_{MG} = 1$	0.13	[0.72]	0.42	[0.52]	1.21	[0.27]
$\hat{\beta}_{DFE}$	0.92	[0.00]	0.89	[0.00]	1.14	[0.00]
$H_0: \hat{\beta}_{DFE} = 1$	0.30	[0.58]	0.45	[0.50]	1.42	[0.23]
Hausman tests						
PMG vs MG	0.04	[0.84]	0.06	[0.81]	2.89	[0.09]
MG vs DFE	0.00	[0.99]	0.00	[0.99]	0.00	[0.99]

Note: Numbers in parenthesis are p values.

The estimates show that when GDP per capita of the USA is taken as leading economy, the coefficients estimated with PMG, MG and DFE estimators for the total sample and the first subperiod were much less than one. Besides the fact that in all cases the Hausman tests show that from the PMG and MG estimators, the PMG estimator is the most efficient with the null hypothesis and both cases reject the null hypothesis that the true parameter is equal to one. On the other hand, the results are very different for the second subperiod of the sample where PMG and MG estimators were equal to 0.90 and 1.00, respectively. The first of these is significant at 1% and the second at 10%. Furthermore, it was not possible to reject the hypothesis that the $\widehat{\beta}$ coefficient is equal to one for either indicator. Thus, the presence of panel cointegration between both variables according to the test of Kao (1999) for the second subperiod, and the fact that it is not possible to reject the hypothesis that the parameter estimated by PMG is equal to one (PMG = 1) for that subgroup shows robust evidence for convergence of Latin American countries toward the USA in the second subperiod. This result is also compatible with that found with the restricted version of the test.

On the other hand, estimates carried out to test for β convergence taking a region's average as leading economy revealed that the PMG estimator of β is very close to one for the whole sample. It is not possible in this case to reject the null hypothesis that said parameter is equal to 1 either. In the same way as in the previous case where GDP per capita of the USA is taken as the leading economy, when the region's average per capita GDP is taken as leading economy Hausman's tests show that in every case the PMG estimator is more efficient than the MG estimator.

As for estimates carried out by subperiods, taking a region's average as leading economy, although all the indicators were statistically significant, in the case of estimates through the PMG it was not possible to accept the null hypothesis that this coefficient is equal to one. For this reason, we do not find evidence of convergence toward the region's average by subperiods despite the fact that we do find evidence of this for the period as a whole.

Table 6

SPEED OF ADJUSTMENT COEFFICIENTS (INDIVIDUAL AND PANEL) ESTIMATED THROUGH THE PMG ESTIMATOR OF IM, PESARAN AND SHIN (1999) RELATIVE TO USA

	Total sample			*First period*			*Second period*		
	$\hat{\theta}_i$	*Std. error*	*z*	$\hat{\theta}_i$	*Std. error*	*z*	$\hat{\theta}_i$	*Std. error*	*z*
Argentina	−0.08	0.06	−1.46	−0.08	0.06	−1.48	−0.03	0.11	−0.28
Bolivia	−0.06	0.03	−2.47	−0.21	0.09	−2.43	−0.01	0.06	−0.23
Brazil	−0.05	0.02	−2.79	−0.03	0.02	−2.17	−0.08	0.09	−0.84
Chile	0.01	0.03	0.45	−0.10	0.08	−1.24	−0.12	0.04	−2.81
Colombia	−0.05	0.03	−1.49	−0.01	0.02	−0.71	−0.06	0.08	−0.75
Costa Rica	−0.12	0.04	−2.75	−0.09	0.03	−3.24	0.03	0.10	0.34
Ecuador	−0.07	0.04	−1.89	−0.03	0.02	−1.45	−0.11	0.09	−1.20
El Salvador	−0.03	0.03	−1.00	−0.09	0.04	−2.36	−0.23	0.08	−2.97
Guatemala	−0.03	0.03	−1.11	−0.03	0.02	−1.25	−0.09	0.06	−1.56
Honduras	−0.09	0.05	−1.94	−0.05	0.05	−0.90	−0.19	0.09	−2.18
Mexico	−0.08	0.03	−2.25	−0.03	0.02	−1.57	−0.37	0.19	−1.94
Nicaragua	−0.01	0.03	−0.25	−0.15	0.10	−1.59	−0.37	0.08	−4.80
Panama	−0.02	0.02	−0.84	−0.03	0.02	−1.21	0.00	0.08	−0.01
Paraguay	−0.04	0.04	−1.20	0.01	0.02	0.32	−0.11	0.07	−1.68
Peru	−0.04	0.04	−0.98	−0.08	0.05	−1.58	0.09	0.08	1.10
Uruguay	−0.12	0.05	−2.18	−0.16	0.08	−1.94	−0.05	0.11	−0.41
Venezuela	−0.02	0.04	−0.54	−0.11	0.05	−2.14	−0.15	0.11	−1.30

Source: Own elaboration.

Table 7

SPEED OF ADJUSTMENT COEFFICIENTS (INDIVIDUAL AND PANEL) ESTIMATED THROUGH THE PMG ESTIMATOR OF IM, PESARAN AND SHIN (1999) RELATIVE TO REGION'S AVERAGE

	Total sample			*First period*			*Second period*	
	$\hat{\theta}_i$	*Std. error*	*z*	$\hat{\theta}_i$	*Std. error*	*z*	$\hat{\theta}_i$	*Std. error*
Argentina	−0.12	0.05	−2.30	−0.12	−2.30	−0.06	0.11	−0.54
Bolivia	−0.07	0.02	−2.96	−0.07	−2.96	−0.38	0.15	−2.57
Brazil	−0.05	0.02	−2.92	−0.05	−2.92	−0.45	0.17	−2.71
Chile	−0.01	0.03	−0.18	−0.01	−0.18	−0.12	0.03	−4.71
Colombia	−0.07	0.04	−1.73	−0.07	−1.73	−0.16	0.13	−1.24
Costa Rica	−0.30	0.07	−4.52	−0.30	−4.52	−0.04	0.09	−0.44
Ecuador	−0.08	0.05	−1.74	−0.08	−1.74	−0.21	0.15	−1.43
El Salvador	−0.03	0.03	−0.91	−0.03	−0.91	−0.20	0.06	−3.14
Guatemala	−0.05	0.04	−1.24	−0.05	−1.24	−0.32	0.12	−2.70
Honduras	−0.11	0.05	−2.37	−0.11	−2.37	−0.29	0.10	−2.99
Mexico	−0.09	0.04	−2.02	−0.09	−2.02	−0.30	0.17	−1.72
Nicaragua	−0.01	0.03	−0.26	−0.01	−0.26	−0.43	0.08	−5.45
Panama	−0.02	0.02	−0.66	−0.02	−0.66	0.10	0.08	1.23
Paraguay	−0.08	0.04	−2.06	−0.08	−2.06	−0.19	0.07	−2.71
Peru	−0.05	0.04	−1.30	−0.05	−1.30	0.08	0.08	1.05
Uruguay	−0.08	0.04	−1.90	−0.08	−1.90	−0.07	0.11	−0.66
Venezuela	0.00	0.03	−0.02	0.00	−0.02	−0.11	0.12	−0.87
United States	−0.02	0.03	−0.93	−0.02	−0.93	−0.04	0.09	−0.41

Source: Own elaboration.

Tables 6 and 7 show adjustment speed coefficients estimated through the PMG estimator, taking GDP per capita of the USA and a region's average as leading economy, respectively.

As can be seen in the tables above, most of the adjustment speed coefficients estimated for the whole period and for the subperiods considered are negative. This tends to corroborate the presence of a long-run steady-state relation between the variables analyzed, despite the fact that some individual adjustment coefficients were not significant.

Thus, through the PMG estimator we find evidence of Latin American convergence toward the USA only for the second subperiod, between 1990 and 2010. On the other hand, through the same estimators we find evidence of convergence toward a region's average only for the total sample; yet, paradoxically, we do not find evidence of convergence toward this indicator when the analysis is carried out by subperiods.

5. CONCLUSIONS

In this paper, we review the convergence hypothesis for individual Latin American countries relative to two references considered as *leading economies,* GDP per capita of the USA and a region's average. In order to prove the convergence hypothesis in Latin American countries relative to the leading economy, we employ restricted and unrestricted versions of the test for the whole period analyzed, 1951-2010, and for two subperiods: the first from 1951 to 1990 and the second from 1990 to 2010. The aim of this was to identify, for the total sample and the subperiods, whether there was a process of convergence toward the leading economy before and after the process of trade liberalization registered in most countries of the region.

With respect to the unrestricted version of the test, MW (1999) and Pesaran (2007) tests carried out with and without trend show that for the total sample and the first subperiod it is not possible to reject the unit root null hypothesis for any case in the panel considered when GDP per capita of the USA is assumed as leading economy. Meanwhile, for the second subperiod in some cases MW and Pesaran tests reject the unit root hypothesis, in this way giving some indications of stationarity in the difference of GDP per capita of each Latin American country relative to that of the USA and, therefore, of convergence between both indicators for the period of trade liberalization. In addition, MW and Pesaran tests applied to the restricted version of the test taking average GDP per capita of the region as leading economy, provide a similar result for both the total sample and for the first subperiod given that it is not possible to reject the unit root null hypothesis of this variable for any case in the panel. And for the second subperiod, both the MW test and that of Pesaran, in some cases, allow for rejecting the unit root hypothesis. These results suggest the presence of some indications of *stationarity* in the difference between GDP per capita of each Latin American country and the indicators considered as *leading economies* and, therefore, of convergence toward both indicators for the second subperiod (1990-2010).

The panel cointegration tests employed for proving cointegration between GDP per capita of Latin American countries and indicators for the *leading economy* show evidence of cointegration across such variables in both cases only for the second subperiod.

Finally, the results found through PMG, MG and DFE estimators applied to the unrestricted version of the test showed that when GDP per capita of the USA is considered as leading economy, the convergence hypothesis is only fulfilled during the second subperiod, which is in line with the results of the restricted test applied to the same indicator. Nonetheless, the estimations carried out to test β convergence taking a region's average as leading economy revealed that the PMG estimator of β had a value very close to one only for the total sample, as well as the fact that it is not possible to reject the null hypothesis that this parameter is equal to one. Estimations for this indicator therefore suggest a process of convergence toward the regional average.

However, these results are not consistent with those found with the restricted version of the test. In general, they are very consistent with those obtained in the works of Rodríguez et al. (2012), Martín-Mayoral (2010) and Barrientos (2007).

Thus, we have found conclusive evidence for the convergence of Latin American countries toward the USA with both tests, restricted and unrestricted, *only* for the second subperiod, where trade liberalization and globalization appear to have had a positive impact. It is important to point out that despite the fact that this empiric evidence provides some support to the version of absolute convergence for countries of the region toward the leading economy of the USA, for the second period of the sample we cannot say there is evidence of absolute convergence given that it is necessary to prove that intercept α_i of Equation 7, homogenous and heterogeneous, as the case may be, is equal to 0, which as far as we know is not possible with the econometric methodology employed here. We therefore confine ourselves to reporting that we found evidence of convergence toward the USA in the second period analyzed.

We also find mixed evidence of convergence toward a region's average for the total sample and for the second subperiod, given that in this case the restricted version tests suggest the presence of convergence in the second subperiod, while the PMG estimator denotes evidence of convergence only for the sample. We therefore believe more research is required in this area using different techniques –linear or non-linear– that help to explain the reasons behind such results.

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Policy Implications for the Application of Countercyclical Capital Buffers When the Government Borrowing Crowds Out Private Sector Credit: The Case of Jamaica

Abstract

This paper investigates the use of conditioning variables in guiding the accumulation and release phases of a capital buffer requirement for Jamaican banks. An important innovation of this study is the inclusion of public sector conditioning variables to explore the role of sovereign risk build-up in designing countercyclical buffers.

Keywords: Countercyclical capital buffers, financial stability, procyclicality, early warning indicators, sovereign risk.

JEL classification: E44, E61, G21.

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1. INTRODUCTION

One of the issues that took center stage in the international debate on the lessons of the global financial crisis of 2008-2009 is that of managing procyclicality of the financial system. Procyclicality of the financial system is defined as the amplification of the cyclical fluctuations of the economy by financial sector activities, most notably bank lending (see, for example, Bernanke et al., 1995; Borio at al., 2001; Geršl and Jakubik, 2006). This behavior can have particularly serious implications in an economic downturn as it can considerably prolong and deepen the recession via a feedback effect on the economy.

Countercyclical policy tools have recently been utilized by central banks to mitigate the negative effects of procyclicality of the banking sector. The proximate objective of a countercyclical capital requirement is to encourage banks to build up buffers in good times that can be drawn down in bad times. Buffers in this context comprise Tier 1 capital in excess of the prudential minimum, so that additional capital is available to absorb losses in the event of a boom-and-bust financial cycle. One of the main issues involved in the policy design process is the choice of conditioning variables that can guide the buildup of the buffer during the periods of expansion. Of equal significance is the identification of variables which point to releasing the capital buffer at the beginning of the bust stage.

This paper examines a range of potential early warning indicators or conditioning variables which may be used by policymakers for setting appropriate time-varying capital requirements to address banking sector procyclicality. Specifically, one aim of this study is to assess the ability of specific macroeconomic and commercial bank-level (conditioning) variables, similar to those explored in Drehman et al. (2011), in reflecting risk buildup in the banking system in Jamaica. The key finding from Drehman et al. (2011) is that the ratio of credit-to-GDP and its long-term trend (the credit-to-GDP gap) performs best as an indicator for the build-up phase of a financial boom-and-bust cycle. The authors exclude public sector debt as its tendency to be counter-cyclical reduced the performance of credit related variables in their sample.

In Jamaica's case, the fact that its banking sector has historically operated within an environment of strong fiscal dominance, which led to public sector crowding out of private sector credit, the role of sovereign risk build-up could be important in designing domestic countercyclical buffers. That is, public sector credit and public sector debt holdings could rise in booms and slowdown in the downswing. Fiscal dominance has been manifested in sustained high interest rates in the context of persistent budget deficits. For last two decades, Jamaica has been caught in a vicious cycle of very low private sector credit and unsustainable public sector debt dynamics. Consistent with the running of persistent budget deficits, along with the price incentive of a high sovereign risk premium, the growing stock of public sector debt has been supported by the oversupply of financing by the banking sector. Over this period, the stock of public sector debt (private sector credit) has remained high (low) by international standards at above 100% of GDP (at around 20% to 30% of GDP). Hence, an important innovation of this study is to include indicators capturing the level of public sector credit and investments in public sector bonds by commercial banks as candidate conditioning variables to explore the role of sovereign risk build-up in designing countercyclical buffers.

Similar to the cyclical experience with private sector credit, sovereign risk is likely underestimated by the banking sector in credit cycle upturns and overestimated in downturns. In an upturn, normally associated with higher public revenues, banks would rapidly expand holdings of public sector credit and bonds, contributing to overpriced public sector bonds and lending spreads along with inadequate bank capital buffers. In the downswing, when sovereign risk increases as public revenues decline, the opposite would tend to occur as banks become overly risk averse. In the context of this paper, the positive correlation between the financial cycle upturn and the accumulation of public sector credit and debt holdings is expected to be stronger in countries such as Jamaica which has historically exhibited high sovereign risk premium relative to the private sector interest rates (that is, *crowding out*).

Against this backdrop, the set of conditioning variables considered in the paper have been tailored to the Jamaican historical environment of strong fiscal dominance and high levels of sovereign debt, in addition to the typical private sector credit variables. These variables are evaluated using both signal extraction and receiver operating characteristics (ROC) methods to determine how effective their deviations from long-term trends (gaps) were in signaling buffer accumulation and release phases around financial crisis episodes. The main conclusion derived from the analysis is that the credit (public and private)-to real GDP gap, investment (in public sector bonds)-to-real GDP gap, private sector credit-to-real GDP gap and public sector credit-to-real GDP gap, all indicate significant signaling value for the accumulation phase. In addition, non-performing loan growth gap and provision for loan loss growth gap reveal significant predictive power for the release phase. However, similar to the finding of Drehman et al. (2011), the overall results of this study do not support the use of any fail-safe conditioning variables to guide policy. Rather, the combination of a set of conditioning variables and judgment is advisable in designing a policy framework for dampening procyclicality.

The paper is organized as follows. In the next section, the data used in the analysis is defined. Sections 3 and 4 compares the performance of different conditioning variables around crisis episodes by using the signals approach and describes the evaluation of these variables using ROC curve analysis, respectively. Section 5 presents the empirical results from the signal extraction method and the ROC curve analysis. The final section concludes and provides some policy implications.

2. DATA DESCRIPTION, INDICATOR MEASUREMENT AND THRESHOLD CHOICE

The period for assessment of the historical performance of conditioning (indicator) variables for application of a countercyclical capital buffer to Jamaica's commercial banking sector covers 1990 to 2012. The data set, which was provided by the Central Bank, was unavailable prior to 1990. In the context of this paper, a crisis episode is defined as the occurrence of a threat to overall stability of banking system characterized by: *1)* significant NPLs, consistent with the effects of procyclicality in the down cycle; and *2)* illiquidity, requiring emergency lending assistance (ELA) by the Central Bank, consistent with financial instability. The data set is suitably long as it covers periods of extensive bank vulnerability as well as credit upswing periods[1]. There are two banking crisis episodes identified within the sample period. Accordingly, the conditioning variables are juxtaposed against a banking crisis indicator variable to assess their signaling ability.

The first crisis episode spans the six-quarter period September 1997 to December 1998, which began with successive runs on two commercial banks affiliated with life insurance companies in December 1996 and February 1997. Due to the close relationship between insurance companies and commercial banks, liquidity and insolvency problems that originated in the insurance sector spread to the banking sector. Severe liquidity shortfalls resulted in the Central Bank providing ELA to four commercial banks. In addition, the Government of Jamaica (GOJ) established the Financial Sector Adjustment Company (Finsac) in January 1997 to resolve the serious problems faced by the financial sector. During 1997, the nonperforming loan (NPL) ratio for the commercial bank sector doubled to 28.9%

[1] Similar studies in the literature, which involve the ranking of indicators, have also been constrained in coverage of banking crises. For example, Giese et al. (2012) assess indicators in the UK context using data covering three past episodes of banking system distress. The authors aptly note, however, that their rankings should be treated with appropriate caution.

by the end of the year. The increase in the NPL ratio followed on above-normal expansion in private sector credit growth of 68.9% in 1993 which subsequently slowed to 25.3% by 1996 and –33.5 in 1997. By end-1998, Finsac had intervened in the operations of most of the domestic commercial banks, over half of the life insurance companies as well as a few merchant banks and building societies.

The second crisis episode began in the September 2008 quarter and also spans six quarters. In October 2008, as a direct consequence a slowdown in lending as well as economic activity triggered by the global financial turmoil and to preserve overall financial stability, the Central Bank offered an emergency temporary lending facility in United States dollars to domestic financial institutions. This facility was primarily intended to provide liquidity to these institutions due to contagion which resulted in a dysfunctional interbank money market as well as large margin calls and cancelled repurchase agreements on GOJ global bonds held with overseas institutions. The stated objectives of the temporary lending facility were to *a)* alleviate significant short-term US dollar liquidity needs of domestic financial institutions, *b)* stabilize GOJ global bond prices which had sharply declined, and *c)* minimize volatility pressures in the domestic foreign exchange market. In addition, the Central bank established a special intermediation facility in the final quarter of 2008 to facilitate the flow of credit among local financial institutions. This facility gave extraordinary access to domestic liquidity to deposit-taking institutions (DTIs) with the appropriate collateral, using funds placed at the Central Bank by DTIs with surplus liquidity for on-lending to the borrowing institutions.

During this period of system-wide stress, Jamaica's economy was severely impacted by the global financial turmoil. Real GDP declined by 1.6% for FY2008/2009, with economic conditions deteriorating sharply in the second half of the year. Bauxite and alumina production and exports fell by about 60%, while remittances –a traditional source of balance of payments support– declined by 33%. The value of the Jamaica dollar vis-à-vis the US dollar depreciated by 10% in the December 2008 quarter compared to 1% average depreciation for the first three quarters of 2008. In addition, similar to other developing countries, the external credit market was closed to Jamaica. This damaged investor confidence, especially with regard to the fiscal and debt dynamics and their sustainability. Notably, growth in NPLs for DTIs was also adversely impacted by the international economic slowdown, rising by over 40% over the crisis period. During the first quarter of 2010, the domestic financial environment returned to relative stability, which was underpinned by the signing of a 27-month stand-by arrangement with the IMF in that quarter.

Regarding the construction of the conditional variables, similar to Borio and Lowe (2002) and Drehmann et al. (2011), this paper is concerned with cumulative processes in contrast to levels or growth rates. Specifically, the focus is on the deviation of variables from their respective long-term trends, above explicit thresholds. Trends are determined using only ex ante information and are measured as deviations from one-sided Hodrick-Prescott filters, calculated recursively up to time t. The respective gaps are computed as the difference between the values of the variable and its trend at t. Consistent with Hodrick and Prescott (1991), to capture the cumulative buildup of imbalances, the smoothing parameter lamda (λ) is set to 1,600 for each of the quarterly data series used. However, this choice of λ is notable different from previous advanced economy studies which find that setting λ equal to 400,000 (which is associated with less frequent crisis episodes compared to business cycles) yields better results in picking up the time trends of conditioning variables.

For robustness, multiple horizons are considered for the accumulation phase. Specifically, crisis signals from indicators are judged to be correct if a crisis occurs *at the end* of one-year-ahead and three-month-ahead horizons. Signals from indicators of the release phase can only occur within a shorter horizon as release of the capital buffer should occur contemporaneously with the period of distress.

A range of thresholds are considered for each indicator. The choice of the ideal threshold involves a trade-off between the cost of missing a crisis (type 1 error) and the cost of calling a crisis which turns out to be false (type 2 error). Minimizing the noise-to-signal threshold has been the popular method of finding optimal thresholds in past studies (pioneered by Kaminsky and Reinhart, 1999). However, this method of signal extraction may not be ideal as highlighted by Demirgüç-Kunt and Detragiache (1998), given the incentives for regulators to overweight the risk of type 1 errors. Borio and Lowe (2002) and Borio and Drehmann (2009) offer the simple alternative of minimizing the noise-to-signal ratio with the proviso that at least two-thirds of the crises are correctly predicted.

This paper relies on a more precise method of balancing the cost-benefit trade-off of choosing indicator thresholds through the construction of a correct classification frontier (CCF) or receiver operating characteristics (ROC) curve (see Jordà and Taylor, 2011; Berge and Jordà, 2011, and Drehmann et al., 2011). In particular, Berge and Jordà (2011) discuss the use of ROC curve analysis to evaluate the historical predictive ability of indicator variables when the utility trade-offs across outcomes are unknown. Jordà (2011) describes the chronology of indicator variables as potentially embodying the latent state of the financial cycle. Observable financial conditions variables are generated by a mixture of distribution with each state (non-crisis and crisis) determined by the indicator chronology. Comparisons of the empirical distributions obtained by sorting the indicator and financial conditions variables by state will determine the information content of each indicator chronology. Berge and Jordà (2011) present two non-parametric statistics which can be used to gauge correct classification, the Kolmogorov-Smirnov (KS) statistic and Wilcoxon-Mann-Whitney (WMW) rank statistic (see Kolmorgorov, 1933; Smirnov, 1939; Mann and Whitney, 1947; Wilcoxon, 1945).

3. BEHAVIOR OF CONDITIONING VARIABLES AROUND DOMESTIC CRISIS EPISODES

The potential conditioning variables are measured based on deviations of variables from their trends to reflect their underlying cyclicality. As discussed above, all gaps are calculated as differences from a one-sided Hodrick-Prescott filter. Hence, the trend considers only historical information up to time t for each variable and excludes the future path of the given variable.

As discussed in Drehmann et al. (2010, 2011), the variables can be classified into three categories: the *macroeconomy*, *banking sector activity* and *funding costs*. The variables evaluated in this paper are similar to those in Drehmann et al. (2010, 2011). However, this paper also considers the relative behavior of *credit to the public sector* as well as *investment in public sector securities* given the dominant role of the public sector in the economy throughout the sample period.

The variables that relate to the macroeconomy include: credit (private and public)-to-real GDP, real credit growth, real investment growth, credit plus investment-to-real GDP, credit plus investment growth, investment to real GDP, private sector credit-to-real GDP and public sector credit-to-real GDP[2,3]. Other macroeconomic series evaluated are inflation, real GDP growth, real M2J growth and JSE Index growth. These variables are typically used as leading credit cycle indicators as they tend to display strong growth preceding systemic financial downturns. As shown in Figure 1, credit-to-real GDP, private sector credit-to-real GDP, public sector credit-to-real GDP and credit plus investment-to-real GDP, all rise leading up to a crisis episode, indicating their usefulness for signaling the accumulation phase. In contrast, real GDP growth declines significantly before a crisis, suggesting that it may be a useful variable for the release phase.

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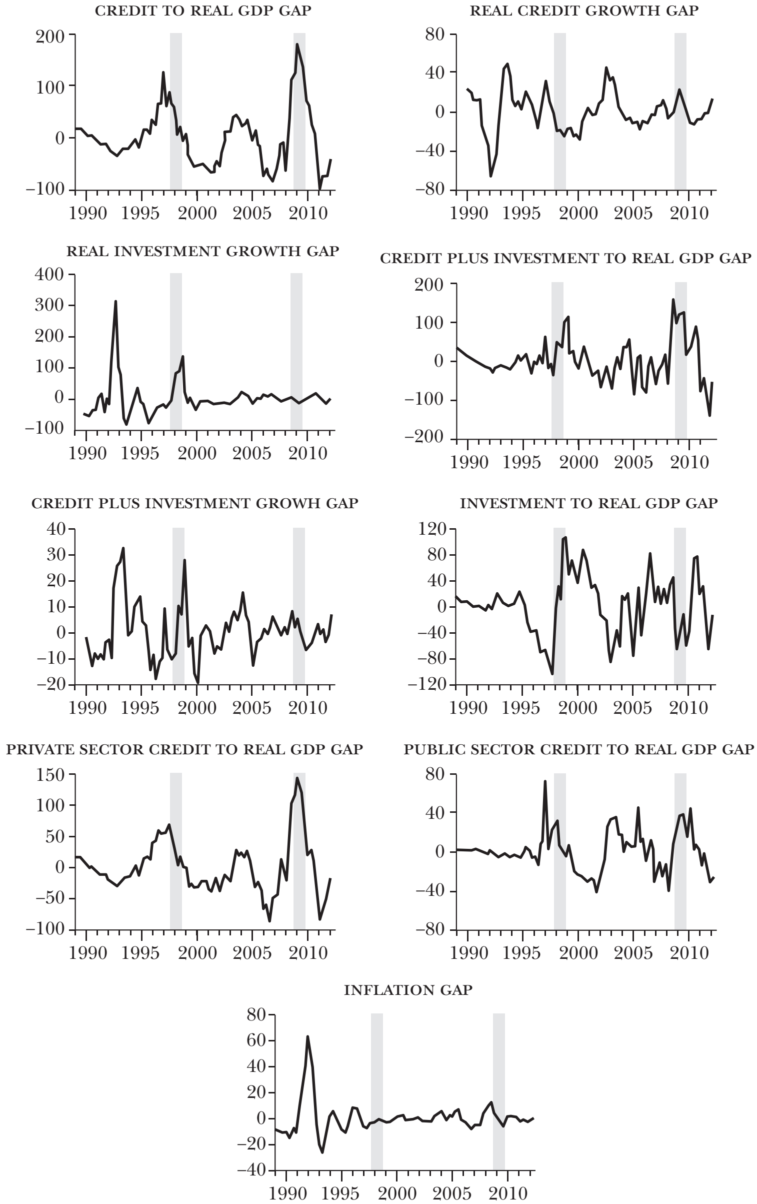
[2] Real GDP is used as the normalizing variable given the unavailability of a long enough official series for nominal GDP.

[3] Growth variables are calculated as the four-quarter change (in percent).

Figure 1a

BEHAVIOR OF CONDITIONING VARIABLES AROUND CRISES[1]

(Percentage)

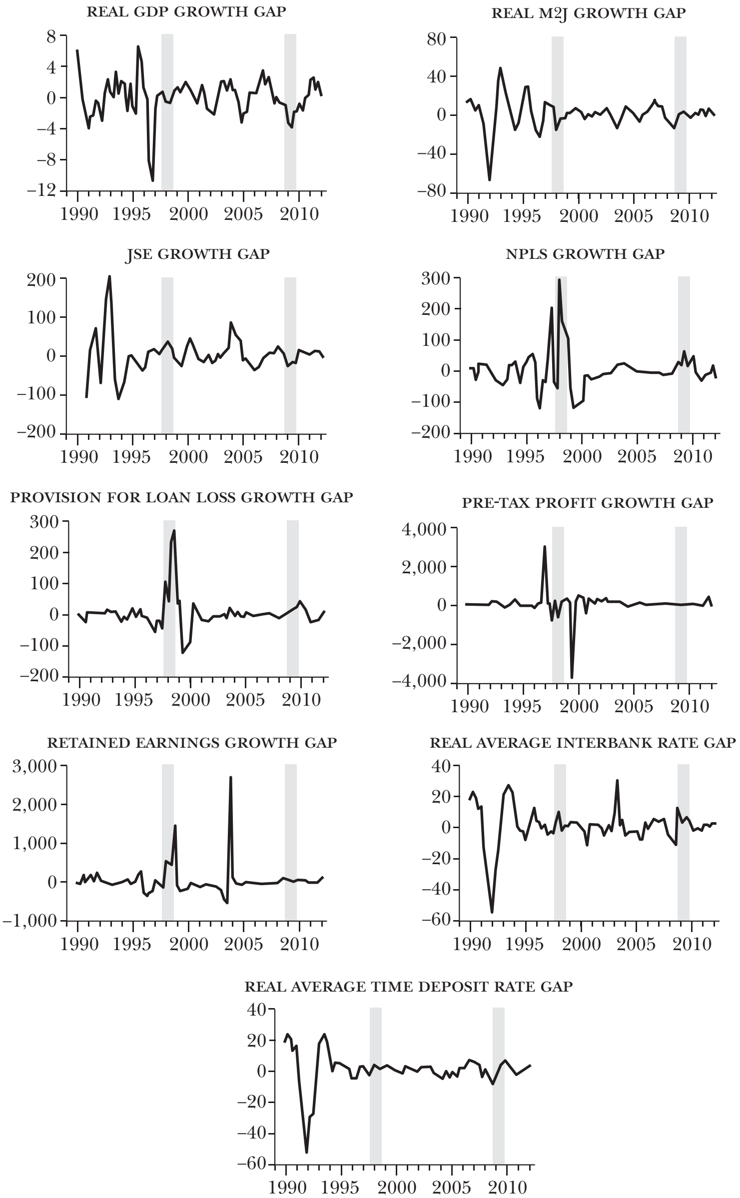


[1] Areas shaded in gray denote crisis episodes.

Figure 1b

BEHAVIOR OF CONDITIONING VARIABLES AROUND CRISES[1]

(Percentage)



[1] Areas shaded in gray denote crisis episodes.

The banking sector variables evaluated are growth in NPLs, provision for loan loss growth, pre-tax profits growth and retained earnings growth. Changes in the two former variables appear to be fairly coincident with the financial cycle. Growth in provision of loan loss, in particular, seems to be a good candidate for the release phase. Pre-tax profits growth and retained earnings growth exhibit weak performance for both the accumulation and release phases, especially for the second crisis episode. Finally, real monthly average (mid-point) interbank and real weighted average time deposit rates are the funding cost variables evaluated. Signals from these measures appear relatively noisy and do not perform well around the crisis episodes.

4. EVALUATION OF INDICATORS AND THRESHOLDS USING ROC CURVE ANALYSIS

Let $S_t \in \{0,1\}$ denote an observed financial conditions variable, with 1 indicating that t is a crisis period (quarter), and y_{t-h} be an indicator variable at time $t-h$ for h=0,1,2...H. Also let $\hat{S}_t(h) = I(y_{t-h} > c_h)$ denote a probability prediction about S_t, where the $I(.)$ indicator function equals 1 if true and c_h denotes the threshold related to the h-period ahead prediction. Assuming h=0, define the following conditional probabilities:

$$TP(c) = P\left[y_t \geq c \mid S_t = 1\right] \quad (1)$$

$$FP(c) = P\left[y_t \geq c \mid S_t = 0\right], \quad (2)$$

where $TP(c)$ is the true positive, sensitivity or recall rate and $FP(c)$ is the false positive, 1-specificity rate or type 1 error. The relationship between $TP(c)$ and $FP(c)$ describes the ROC curve. The threshold or cut-off value provides the decision rule to divide the conditioning variable according to the crisis states (see Table 1).

The ROC curve plots the combinations $\{TP(c), FP(c)\}$ for $c \in \{-\infty, \infty\}$. When $c \to \infty, TP(c) = FP(c) = 0$ and, alternatively, when $c \to -\infty, TP(c) = FP(c) = 1$. The ROC curve may be represented with the Cartesian convention $\{ROC(r), r\}_{r=0}^{1}$, where $ROC(r) = TP(c)$ and $r = FP(c)$. If y_t is uninformative regarding the crisis period, $TP(c) = FP(c) \forall c$ and the ROC curve would be the 45° line in [0, 1]×[0, 1] space. Conversely, if y_t is perfectly informative, then the ROC curve would hug the north-east corner in [0, 1]×[0, 1].

Table 1

RESULTS FROM DECISION RULE

		Observed	
		Crisis	*No crisis*
Decision	Above threshold	True positive prediction (sensitivity)	False Positive prediction (1-specificity)
	Below threshold	False negative prediction (1-sensitivity)	True negative prediction (specificity)

As an alternative to the noise-to-signal approach for indicator evaluation, consider the expected utility given the cost-benefit trade-off of each type of error given by:

$$U(r) = U_{11} ROC(r)\pi + U_{01}(1 - ROC(r))\pi + U_{10} r (1-\pi), \\ U_{00}(1-r)(1-\pi) \tag{3}$$

where U_{ij} is the utility associated with the prediction i given that the true state is j, $i, j \in \{0,1\}$ and π is the unconditional probability of observing a crisis episode over a specific horizon.

Maximization of [3] indicates that the optimum, c^*, can be obtained by solving:

$$\frac{dROC}{dr}=\frac{U_{00}-U_{10}}{U_{11}-U_{01}}\frac{(1-\pi)}{\pi} \quad , \tag{4}$$

which is the point where the slope of the ROC curve equals the expected marginal rate of substitution between net utility of accurate crisis and non-crisis prediction.

In addition, the slope of the ROC curve is the likelihood ratio of probability density function (*pdf*), given by θ, for the sub-sample of $y_t\left(y_t^c\right)$ for which S_t=1 and the *pdf* for the sub-sample of $y_t\left(y_t^{nc}\right)$ for which S_t=0 given by φ, so that:

$$\frac{dROC}{dr}=\frac{\varphi\left(\Theta^{-1}(1-r)\right)}{\theta\left(\Theta^{-1}(1-r)\right)}, \tag{5}$$

where Θ is the cumulative *pdf* associated with θ. Furthermore, the (KS) statistic is used to determine the optimal operating point (c^*) by the maximization of the distance between *TP(c)* and *FP(c)*, under the assumptions $U_{ii}=1$, $U_{ij}=-1$ and $\pi=0.5$ (see Figure 2).

The measure of overall classification ability is the area under the ROC (AUROC) curve:

$$AUROC=\int_0^1 ROC(r)\,dr; \qquad AUROC\in[0.5,1], \tag{6}$$

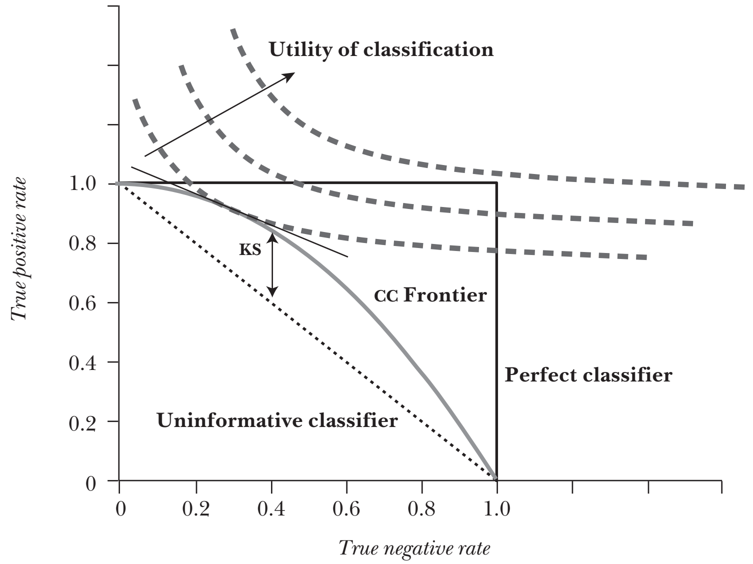
which may be computed as the rank-sum statistic:

$$\widehat{AUROC}=\frac{1}{n_0 n_1}\sum_{j=1}^{n_0}\sum_{i=1}^{n_1}\left\{I\left(y_j^{nc}<y_i^c\right)+\frac{I\left(y_j^{nc}=y_i^c\right)}{2}\right\}, \tag{7}$$

where $I(\cdot)$ is the indicator function that equals 1 when the argument is true and 0 otherwise, n_0 and n_1 are the number of observations in y_j^{nc} and y_i^c, respectively, and the latter term in

Figure 2

RECEIVER OPERATING CHARACTERISTICS CURVE



Source: O. Jordà, *Discussion of Anchoring Countercyclical Capital Buffers: The Role of Credit Aggregates*, Working Paper, University of California, Davis, 2011.

7 is used to correct tied ranks (see Jordà and Taylor, 2010). The AUROC is a WMW rank statistic which is equal to 1 in the case of a perfect classifier and 0.5 (450 line) for a completely uninformative classifier. In addition, under standard regularity conditions (see Hsieh and Turnbull, 1996):

$$\sqrt{n_1}\left(\widehat{AUROC}-0.5\right)\xrightarrow{d}N\left(0,\sigma^2\right) \tag{8}$$

$$\sigma^2=\frac{1}{n_0n_1}\left[\begin{array}{l}AUROC\left(1-AUROC\right)+\left(n_1-1\right)\left(\phi_1-AUROC^2\right)+\\+\left(n_0-1\right)\left(\phi_2-AUROC^2\right)\end{array}\right]^{1/2}$$

where $\phi_1 = AUROC/\left(2-AUROC\right)$

and $\phi_2 = 2AUROC^2/\left(1+AUROC\right)$.

5. EMPIRICAL RESULTS

Before conducting the ROC curve assessment, the signal extraction method was employed to assess the performance of potential conditioning variables over different thresholds and horizons. Specifically, the values of thresholds to be examined for each indicator were based on visual assessments of the data vis-à-vis the crisis periods (see Figure 1). Signals, $S(y_{t-h})$, can either take on the value of 0 or 1 depending on whether y_{t-h} is below or above the threshold value, c_h. A signal of 1 (0) was judged to be correct only if a crisis (no crisis) occurred at the end of the prediction horizon[4]. One-year-ahead, three-months-ahead and zero-year-ahead prediction horizons were examined. Notably, these horizons, particularly the latter two, would give the Central Bank a relatively short lead time to implement capital buffers. Longer horizons of two and three years were also examined, but with inferior results. This shortcoming of relatively high volatility in the indicator series may be a feature of small developing economies.

As discussed earlier, given that the preferences of regulators are not observed, the best threshold is determined when using the signals extraction method by minimizing the noise-to-signal ratios conditional on at least two-thirds of the crises being correctly predicted (see Borio and Drehmann, 2009). As depicted in Table 2, bold fonts are used in the columns labeled *Predicted* to indicate threshold values that are consistent with a condition of a crisis prediction rate of at least 66%. In addition, bold fonts and shaded cells in columns labeled N/S indicate the lowest noise-to-signal ratio for threshold values that satisfy the condition.

For the one-year-ahead horizon, private sector credit-to-real GDP gap at the 20% threshold value, achieved the lowest noise-to-signal ratio of 22% as well as the highest percent of correct

[4] This is a more conservative definition compared to Borio and Lowe (2002) and Drehmann et al (2010, 2011) where signals of 1 (0) are judged to be correct if a crisis (no crisis) occurred *at any time* within the prediction horizon.

predictions of 81%. Thresholds of 30% and 40% for this variable also achieve above two-thirds successful predictive rates, albeit, at slight higher noise-to-signal ratios. Credit (private and public)-to-real GDP gap is the only other variable to satisfy the condition of a crisis prediction rate of at least 66% (75%) and achieved a noise-to-signal ratio of 29% at a 25% threshold value.

At the three-month-ahead horizon, the results are a bit different. Credit-to-real GDP gap still satisfies the condition of a crisis prediction rate of at least 66%, but now at both the 25% threshold value (with noise-to-signal ratio of 21%) and 50% threshold value (with noise-to-signal ratio of 26%). However, in contrast to results for the one-year-ahead horizon, private sector credit-to-real GDP gap did not attain the minimum condition for the prediction ratio.

The results at contemporaneous horizon are similar to those for the three-month-ahead horizon. Only credit-to-real GDP gap satisfies the condition of a crisis prediction rate of at least 66% (81%). Similar to the results for the three-month-ahead horizon, this condition is held at both the 25% and 50% threshold values.

Table 3 presents the AUROC for each indicator over the three horizons. Consistent with the signal extraction method discussed above, the AUROC for the HP-filtered credit-to-real GDP gap, credit plus investment-to-real GDP gap, private sector credit-to-real GDP gap and public sector credit-to-real GDP gap all have significant predictive value for crisis episodes. In contrast to the alternative method, however, is the fact that significant predictive values for these variables are attained for all horizons considered.

Table 2

PERFORMANCE OF POTENTIAL CONDITIONING VARIABLES FOR DIFFERENT SIGNALING HORIZONS

Conditioning variables	*Threshold*	*1-year-ahead prediction (%)*				*3-month-ahead prediction (%)*				*0-year-ahead prediction (%)*			
		Type 1	*Type 2*	*Predicted*	*N/S*	*Type 1*	*Type 2*	*Predicted*	*N/S*	*Type 1*	*Type 2*	*Predicted*	*N/S*
Credit to real GDP	25	14	25	**75**	**29**	9	19	**81**	**21**	9	25	**75**	28
	50	8	38	63	41	3	25	**75**	26	1	25	**75**	**25**
	75	5	69	31	73	1	56	44	57	1	56	44	57
Real credit growth	15	18	88	13	106	15	88	13	103	15	88	13	103
	20	16	88	13	105	14	88	13	101	14	88	13	101
	30	11	94	6	105	8	94	6	102	8	94	6	102
Real investment growth	15	18	94	6	114	12	81	19	93	11	75	25	84
	20	16	94	6	112	11	81	19	91	10	75	25	83
	25	15	94	6	110	10	81	19	90	8	75	25	82
Credit and investment to real GDP	30	21	69	31	87	12	38	63	43	12	38	63	43
	40	14	63	38	72	4	50	50	52	4	44	56	46
	50	11	69	31	77	3	56	44	58	4	56	44	59
Credit and investment growth	5	26	94	6	127	22	81	19	104	21	81	19	102
	10	15	100	0	118	12	100	0	114	11	94	6	105
	15	10	100	0	111	7	100	0	107	5	94	6	99

	70	12	100	0	114	10	100	0	111	8	94	6	102
Investment to real GDP	75	11	100	0	112	8	100	0	109	7	94	6	101
	80	11	100	0	112	4	100	0	104	3	94	6	96
	20	13	19	**81**	**22**	17	44	56	53	19	56	44	70
Private sector credit to real GDP	30	6	25	**75**	27	12	63	38	71	14	75	25	88
	40	4	25	**75**	26	9	63	38	69	12	75	25	85
	20	14	69	31	80	8	56	44	61	10	56	44	62
Public sector credit to real GDP	35	8	88	13	95	3	94	6	96	3	75	25	77
	40	5	94	6	99	3	94	6	96	3	94	6	96
	30	11	75	25	84	5	69	31	73	5	63	38	66
Non-performing loans growth	40	11	75	25	84	5	63	38	66	4	56	44	59
	50	16	75	25	90	3	63	38	64	1	56	44	57
	15	21	88	13	110	14	63	38	72	11	50	50	56
Provision for loan loss growth	30	11	88	13	98	5	75	25	79	3	63	38	64
	60	5	94	6	99	1	88	13	89	0	81	19	81
	5	16	63	38	75	15	75	25	88	18	88	13	106
Inflation	10	10	88	13	97	8	88	13	95	8	94	6	102
	-	-	-	-	-	-	-	-	-	-	-	-	-

continues

Conditioning variables	Threshold	1-year-ahead prediction (%)				3-month-ahead prediction (%)				0-year-ahead prediction (%)			
		Type 1	Type 2	Predicted	N/S	Type 1	Type 2	Predicted	N/S	Type 1	Type 2	Predicted	N/S
Real GDP growth	3	10	88	13	97	10	100	0	111	10	100	0	111
	3.5	5	88	13	93	5	100	0	106	5	100	0	106
	4	4	88	13	91	4	100	0	104	4	100	0	104
Real M2J growth	10	16	88	13	105	15	94	6	110	15	94	6	110
	15	11	94	6	105	10	100	0	111	10	100	0	111
	20	10	94	6	104	8	100	0	109	8	100	0	109
Real monthly average inter-bank rate	3	30	56	44	81	30	69	31	98	29	63	38	88
	5	19	81	19	101	16	81	19	97	16	81	19	97
	10	18	94	6	114	16	100	0	120	16	100	0	120
Real weighted average time deposit rate	3	33	88	13	130	32	94	6	137	30	88	13	125
	4	26	100	0	135	23	100	0	130	23	94	6	122
	5	19	100	0	124	16	100	0	120	16	94	6	112

Notes: A signal of 1 (0) was judged to be correct only if a crisis (no crisis) occurred at the end of the prediction horizon.
Type 1 error refers to when no signal is issued and a crisis occurs.
Type 2 error refers to when a signal is issued and no crisis occurs.
Predicted refers to the percentage of crises correctly predicted. Values in bold font in this column indicate that more than 66% of crisis quarters were correctly predicted.
The noise-to-signal ratio (N/S) is defined as the fraction of type 2 errors divided by one minus the fraction of type 1 errors. Values in bold font and shaded cells in this column indicate the lowest N/S ratio among the threshold values that are associated with indicators showing a correct prediction rate of at least 66 percent.

Table 3

PERFORMANCE OF POTENTIAL CONDITIONING VARIABLES USING THE AUROC CURVE FOR DIFFERENT SIGNALING HORIZONS

Conditioning variables	*0 Year*	*3 Months*	*1 Year*
Credit to real GDP	**0.95**	**0.94**	**0.87**
Real credit growth	0.53	0.50	0.54
Real investment growth	0.53	0.51	0.43
Credit and investment to real GDP	**0.81**	**0.81**	**0.73**
Credit and investment growth	0.47	0.45	0.41
Investment to real GDP	0.27	0.29	0.30
Private sector credit to real GDP	**0.66**	**0.71**	**0.82**
Public sector credit to real GDP	**0.86**	**0.77**	**0.64**
Non-performing loans growth	**0.73**	**0.68**	**0.64**
Provision for loan loss growth	**0.64**	0.61	0.58
Inflation	0.42	0.48	0.52
Real GDP growth	0.24	0.24	0.35
Real M2J growth	0.39	0.34	0.37
Real monthly average inter-bank rate	0.53	0.49	0.52
Real weighted average time deposit rate	0.44	0.40	0.40

Notes: AUROC curve of conditioning variables relative to crisis periods for 0-year-ahead, three months-ahead and one year-ahead predictions.
Areas statistically different from 0.5 using the one-tailed WMW test are denoted by bold font and shaded cell at the 99% level of significance and bold font at the 95% level of significance.

Furthermore, credit plus investment-to-real GDP gap, public sector credit-to-real GDP gap, NPLs growth gap and provision for loan loss growth gap all show significant predictive power especially for the contemporary horizon. Notably, these indicators were not supported as being useful under the conditions of the signal extraction method. Notwithstanding, the more robust AUROC method provides strong support for the two latter indicator variables, in particular, to be used as lagging indicators to guide the release phase. Specifically, as indicated by the BCBS, release of the buffer add-on should be considered when in a situation of system-wide banking system losses. Accordingly, NPLs growth gap and provision for loan loss growth gap both satisfy this scenario in sufficiently promptly signaling the timing of the release.

Basel Committee (2010) offers guidelines for countries operating the countercyclical capital buffer regime. The Committee also developed a formula that offers a buffer level that varies with the size of the deviation of the cyclical components of conditioning variables from their long-term trends. The formula links a conditioning variable to a capital adjustment factor. This add-on factor equals zero in bad times and increases linearly in the conditioning variable to a set maximum level. In practice, each national authority makes its own decision on the choice of conditioning variables and the statistical tool that splits these variables into their trend and cyclical components.

The formula for the countercyclical add-on may be presented as:

$$k_t = \begin{cases} 0 & \text{if } y_t < L \\ \dfrac{y_t - L}{H - L} k_{max} & \text{if } L \leq y_t \leq H \\ k_{max} i & \text{if } H < y_t \end{cases} \quad (9)$$

The choice of lower and upper threshold gap levels, L and H, are critical to the speed and timing of buffer adjustment in relation to the buildup of systemic risk. The Basel Committee has established broad criteria to determine threshold gap levels as a starting guide to the relevant authorities for deciding the buffer add-on (BCBS, 2010):

> *1)* L should be low enough, so that banks are able to build up capital in a gradual fashion before a potential crisis. As banks are given one year to raise additional capital, this means that the indicator should breach the minimum at least 2-3 years prior to a crisis,
>
> *2)* L should be high enough, so that no additional capital is required during normal times,
>
> *3)* H should be low enough, so that the buffer would be at its maximum prior to major banking crises.

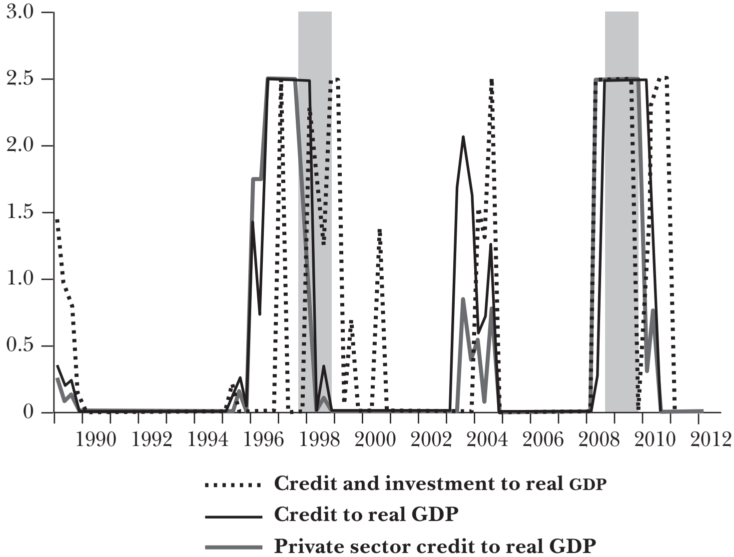
Figure 3 illustrates how the countercyclical buffers would have affected Jamaica's commercial banks using the HP-filtered credit-to-real GDP gap, credit and investment-to-real GDP gap and private sector credit-to-real GDP gap as conditioning variables (as supported by the AUROC method) over the sample period of this study. In accordance with the Basel (2010) guidelines, the maximum buffer add-on (Kmax) was set at 2.5% of risk-weighted assets. The Figure depicts that evolution of capital add-on when L=15% and H=50% for purely exposition purposes. For both crisis periods, the buffer would reach the maximum value prior to the onset of the crisis. This feature of the conditioning variables provides justification for setting λ = 16,000 which is below the λ = 400,000 used for studies on advanced countries.

Whereas the build-up phase associated with these conditioning variables is sufficient for the first crisis episode (two years), it is short (one quarter) in the case of the second crisis period. Interestingly, the conditioning variables indicate a build-up of buffer capital in the 2003 to 2004 period, albeit, with a shorter duration and smaller magnitude compared to the crisis episodes. However, this period is not considered a crisis episode given the maintenance of low NPL levels in the banking sector as well as the presence of abundant market liquidity. Notwithstanding the absence of an *official* crisis, commercial banks

Figure 3

HISTORICAL PERFORMANCE OF COUNTERCYCLICAL CAPITAL BUFFERS FOR JAMAICA'S COMMERCIAL BANKS

(Percentage)



Source: O. Jordà, *Discussion of Anchoring Countercyclical Capital Buffers: The Role of Credit Aggregates*, Working Paper, University of California, Davis, 2011.

operated within a severely challenging macroeconomic environment within this period triggered by the announcement of a large fiscal disjuncture and a downgrade in the rating of Jamaica's sovereign debt by Standard and Poor's at the end of 2002. Given the deteriorated domestic financial conditions, particularly in the foreign exchange market, the Central Bank instituted a Special Deposit reserve requirement for DTIs on 10 January 2003 and adjusted interest rates sharply upward on three occasions during the first half of 2003 in order to constrict the excess market liquidity. Hence, in the context of the tightening in monetary policy during 2003, it can be reasonably argued that the actions of the Central Bank averted a looming boom-bust cycle at that time of weakened sovereign creditworthiness.

6. CONCLUSION AND POLICY IMPLICATIONS

This paper provides support for the findings of other studies (eg., Borio and Drehmann, 2009) that policymakers can be guided by conditioning variables at one-year and three-month horizons such as credit-to-GDP, NPLs growth and provisions for loan loss growth in their design of countercyclical capital buffers. It is acknowledged that reliance on these relatively short horizons, which may be due to relatively high volatility in the indicator series, would give policymakers relatively little implementation lead time. This shortcoming may be a feature of small developing economies.

The novelty of this paper comes from the finding that banking sector variables reflecting sovereign risk build-up (namely the level of public sector credit and investments in public sector securities) perform successfully as conditioning variables for Jamaica. Hence, other economies with a history of fiscal dominance and public sector crowding out of private sector credit should explore variables that reflect sovereign risk build-up in guiding the accumulation and release phases of a capital buffer requirement for their banking sectors.

Importantly, the accurate timing of implementing a countercyclical capital buffer would be crucial, as it would have to be established only in a clear up-cycle period. Otherwise, it could have negative implications in terms of banks' financial strength, stakeholders' perceived confidence in the sector and the reputation of the central bank. Against this pre-requisite, although this paper focuses on computing the long-run trend by the HP filter as a guide for the buffer to be consistent with the proposed method of the BCBS, alternative statistical filters may be applied to obtain comparative results for robustness checks[5]. Nonetheless, experimenting with other statistical detrending approaches is unlikely to dramatically improve the performance of the indicators. Indeed, an alternative approach such as that proposed by Geršl and Seidler (2010)

[5] Alternative filters include Beveridge and Nelson (1981) and band-pass, among others.

could be explored which relies on an out-of-sample technique to estimate the fundamental-based equilibrium credit level and may be more appropriate for small developing economies such as Jamaica.

In addition, Jamaica's macroprudential authorities will need to build up a longer time series of data on these indicators to strengthen the decision-making framework regarding implementing countercyclical capital buffers. Then further disaggregation of variables should be explored to refine the efficiency of relevant information contained in the indicators. For example, credit could be further broken down by institution size, currency and economic sector.

Importantly, the regulatory approach to mitigating procyclicality of the financial system should be all-inclusive, covering all financial institutions to mitigate arbitrage opportunities. In addition to the countercyclical buffer requirement, other elements of the prudential framework should also be utilized. For instance, excessive credit growth (and subsequent downward shift in credit quality) stems essentially from inadequate risk management practices. While the central bank may be in the best position to assign the capital requirements commensurate to the degree of risk taken by banks during times of credit growth, it should not be left as a holistic rule-based mechanism.

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