CENTER FOR LATIN AMERICAN MONETARY STUDIES
Regional Association of Central Banks

MONEY AFFAIRS

Volume XXIII, Number 1
January-June 2010

MONEY AFFAIRS is a biyearly publication of the Center for Latin American Monetary Studies (CEMLA), Durango nº 54, Mexico City, D. F., 06700. Printed in Master Copy, S. A. of C. V., Av. Coyoacán nº 1450, Mexico City, D. F., 03220. 310 copies. ISSN 0187-7615.

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VOLUME XXIII, NUMBER 1, JANUARY-JUNE 2010

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The dynamics of bank spreads in the Jamaican banking sector: an empirical assessment

1. INTRODUCTION

A key aspect of monetary policy transmission is the extent to which policy rates affect market rates, in particular money market rates, and how these changes affect banks' interest rates. This issue is important in assessing the effectiveness of the monetary policy since the pass-through of market rates to bank retail rates is a critical element in the monetary transmission process. A common finding in the international literature is that market conditions are not passed on to bank interest rates immediately.

The empirical literature provides evidence that corporate lending rates, in particular, respond sluggishly to market rates (see Cottarelli and Kourelis, 1994; Borio and Fritz, 1995; Mojon, 2001). For instance, when the central bank takes a monetary policy stance, there is the presumption that these

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official rate changes will feed through to influence the array of short-term money market rates and the rates set on retail products, such as deposit and loan accounts and mortgages. However, the extent to which monetary policy can be effective is heavily influenced by factors such as banks' price setting behavior.

It is widely established that an important relationship exists between banks price setting behavior and the transmission of monetary policy. For instance, as banks price their products more in line with the market, the transmission of monetary policy is typically smoother.[1] In addition, studies on the monetary transmission process in Jamaica have found this to be the case. Robinson (2000) found that the absolute size of banking spreads in Jamaica is an outcome of the factors that have defined the economic environment, such as uncertainty, market structure and inefficiency. Ultimately, banks pass on these costs in terms of higher (lower) interest premiums on loans (deposit) rates.[2]

The focus in this paper will be on the price-setting behavior of Jamaican banks as well as the pass-through mechanism from changes in official policy rates through market rates to bank rates. By applying the econometric framework originally developed by Ho and Saunders (1981), the paper estimates the dynamic adjustment of bank spreads (i.e. the difference between the bank interest rate and its corresponding market rate for various bank loan and deposits categories) to changes in monetary policy as a function of various exogenous factors, such as bank competition and financial structures.[3]

The paper is structured as follows: section 2 presents a review of the literature underlying bank spreads. Section 3 breaks down the determinants of bank spreads and section 4 gives a description of variables used in the study. Section 5 outlines the econometric framework employed in the investigation of the pass-through and its determinants and section 6 describes the data used in the study. Estimation results are

[1] See Hannan and Berger (1991).

[2] Robinson (2000) showed that cash reserve requirements have minimal impact on bank spreads in Jamaica. In particular, the findings showed that cash reserve requirements represent only 2.0 percentage points of a loan rate of 22%.

[3] See Courvoisier and Gropp (2001).

shown in section 7, robustness checks are presented in section 8, and section 9 outlines the conclusions and policy implications of the study.

2. LITERATURE REVIEW

In recognizing the two-sided nature of bank spreads, several authors model lending and deposit rates simultaneously. One of these models is the dealership approach, originally proposed by Ho and Saunders (1981, 1982). Ho and Saunders (1982) advocated a two-step procedure in explaining the determinants of bank interest spreads. In the first-step, bank interest margin is regressed against a set of bank-specific variables such as non-performing loans, operating costs, capital to asset ratio and time dummies. The time dummy coefficients of this regression are interpreted as being measures of the *pure* component of a country's bank spread. In the second-step, the constant terms are regressed against variables reflecting macroeconomic factors. For this second step, the inclusion of a constant term aims at capturing the influence of factors such as market structure or risk-aversion coefficient, which reflect neither bank-specific observed characteristics nor macroeconomic elements.

Following the work of Ho and Saunders (1981, 1982), Hannan and Berger (1991) showed that the pace of adjustment of deposit rates to policy rates depends on the elasticity of deposit supply. Further, the elasticity of supply may depend on factors such as market concentration and the depositor base of the bank. Overall, the studies found that banks tend to adjust rates in asymmetric fashion, as deposit rates tend to be more rigid in the case of interest rate increases than in periods of decreasing interest rates.

Scholnick (1996) argued that the issue of interest rate rigidity is best examined using the cointegration and error correction methodology, by utilizing results on speeds of adjustment of retail (lending and deposit) rates to changes in wholesale (interbank or money market) rates. A further innovation by Scholnick (1996) is the use of an asymmetric error correction methodology, which makes it possible to examine whether retail rates have greater rigidity upwards or downwards.

Angbazo (1997) studied the determinants of bank net interest margins using a sample of US banks' data over the period 1989 to 1993. The empirical model for the net interest margin/bank spreads is postulated to be a function of a wide cross-section of variables that impact banks' price setting behavior. The variables covered include default risk, interest rate risk, an interaction term for default and interest risk, liquidity risk, leverage, implicit interest payments, opportunity cost of non-interest bearing reserves, management efficiency and a dummy variable for states with branch restrictions. The results for the pooled sample suggest that the proxies for default risk (ratio of net loan charge-offs to total loans), the opportunity cost of non-interest bearing reserves, leverage (ratio of core capital to total assets) and management efficiency (ratio of earning assets to total assets) are all statistically significant and positively related to bank interest margins. The ratio of liquid assets to total liabilities, a proxy for liquidity risk, was inversely related to bank net interest margin. The other variables were not statistically significant.

Demirguc-Kunt and Huizinga (1999) also investigated the determinants of bank interest margins using bank-level data. This study covered 80 countries over the period 1988-1995 and utilized regressors capturing bank characteristics, macroeconomic conditions, explicit and implicit bank taxation, deposit insurance regulation, financial structure as well as legal and institutional indicators. Their findings showed that bank interest margins are positively influenced by the ratio of equity to the lag of total assets, the ratio of loans to total assets, a foreign ownership dummy, bank size, the ratio of overhead costs to total assets, the inflation rate and the short-term market real interest rate. The ratio of non-interest earning assets to total assets, on the other hand, was negatively related to bank interest margin, while output growth did not have an impact on bank spreads.

In investigating the determinants of banks' interest margins, Brock and Rojas-Suarez (2000) applied the two-step procedure developed by Ho and Saunders (1982) for a sample of Latin American countries. For each country, the first-stage of regressions for bank spread included variables such as the slope of the yield curve and time dummies as well as various microeconomic variables covering non-performing loans

(NPLs), capital ratio, operating costs and a measure of liquidity. Their findings show positive and significant results for the capital, cost and liquidity ratios. However, the evidence was mixed regarding the impact of non-performing loans. They explained that this finding reflected inadequate provisioning for loan losses, which was used as a proxy for NPLs, thereby lowering the spread in the absence of adequate loan loss reserves.

3. HO AND SAUNDERS (1981) & MAUDOS AND FERNÁNDEZ (2004) MODEL OF BANKS' PRICE-SETTING BEHAVIOR

In this paper, the determinants of banks' price-setting behavior are analyzed using the influential model developed by Ho and Saunders (1981). This paper also builds on the work of Maudos and Fernández (2004), which extended the original model of Ho and Saunders (1981) to include the production costs associated with the process of intermediation between deposits and loans. The theoretical model captures a number of factors that influence banks' price setting behavior such as the competitive structure of the market, operating costs, the volatility of money market rates, credit risk as well as the interaction between interest rate risk and credit risk.

Similar to the Ho and Saunders model, a bank is viewed as a dealer in the credit market and acts as an intermediary between the demanders and suppliers of funds. Furthermore, decisions are assumed to be made in a finite horizon, where the bank maximises the expected utility of terminal wealth. The bank has three components to its wealth portfolio. The first component is its initial wealth, W_0, which is invested in a diversified portfolio. Wealth is determined by the difference between the assets and liabilities. Assets comprise of the sum of loans (L) and money market assets (M), while liabilities consist of deposits (D). Thus, initial wealth is, $W_0 = L_0 - D_0 + M_0$. The second component is a net credit inventory, I, which is the difference in market values of loans and deposits, $I = L - D$. It is assumed that the credit inventory will be subject to interest rate risk. The third component is the banks' money market position (M).

The operating or production costs of a banking firm are assumed to be a function of the deposits captured, $C(D)$, and the loans made, $C(L)$, so that the cost of net credit inventory $C(I)$ can be expressed as $C(I) = C(L) + C(D)$.[4] Therefore, the bank's wealth portfolio at the end of the decision period is the sum of initial wealth, money market position, and net credit inventory less the cost of these net credit inventories. This can be expressed as follows:

$$(1) \qquad \begin{aligned} W &= (1 + r_I + Z_I)I_0 + (1 + r + Z_M)M_0 - C(I_0) \\ &= I_0 + I_0 r_I + I_0 Z_I + M_0 + M_0 r + Z_M M_0 - C(I_0) \\ &= W_0(1 + r_w) + I_0 Z_I + M_0 Z_M - C(I_0)\ , \end{aligned}$$

where, r_w, r_I, r are the expected rates of return on initial wealth, net credit inventory and the net cash position, respectively.[5] Uncertainty faced by the banks is captured by Z_L and Z_M, which represent interest rate risks and credit risks, respectively. The variables Z_M and Z_L are random variables distributed $Z_M \sim N\left(0, \sigma_M^2\right)$ and $Z_L \sim N\left(0, \sigma_L^2\right)$, respectively. The joint distribution of interest rate and credit risk assumes a bivariate normal function.

Through the intermediation process, banks continue to accumulate wealth based on the intermediation margins on new deposits and loans. As such, banks set loan and deposit prices, p_L and p_D, respectively, and the quantity is determined exogenously, where:

$$(2) \qquad p_L = r + b \text{ and } p_D = r - a$$

where a and b are the margins for deposits and loans, respectively, relative to the money market interest rate.

The bank's decision problem in the face of these transaction and interest-rate risks is to determine the expected utility-maximizing deposit and loan rates, where spreads are determined by the margins on deposits and loans, $S = a + b$. The

[4] See Maudos and Fernández (2004).

[5] In equation (1) $r_I = \dfrac{r_L L_0 - r_D D_0}{I_0}$ and $r_w = r_I \dfrac{I_0}{W_0} + r\dfrac{M_0}{W_0}$ are the respective average profitability on net credit inventory and bank's initial wealth and $Z_I = Z_L \dfrac{L_0}{I_0} + Z_D \dfrac{D_0}{I_0} = Z_L \dfrac{L_0}{I_0}$ is the average risk of net credit inventory.

expected utility of wealth at the end of the period is approximated using the Taylor series expansion around the level of wealth, W, where $\bar{W} = E(W)$, and the expected utility of wealth is given by:

$$(3) \qquad EU(W) = U(\bar{W}) + U'(\bar{W})E(W - \bar{W}) + \frac{1}{2}U''(\bar{W})E(W - \bar{W})^2,$$

where it is assumed that the bank's utility function is concave, such that $U' > 0$ and $U'' < 0$ and, therefore, that the bank is risk averse.[6] When a new deposit, D, is made, if no additional credit is granted, whatever funds that are captured by the bank will be invested in the money market obtaining a return of $(r + Z_M)D$. Moreover, taking into consideration that $W - \bar{W} = L_0 Z_L + M_0 Z_M$ and given the existence of operating costs in the capture of deposits $C(D)$, substituting the new value of the final wealth in (3), the increase in expected utility associated with the new deposit is:[7]

$$(4) \quad \Delta EU(W_D) = EU(W_T) - EU(W)$$

$$= U'(\bar{W})\left[aD - C(D)\right] + $$

$$+ \frac{1}{2}U''(\bar{W})\begin{bmatrix} (aD - C(D))^2 + (L + 2L_0)L\sigma_L^2 + \\ (L - 2M_0)L\sigma_M^2 + 2(M_0 - L_0 - L)L\sigma_{LM} \end{bmatrix}.$$

Similarly, if a new request for credit is made for which there is also a cost of production, $C(L)$, the increase in expected utility for new loans is given as:

$$(5) \quad \Delta EU(W_L) = EU(W_T) - EU(W)$$

$$= U'(\bar{W})\left[bL - C(L)\right] + $$

$$+ \frac{1}{2}U''(\bar{W})\begin{bmatrix} (bL - C(L))^2 + (L + 2L_0)L\sigma_L^2 + \\ (L - 2M_0)L\sigma_M^2 + 2(M_0 - L_0 - L)L\sigma_{LM} \end{bmatrix}.$$

Similar to the Ho and Saunders (1981) model, it is assumed that loans and deposits are made randomly according to a Poisson process. As such, the probability of granting a

[6] If the bank were risk neutral, the bank would be an expected wealth maximizer. That is, the bank faces no risk associated with market rates or credit facilities.

[7] See Appendix A.

loan or capturing a deposit is represented as a decreasing function of the margins applied by the bank:

$$\begin{aligned} \Pr_D &= \alpha_D - \beta_D a, \\ \Pr_L &= \alpha_L - \beta_L b. \end{aligned} \tag{6}$$

The maximization problem, which is the linear combination of equations (4), (5) and (6), therefore becomes:

$$Max_{a,b} EU(\Delta W) = \left(\alpha_D - \beta_D a\right)\Delta EU(W_D) + \left(\alpha_L + \beta_L b\right)\Delta EU(W_L)\,, \tag{7}$$

where total spreads, s, is equal to:

$$\begin{aligned} S = {} & \frac{1}{2}\left(\frac{\alpha_D}{\beta_D} + \frac{\alpha_L}{\beta_L}\right) + \frac{1}{2}\left(\frac{C(L)}{L} + \frac{C(D)}{D}\right) - \\ & -\frac{1}{4}\frac{U''(\bar{W})}{U'(\bar{W})}\left[\left(L + 2L_0\right)\sigma_L^2 + (L + D)\sigma_M^2 + 2(M_0 - L)\sigma_{LM}\right]. \end{aligned} \tag{8}$$

In the model, the competitive structure of the market is captured by the β terms. This term measures the elasticity of the demand for loans and the elasticity of deposits supply. Therefore, the less elastic the demand for credit, the less will be the value of β and the bank will be able to apply a higher margin if it exercises monopoly power. Hannan and Berger (1991) summarize these arguments in literature on the structure-conduct-performance hypothesis (SCP), which asserts that higher market concentration leads to less favorable pricing to consumers due to some form of collusion among banks. That is, the interest income earned on loans are generally higher for institutions that have a larger share of the market, while interest expenses tend to be lower for these institutions.

The Maudos and Fernández (2004) model yielded an additional term, which captures the average operating costs of banks in the determination of interest spreads. Firms that incur high unit costs will logically need to work with higher margins to enable them to cover their higher operating costs.

Another conclusion from the Maudos and Fernández (2004) model was that spreads are affected by the volatility of money market rates, σ_M^2 in equation (8). That is, the more volatile the rates in the money market, the greater will be the market risk, which will therefore cause banks to want to operate with a higher premium for this uncertainty. From most of the empirical literature on bank spreads, the relationship between spreads and interest rate risk is statistically significant.

Credit risk in the Maudos and Fernández (2004) model is captured by σ_L^2 in equation (8), which is defined as the risk associated with the volatility of the expected return on loans. This was included on the basis that the probability of borrowers defaulting on loans as well as the possibility of a loss of capital and interest, will likely result in a premium charged to cover the likelihood of a default. The interaction of credit and market risk, which is also a measure of default probability, was brought out in the model as having a meaningful role in the determination of bank spreads.

4. DESCRIPTION OF VARIABLES

A number of variables was employed in assessing the response of bank spreads to policy rates. The policy rate variable is proxied by the 180-day BOJ open market operation (OMO) rate, which has a strong influence on market rates given that it serves as signal rate to market participants.[8] Proxies for the variable used to capture the theoretical model on banks' price setting behavior cover the market structure, market risk and credit risk as well as the interaction between credit and market risk and operating costs.

4.1. Market structure

In attempting to capture the market structure based on the theoretical model, two alternative measures were selected. As

[8] Other rates were considered as a proxy for policy rates such as, the three-month money market rate (Gropp, Sorensen and Lichtenberger, 2007), as well as the overnight rate, this is the interest rate at which major financial institutions borrow and lend one-day (or *overnight*) funds among themselves. The Bank sets a target level for this rate. However, in a study of the lead lag structure of interest rates in Jamaica (McLeod, 2008) discovered that the 180-day t-bill rates was used more than any other BOJ rates in the pricing of private rates. Moreover, the study revealed that the 180-day rate had more influence on the market and was viewed as the signal rate by market participants.

a proxy for market structure, the Lerner Index, which measures the degree of competition in the sector was used. The Lerner index is measured as the difference between the price of output (asset) and marginal cost as a share of the price of the asset (see equation 11). The price of asset is computed as total revenues divided by total assets.

$$LI_i = \frac{p_i - MC_i}{p_i}. \tag{11}$$

The marginal cost is based on the estimation of the cost function:

$$\begin{aligned} \ln(TC_i) &= \alpha_0 + \alpha_1 \ln A_i + \frac{1}{2}\alpha_k (\ln A_i)^2 + \sum_{j=1}^{3} \beta_j \ln w_{ji} + \frac{1}{2}\sum_{j=1}^{3}\sum_{k=1}^{3} \beta_{jk} \ln w_{ji} \ln w_{ki} \\ &+ \frac{1}{2}\sum_{j=1}^{3} \gamma_j \ln A_i \ln w_{ji} + \mu_1 trend + \mu_2 \frac{1}{2} trend^2 + \mu_3 trend \ln A_i \\ &+ \sum_{j=1}^{3} \lambda_j trend \ln wji + \ln u_i, \end{aligned} \tag{12}$$

where TC_i denotes total costs, A_i represents total assets, where $i = 1...14$ is the number of institutions in the sector. On the other hand, w_j is the price of the factors of production, and jk is the cross-product of the price of input, $\forall j, k = 1...3$, where: $w_1 =$ price of labor: personnel costs/total assets; $w_2 =$ price of physical capital: operating costs/fixed assets; and $w_3 =$ price of deposits: financial costs/deposits.

The cost function is estimated by including fixed effects for individual banks to capture the influence of variables specific to each bank. A trend component is used to capture technical change and shifts in the cost function over time. As usual, the estimation is done under the restrictions of symmetry and homogeneity in the prices of inputs.

The estimated coefficients of the cost function are then used to compute the marginal cost. The marginal cost can be expressed as:

$$MC = \frac{TC_i}{A_i} \cdot \frac{d \ln TC_i}{d \ln A_i}, \tag{13}$$

where the derivative of the logarithm of the total cost with respect to the logarithm of output is computed using the cost function specified in equation (12). Equation (14), shows the derivative of the cost function in equation (12) with respect to total assets:

$$\frac{d \ln TC_i}{d \ln A_i} = \alpha_j + \alpha_k \cdot \ln A_i + \frac{1}{2}\sum_{j=1}^{3} \gamma_j \cdot \ln w_{ji} + \mu_3 \cdot trend \,. \tag{14}$$

4.2. Market risk

From the theoretical model, the volatility in market interest rates causes uncertainty in the money markets. As such, in proxying for this variable in the empirical model, the monthly standard deviation in the 180-day Treasury bill (t-bill) rates is used.[9]

4.3. Credit risk

In this study, credit risk is measured as the ratio of non-performing loans to total loans. This variable is a measure of the willingness and ability of borrowers to repay their loans.[10]

4.4. Operating cost

Equation (8) reflects the importance of operating costs and quality of management in the price setting behavior of banks. As such, both of these variables are captured by estimating a cost efficiency measure based on the translogorithmic cost function specified in equation (15):

$$\begin{aligned} \ln tc = {} & \alpha_0 + \sum_{i=1}^{2} \alpha_i \ln(y_i) + \sum_{j=1}^{3} \beta_j \ln(p_j) + 1/2 \sum_{i=1}^{2}\sum_{k=1}^{2} \alpha_{ik} \ln(y_i)\ln(y_k) + \\ & + 1/2 \sum_{j=1}^{3}\sum_{h=1}^{3} \beta_{jh} \ln(p_j)\ln(p_h) + \sum_{i=1}^{2}\sum_{j=1}^{3} \delta_{ij} \ln(y_i)\ln(p_j) + \in, \end{aligned} \tag{15}$$

where tc is total operating and interest costs, y_1 is total loans,

[9] This information was taken from Bloomberg as well as Jamaica Money Market Brokers (JMMB), one of the largest stockbrokers and securities dealers in Jamaica. JMMB is also considered by many to be one of the most active players in the money market and has been collecting information on GOJ bond yields from 1999 for the client purposes.

[10] Other variable were considered such as, the slope of the yield curve and was calculated as the difference in five-year government bond yields and three-month interbank deposit rates.

y_2 is all other earning assets, and p_1, p_2, p_3 are the respective prices of labor, capital and borrowed funds. It is assumed that a higher quality of management translates into a profitable composition of assets and a low cost composition of liabilities.[11] As a result, the cost of doing business would be captured as well as the efficiency of management.

4.5. Interaction between credit risk and market risk

The interaction between credit risk and market risk is proxied as the product of the measures of credit risk and interest rate risk.

5. ECONOMETRIC FRAMEWORK

The paper employs a single-stage approach to assess the adjustment of bank spreads to changes in monetary policy, similar to what was employed by McShane and Sharp (1985) and Gropp, Sorensen and Litchtenberger (2007).[12] The model is expressed as:

$$S_{it} = \varphi_0 + \varphi_t PR_t + \sum_{i=1}^{t} \varphi_b X_{bit} + \sum_{i=1}^{t} \varphi_c X_{cit} + v_b + \varepsilon_{it}\,, \tag{16}$$

where $\varepsilon_{it} \sim$ i.i.d and S_{it} represents the spread of bank products $i = 1,\ldots,N$ (savings deposits, time deposits, and the different types of loans) in period $t = 1,\ldots,T$. Policy rate, PR represents the official rate of the central bank and is used to indicate policy direction at a particular time, t. The variable X_{bit} represents the determinants of bank spreads used in the study, while X_{cit} are a set of bank specific control variables.

To facilitate a robust test of the dynamic adjustment of bank spreads, *S*, in response to the level of the policy rate and permit a better identification of the model, equation (16) is

[11] For further discussion on the estimation of technical inefficiencies using the translogorithmic cost function in equation (15) above see Bailey (2007).

[12] In this single step method variables captured in the theoretical model were incorporated as well as an additional variable capturing movement in policy rate.

estimated in first differences and is represented in equation (17):[13]

$$\Delta S_{it} = \varphi_0 + \varphi_1 \Delta PR_{it} + \sum_{i=1}^{t} \varphi_b \Delta X_{bit} + \sum_{i=1}^{t} \varphi_c \Delta X_{it} + v_b + \varepsilon_{it} \ , \tag{17}$$

where Δ denotes first differences and ΔPR_t represents the innovation of the policy rate in period *t*. The innovation in policy rate is accomplished by taking the first difference of a 180-day OMO rate, which would mean considering the expected and the unexpected component of monetary policy.[14] One caveat of estimating the model in first differences is that this would result in an elimination of structural control variables, leaving only cyclical and other time-varying variables as controls.[15]

In assessing the dynamic adjustment of bank spreads to policy rates, the framework is refined to include asymmetries in the adjustment process as well as the movement in spreads across different bank products (see equation 17a). Given that bank products may exhibit varying adjustment dynamics to policy rates, an additional estimation is conducted to capture asymmetries in the adjustment in bank spreads.[16] Based on equation (17a), when the indicator variable I^{up} is equal to one, this translates to a tightening in monetary policy.

$$\begin{aligned} \Delta S_{it} &= \varphi_0 + I^{up} * \sum_i \varphi_i \Delta PR_{it} + (1 - I^{up}) * \sum_i \varphi_i \Delta PR_t + \\ &\quad + \sum_{i=1}^{t} \varphi_b \Delta X_{bit} + \sum_{i=1}^{t} \varphi_c \Delta X_{cit} + v_b + \varepsilon_{it} . \end{aligned} \tag{17a}$$

As such, this specification allows for different dynamics

[13] See Gropp, Sorensen and Litchtenberger (2007).

[14] In this context, one would say that the difference between an expected and an unexpected monetary policy is that the former is well communicated to the market.

[15] One could argue that by first differencing the bank specific effects would disappear as well. However, the equation is estimated in differences given that even in first differences there may be unobserved bank specific factors.

[16] This was done to determine whether a downward change in the policy rates results in a slower adjustment in loan rates and an upward change in the policy rate would result in a faster adjustment in loan rates.

based on the direction of the policy change. In this context, the framework is useful in ascertaining whether a downward change in the policy rate results in a slower adjustment for loan rates compared to deposit rates and whether an upward change in the policy rate results in a faster adjustment for loan rates.

6. DATA AND DESCRIPTIVE STATISTICS

The paper employs quarterly data for the period March 1996 to June 2008. Spreads are computed on three types of loans including personal credit, instalment and mortgage credit as well as four types of deposits, namely, demand, savings, short- and long-term time deposits.

It is found that policy rates (PR_t) as well as variables capturing interest rate risk, credit risk, the interaction between credit and market risk and efficiency indicators exhibit positive skewness and a peaked distribution (see table A.2, Appendix A). This means that policy rates exhibit leptokurtic behavior, which is typical of interest rate data. Positive skewness is an indication that the probability of observing a large positive jump usually exceeds the probability of observing a large negative jump in policy rates during the sample period.

7. RESULTS

The bank spread equations are estimated in first differences with the introduction of fixed effects. The results from the baseline model *(model 1)* for commercial banks, merchant banks and building societies (see tables 1, 2 and 3) show that at the 1% level of significance, current changes in loan spreads are negatively related to changes in the 180-day money market rate, while the opposite is true for deposit spreads.[17] That is, in the current period, when policy rate changes are made, whether upwards or downwards, banks are slow to react to these changes, hence there is a narrowing in

[17] This based on the assumption that there is almost a seamless pass-through from policy rate changes to market rates.

TABLE 1. ESTIMATION RESULTS: BASELINE MODEL COMMERCIAL BANKS

	Model 1 *Bank fixed effects*			*Model 2* *Bank fixed effects*		
	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*
Policy rate						
Loans	–0.69[c]	–0.138	0.31			
	(0.06)	(0.05)				
Deposits	0.881[c]	–0.012	0.13			
	(0.06)	(0.05)				
Personal				–0.771[c]	–0.11	0.23
				(0.07)	(0.07)	
Installment				–0.19[b]	–0.26[c]	0.56
				(0.07)	(0.07)	
Commercial				–0.69[c]	–0.10	0.31
				(0.07)	(0.07)	
Savings				0.91[c]	0.04	0.09
				(0.07)	(0.07)	
Time deposit (st)				0.90[c]	–0.19[b]	0.30
				(0.07)	(0.07)	
Time deposit (lt)				0.88[c]	–0.07	0.12
				(0.07)	(0.07)	
Bank soundness	–0.239			–1.196		
	(0.95)			(0.79)		
Credit risk	0.707			3.378		
	(2.51)			(2.08)		
Interest rate risk	–0.38[b]			–0.11		
	(0.13)			(0.17)		
Competition	–1.008[a]			–2.821[a]		
	(2.34)			(1.95)		
Efficiency	–0.168			–1.002		
	(2.24)			(1.86)		
Observations	576			1,728		
Wald statistic	59.82[c]			46.86[c]		
R^2	0.49			0.31		

NOTES: Models were estimated using fixed-effects across banks. Standard errors are in parenthesis. [a, b, c] indicates significance at the 10%, 5% and 1% level, respectively. The column *pass-through* reports the share of changes in bank rates after two quarters to the change in the policy rate.

the spreads. Gropp et al. (2007) argue that if there had been a swift pass-through, changes in the market rate would fully reflect changes in bank rates, thus leaving the spread unchanged. Second, if bank rates adjust fully to changes in market rates after a lag then we would expect the sum of the response to current and lagged changes to be equal to zero.

In contrast to results by Gropp et al. (2007), it is determined that in the Jamaican commercial banking sector, deposit and lending spreads only react to a temporary shock to money market rates in the current period, as the lagged changes are largely insignificant.[18]

While bank retail rates adjust sluggishly for both loans and deposits, the pass-through is more complete for lending rates than for deposit rates. Commercial banks' lending spreads are estimated to decrease by, on average, around 69 basis points (bps) following an increase of 100 bps in market rates in the same quarter, indicating that lending rates would increase by 32 bps. In the merchant banking sector, the results suggest that a complete pass-through in lending rates is attained after two quarters (see table 2).[19] In addition, an assumed shock of 100 bps in market rates among building societies would cause only 0.02 bps increase in their lending rates (see table 3).

On average, commercial banks' deposit spreads increase by 88 bps following an increase of 100 basis points in market rates, suggesting that deposit rates increase by only 12 bps after one quarter. In contrast, deposit spreads in the merchant banking sector are estimated to increase by, on average, 71 bps following an increase of 100 bps in market rates in the same

[18] Gropp et al. (2007) found in a similar study, that when lending rates adjust with a lag to a given *one off* change in market rates, for example an increase, they would expect to observe a decrease in the spreads this period (as bank rates adjust upwards more slowly). That is, a negative relationship between the change in the market rate and the change in spread. As lending rates eventually rise there is, however, a positive relationship between bank spreads and the lagged change in the market rate. Conversely, they found that deposit spreads are positively related to current changes in market rates and negatively related to the lagged change in market rates.

[19] Anecdotal evidence suggest that the products and services being offered in this sector as well as competition from the other sectors would play a significant role in the rate of pass-through from money market rates to retail rates.

TABLE 2. ESTIMATION RESULTS: BASELINE MODEL MERCHANT BANKS

	Model 1 Bank fixed effects			*Model 2 Bank fixed effects*		
	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*
Policy rate						
Loans	–1.1371[c]	1.1906[c]	1.05			
	(0.16)	(0.14)				
Deposits	0.7142[c]	–0.729[c]	1.01			
	(0.16)	(0.14)				
Personal				–0.98[c]	0.97[c]	0.99
				(0.14)	(0.12)	
Commercial				–0.94[c]	1.00[c]	1.07
				(0.14)	(0.12)	
Time deposit (st)				0.71[c]	–0.67[c]	0.96
				(0.14)	(0.12)	
Time deposit (lt)				0.63[c]	–0.73[c]	1.09
				(0.14)	(0.12)	
Bank soundness	–3.00			–1.156		
	(2.24)			(1.47)		
Credit risk	2.1385			–1.279		
	(3.25)			(2.14)		
Interest rate risk	0.5453[b]			0.295[c]		
	(0.28)			(0.18)		
Competition	0.6995			0.485		
	(1.78)			(1.17)		
Efficiency	0.7366			–1.817		
	(2.02)			(1.33)		
Observations	392			784		
Wald statistic	12.83[c]			17.18[c]		
R^2	0.25			0.23		

NOTES: Models were estimated using fixed-effects across banks. Standard errors are in parenthesis. [a, b, c] indicates significance at the 10%, 5% and 1% level, respectively. The column *pass-through* reports the share of changes in bank rates after two quarters to the change in the policy rate.

period (suggesting that deposit rates increase by only 29 bps), but decrease by, on average, 72 bps in response to the lagged increase of 100 bps in market rates. The combined impact thus indicates that an increase of market rates by 100 bps results in an upward adjustment of deposit rates after two quarters by approximately 100%, further indicating that there is full pass-through in this sector after two quarters. The building societies, on the other hand, display a more sluggish pass-through in their deposit rates. The results indicate that a 100 bps increase in market rates would cause deposit spreads to increase by 87 bps. As such, deposit rates increase by roughly 10 bps after one quarter.

The control variables, namely, bank soundness, credit risk, interest rate risk, competition and efficiency, are largely

TABLE 3. ESTIMATION RESULTS: BASELINE MODEL BUILDING SCOCIETIES

	Model 1 Bank fixed effects			*Model 2 Bank fixed effects*		
	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*
Policy rate						
Loans	–0.8546[c]	–0.126[b]	0.02			
	(0.16)	(0.14)				
Deposits	0.8794[c]	0.0162	0.10			
	(0.16)	(0.14)				
Mortgage				–0.964[c]	0.03	0.06
				(0.06)	(0.06)	
Savings				0.824[c]	0.12[a]	0.05
				(0.06)	(0.06)	
Time deposit (st)				0.82[c]	0.00	0.18
				(0.06)	(0.06)	
Time deposit (lt)				0.80[c]	0.10[c]	0.10
				(0.06)	(0.06)	
Bank soundness	–0.61			–1.808		
	(2.16)			(1.73)		
Credit risk	–9.995			–3.508		
	(8.08)			(6.45)		
Interest rate risk	–0.0266			0.235[a]		
	(0.18)			(0.14)		
Competition	–21.916[b]			0.124		
	(9.50)			(2.29)		
Efficiency	–0.4339			–15.12[a]		
	(2.86)			(7.59)		
Observations	384			784		
Wald statistic	58.27[c]			64.63[c]		
R^2	0.61			0.55		

NOTES: Models were estimated using fixed-effects across banks. Standard errors are in parenthesis. [a, b, c] indicates significance at the 10%, 5% and 1% level, respectively. The column *pass-through* reports the share of changes in bank rates after two quarters to the change in the policy rate.

insignificant across all sectors.[20] An increase in the interest rate risk facing commercial banks, as measured by the change in the monthly standard deviation of the 180-day T-bills, has a negative impact on bank spreads. This result implies that commercial banks facing higher uncertainty regarding interest rate developments tend to operate with lower spreads relative to market rates. This result largely reflects commercial banks ability to access to cheap funds, which enables them to absorb the costs, associated with this higher risk without charging higher premiums. Further results indicate that competition plays an important role in the commercial bank's pricing mechanism, as reflected by the significance of the Lerner index for competition in the model. However, in the merchant banking sector, the sign on the interest rate risk variable is positive indicating that merchant banks facing higher uncertainty regarding interest rate developments tend to operate with higher spreads relative to market rates. Intuitively, uncertainties faced by the merchant banks regarding market interest rate developments is likely to be transferred to the consumers in the form of higher premiums.

Finally, with respect to the building societies sector, the results show that changes in bank spreads are negatively related to competition. That is, as the building societies sector becomes more competitive, bank spreads are likely to fall in line with market rates.

In order to assess how individual bank products react to changes in market rates, *model 2* shows the disaggregation across different products for all sectors. The results are considerably different in some cases depending on the loan and deposit categories, as well as depending on the final pass-through after two quarters. For loans in the commercial banking sector, the pass-through is sluggish, except for instalment credit, which shows that after two quarters, the pass-through would approximate 56 bps after a 100 bps increase in market rates. Consistent with a priori expectations, loans in the building societies sector have a similar sluggish pass-through given that their loan portfolio is highly dominated by mortgage-related loans. Loans in the merchant banking sector, across all categories, have full pass-through after two quarters.

[20] Henceforth, only those variables that are significant will be discussed.

Across the sectors, the pass-through in rates is generally more sluggish and significantly less complete for deposit rates relative to loan rates. However, for short-term time deposits, the pass-through amount to 30 bps, 96 bps and 18 bps for commercial banks, merchant banks and building societies. These figures indicate a swift pass-through relative to the other deposit segments that have an average pass-through of 0.06 bps across all sectors.

7.1. Extension of model

In order to investigate whether the pass-through is asymmetric, equation (17a) was estimated with different slopes for periods when market rates increased and when they decreased across all three banking sectors.[21] According to Gropp et al. (2007), the pass-through to retail rates could be asymmetric if the price elasticity of demand is low or if competition is less than perfect. As such, banks would adjust loan rates more quickly when interest rates are increasing than when they are decreasing and viceversa for deposit rates.

The results obtained suggest that there is some evidence of asymmetry in the pass-through. Model 3 shows that in the case of commercial banks, while loan rates adjusted upwards quickly in response to market rate increases, the same can be said of loan rates when market rates adjust downwards.[22] The results, for building societies were largely similar to those of the commercial banks for the parsimonious model (see Appendix A, table A.5). For the merchant banks, the results indicate that loan rates adjusted faster and more completely when rates adjusted upwards than when they were moving downwards. Conversely, deposit rates tended to adjust more completely after two quarters when interest rates were declining (see Appendix A, table A.4).

The product specific effects of the parsimonious model indicate that rates on personal, commercial and instalment

[21] Over the sample period the 180-day money market rate (PR_t) increased approximately 56% of the total number of quarters.

[22] In the case of the deposit rates, the movements were largely in accordance with those of Hannan Berger (1991) and Gropp et al. (2007) in which they found that deposit rates tended to adjust faster and more completely after two quarters when interest rates are declining.

loans were insensitive to declines in market rates.

On the other hand, savings and time deposits rates adjust more quickly and completely when market rates adjusted downwards, which is consistent with the findings of Gropp (2007). For the merchant banks and the building societies, the results were uni-directional for loan rates (mortgages, personal, commercial, and instalment) as the statistical test indicated that when market rates adjusted downwards, there were minimal movements in loan rates.

TABLE 4. ESTIMATION RESULTS: BASELINE MODEL COMMERCIAL BANKS

	Model 3 Asymmetry			*Model 4 Asymmetry*		
	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*
Loans						
Up	-0.187^{c} (0.02)	0.0188^{b} (0.02)	0.83			
Down	0.234^{c} (0.04)	-0.26^{c} (0.04)	0.97			
Deposits						
Up	0.157^{c} (0.02)	–0.012 (0.02)	0.84			
Down	-0.179^{c} (0.04)	0.1829^{c} (0.04)	1.00			
Personal						
Up				-0.202^{c} (0.03)	0.0215 (0.03)	0.80
Down				0.08 (0.05)	–0.197 (0.05)	0.88
Commercial						
Up				-0.172^{c} (0.03)	0.02^{c} (0.03)	0.83
Down				0.15^{b} (0.05)	–0.226 (0.05)	0.93
Installment						
Up				-0.083^{c} (0.03)	0.0757^{c} (0.03)	0.99
Down				-0.10^{a} (0.05)	–0.063 (0.05)	1.04
Savings						
Up				0.167^{c} (0.03)	0.003 (0.03)	0.83

Down		-0.21^{c} (0.05)	0.2039^{c} (0.05)	1.21
Time deposit (st)				
Up		0.13^{c} (0.03)	-0.047^{b} (0.03)	0.91
Down		-0.18^{c} (0.05)	0.2814^{c} (0.05)	1.18
Time deposit (lt)				
Up		0.1491^{c} (0.03)	-0.022 (0.03)	0.85
Down		-0.156^{c} (0.05)	0.2387^{c} (0.05)	1.16
Observations	576	1,728		
Wald statistic	38.89^{c}	20.26^{c}		
R^2	0.22	0.14		

NOTES: Models were estimated using fixed-effects across banks. Standard errors are in parenthesis. [a, b, c] indicates significance at the 10%, 5% and 1% level, respectively. The column *pass-through* reports the share of changes in bank rates after two quarters to the change in the policy rate.

8. ROBUSTNESS CHECKS

To permit robustness checks, the baseline models for the sectors were estimated under different conditions to ensure consistency under different specifications (see Appendix). The models were estimated with and without fixed effects as well as with random effects across sectors and product segments (see Appendix A, tables A.3, A.4 and A.5; model R1). Furthermore, the models were estimated using a seemingly unrelated regression (SUR) (see Appendix A, tables A.3, A.4 and A.5; model R2). The results obtained with these alternative specifications were essentially the same as with the results obtained with our baseline *models 1* and *2*.

9. CONCLUSION

It is a well-known feature of monetary policy operations that central banks aim to exercise control over short-term interest rates by adjusting the official rate. Moreover, it is also commonly assumed that there is complete transmission to short-term rates within a short period. Furthermore, studies on bank spreads are crucial given that with complete pass-through monetary policy can be more efficient in its ability to control inflation.

The results of this study are generally consistent with the empirical literature on pass-through and bank spreads. It was determined that bank spreads tended to adjust very slowly to official policy rate changes. The findings may suggest that the stickiness of deposit spreads largely reflect the fact that banks exert a moderate degree of market power in the market for retail products. The results also showed that there are significant differences in the adjustment processes for the different categories of loan and deposit products. The rates on saving deposits displayed a high degree of rigidity and, as a result, reactions to changes in market rates were almost non-existent.

Findings from this study also suggest that commercial banks hold a fair degree of market power in the market for loans and deposits due to their dominance in the banking sector. As such, there should be a concerted effort to enhancing the competitive environment for banks by encouraging the availability of alternative capital market-based instruments for financing investment in order to increase access to financing (e.g., for small and medium size enterprises). This can be done by promoting innovation in the non-bank financial sector.

In addition, the results provide evidence of asymmetry in the pass-through process, as banks tend to adjust loan rates more quickly in relation to changes in policy rates when rates are increasing than when they are declining, while the opposite holds for deposit rates. Additionally, results from the study indicates that if banks' loan portfolio comprises largely insensitive assets then monetary policy would be less effective under such conditions and vice versa.

The findings of Maudos and Fernández (2004), and Gropp et al. (2007) suggest the potential benefits to be gained from enhanced risk management practices. Strengthened risk management practices enables banks to charge lower premia, which will result in lower spreads, thus amplifying the effects of monetary policy changes on bank interest rates. However, we find that in Jamaica risk premia may not have such a significant impact on banks' price setting behavior.

Appendix A

Taking into account equation (5), the expected utility of the bank is:[23]

(A.1) $$EU(W) = U(\bar{W}) + U'(\bar{W})E(L_0Z_L + M_0Z_M) + \frac{1}{2}U''(\bar{W})E(L_0Z_L + M_0Z_M)^2$$

$$= U(\bar{W}) + \frac{1}{2}U''(\bar{W})(L_0^2\sigma_L^2 + M_0^2\sigma_M^2 + 2L_0M_0\sigma_{LM}).$$

When a new deposit, D, is made, the banking firm has to pay $r_D D$ and operating costs $C(D)$, and will obtain a return $(r + Z_M)D$ in the money market. In this way, the bank's final wealth will be:

(A.2) $$\begin{aligned} W_T &= (1 + r_I + Z_I)I_0 - (1 + r_D)D + (1 + r + Z_M)M_0 + \\ &\quad + (1 + r + Z_M)D - C(I_0) - C(D) \\ &= W_0(1 + r_w) + L_0Z_L + aD + (M_0 + D)Z_M - C(I_0) = C(D), \end{aligned}$$

and expected utility after the new deposit has been made is given by the following expression:

(A.3) $$\begin{aligned} EU(W_T) &= U(\bar{W})E(W - \bar{W}) + \frac{1}{2}U''(\bar{W})E(W - \bar{W})^2 \\ &= U(W) + U'(\bar{W})\left[aD - C(D)\right] \\ &\quad + \frac{1}{2}U''(\bar{W})\Big[(aD - C(D))^2 + L_0\sigma_L^2 + (M_0 + D)\sigma_M^2 + \\ &\quad + 2L_0(M_0 + D)\sigma_{LM}\Big]. \end{aligned}$$

Given the level of wealth after the arrival of the new deposit, the increase in expected utility is as follows:

(A.4) $$\begin{aligned} \Delta EU(W_D) &= EU(W_T) - EU(W) \\ &= U'(\bar{W})\left[aD - C(D)\right] + \\ &\quad + \frac{1}{2}U''(\bar{W})\begin{bmatrix} \left(aD - C(D)\right)^2 \\ +(D + 2M_0)D\sigma_M^2 + 2L_0D\sigma_{LM} \end{bmatrix}. \end{aligned}$$

[23] Given by $\bar{W} = E(W) = E(W_0(1 + r_w) + L_0Z_L + M_0Z_M - C(I_0) = W_0(1 + r_w) - -C(I_0)$.

In the same way, if the bank grants a new credit for an amount L it will receive an income $r_L L = (r + b + Z_L)L$, and incur operating costs $C(L)$and costs of financing the granting of credits $(r + Z_M L)$.

Analogously to the receiving of deposits, the increase of the bank's expected utility due to the granting of an additional credit will be:

$$\text{(A.5)}\quad \Delta EU(W_T) = EU(W_T) - EU(W)$$
$$= U'(\bar{W})\left[bL - C(L)\right] +$$
$$+\frac{1}{2}U''(\bar{W})\begin{bmatrix}\left(bL - C(L)\right)^2 + (L + 2L_0)L\sigma_L^2 \\ +(L - 2M_0)L\sigma_M^2 + 2(M_0 - L_0 - L)L\sigma_{LM}\end{bmatrix}.$$

Bearing in mind the probabilities of granting credits or capturing deposits reflected in equation (8), the problem of maximization of (9) can be written:

$$Max_{a,b}EU(\Delta W) = (\alpha_D - \beta_D a)\begin{bmatrix} U'(\bar{W})\left[aD - C(D)\right] \\ +\frac{1}{2}U''(\bar{W})\begin{bmatrix}(aD - C(D))^2 \\ +(D + 2M_0)D\sigma_M^2 + 2L_0 D\sigma_{LC}\end{bmatrix}\end{bmatrix}$$
$$+ (\alpha_L - \beta_L b)\begin{bmatrix} U'(\bar{W})\left[bL - C(L)\right] \\ +\frac{1}{2}U''(\bar{W})\begin{bmatrix}(bL - C(L))^2 + (L + 2L_0)L\sigma_L^2 \\ +(L - 2M_0)L\sigma_M^2 + 2(M_0 - L_0 - L)L\sigma_{LM}\end{bmatrix}\end{bmatrix}$$

The first-order conditions with respect to a and b is as follows:

$$a = \frac{1}{2}\frac{\alpha_D}{\beta_D} + \frac{1}{2}\frac{C(D)}{D} - \frac{1}{4}\frac{U''(\bar{W})}{U'(\bar{W})}\left[(D + 2M_0)\alpha_M^2 + 2L_0\alpha_{LM}\right],$$

$$b = \frac{1}{2}\frac{\alpha_L}{\beta_L} + \frac{1}{2}\frac{C(L)}{L} - \frac{1}{4}\frac{U''(\bar{W})}{U'(\bar{W})}\left[(L + 2L_0)\alpha_L^2 + (L - 2M_0)\alpha_M^2 + 2(M_0 - L)\alpha_{LM}\right].$$

The first order condition with respect to a and b give rise to the optimal spreads in equation (8):[24]

[24] It is assumed, following Ho and Saunders (1981) and subsequent extensions that the second-order terms of the margins and costs of the Taylor's expansion of expressions (6) and (7) are negligible.

TABLE A.1. DEPOSIT TAKING INSTITUTIONS: 1996-2008

Institution name	*Abbr. name*	*Previous name*
Bank of Nova Scotia	BNS	None
National Commercial Bank	NCB	None
Royal Bank of Trinidad and Tobago	RBTT	Union Bank of Jamaica UBJ, now RBIT was a result of the transfer of assets and liabilities of six (6) financial institutions to Citizens Bank. The amalgamated entities were Citizens Merchant Bank Ltd., Corporate Merchant Bank, Island Life Merchant Bank. Workers Savings and Loan Bank, Island Victoria Bank and Eagle Commercial Bank.
First Caribbean International, Bank Jamaica Ltd.	FCIBJ	Canadian Imperial Bank of Commerce CIBC. CIBC later became Bank of Commerce Jamaica Ltd. On November 12, 1975, the bank was incorporated as CIBC West Indies Holdings Limited (incorporated in Barbados) purchased CIBC's 55.2 percent share in CIBC Jamaica Ltd. On a share exchange basis. The metamorphosis continued on October 30, 2002 when the bank was incorporated locally as First Caribbean International Bank (FCIB) Jamaica Ltd. FCIB is currently an amalgamation of the retail, corporate and offshore banking operations of CIBC West Indies Holding Ltd. And Barclays Bank, PLC in the Caribbean, its majority shareholders
First Global Bank	FGB	FGB was formerly known as First Jamaica National Bank (FJNB) Ltd. In December 1992, Trafalgar Development Bank acquired GJNB from Jamaica National Building Society. The institution was renamed Trafalgar Commercial Bank (TCB) on the 26 of June 1993. As part of a rebranding exercise, TCB had its name changed to First Global Bank Limited, with effect from 11 December 2001.
Citibank		
FIAS		
Capital and Credit Merchant Bank	CCMB	None
Citi Merchant Bank Ltd.	CITIMER	
DB&G Merchant Bank Ltd.	DR&G	DB&G formerly Billy Craig merged the assets and liabilities of Issa Trust and Merchant Bank, in August 2003.
MF&G Trust & Finance	MF&G	None
Pan-Caribben Merchant Bank	PCMB	PCMB is the outcome of the merger of the assets and liabilities of Pan Caribbean Merchant Bank and Manufacturers Sigma Merchant Bank MSMB on June 1, 2004. MSMB itself was the outcome of the merger of Manufacturers Merchant Bank (MMB) and Sigma Management Systems (SIGMA) Ltd.

TABLE A.1 *(finish)*

Institution name	*Abbr. name*	*Previous name*
Building Sicieties		None
First Caribbean International Building Society	FCIBS	FCIBS today is a result of the rebranding of CIBC Building Society following the merger of its retail, corporate and offshore banking operations of CIBC and Barclays Bank PLC in the Caribbean on October 30, 2002
Jamaica National Building Society	JNBS	None
Scotia Building Society	SJBS	None
Victoria Building Society	VMBS	None

TABLE A.2. DESCRIPTIVE STATISTICS

	Policy rate	*Interest risk*	*Credit risk*	*Credit × interest*	*Lindex*	*Efficiency*
Mean	19.23	0.93	0.08	0.08	0.56	1.45
Standard error	0.31	0.05	0.01	0.01	0.03	0.04
Median	16.08	0.72	0.04	0.02	0.58	1.12
Mode	12.00	1.37	–	–	–	2.04
Standard deviation	8.10	1.33	0.14	0.18	0.82	1.00
Sample variance	65.64	1.78	0.02	0.03	0.66	1.00
Kurtosis	3.12	20.64	41.87	29.93	64.00	38.81
Skewness	1.76	4.16	5.20	4.85	–6.52	5.38
Range	35.58	8.65	1.78	1.76	13.85	10.87
Minimum	12.00	0.02	–	–	–9.76	1.00
Maximum	47.58	8.67	1.78	1.76	4.09	11.87
Sum	13,458.17	650.57	57.44	59.23	393.41	1,016.87
Count	700	700	700	700	700	700

TABLE A.3. ESTIMATION RESULTS: BASELINE MODEL COMMERCIAL BANKS

	Model R1 No effects plus SUR			*Model R2 No effects plus SUR*		
	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*
Policy rate						
Loans	–0.673[c]	–0.118[b]	0.21			
	(0.06)	(0.05)				
Deposits	–0.867[c]	–0.027	0.16			
	(0.06)	(0.05)				
Personal				–0.795[c]	–0.05	0.21
				(0.07)	(0.07)	
Installment				–0.21[b]	–0.20[c]	0.59
				(0.07)	(0.07)	
Commercial				–0.72[c]	–0.04	0.25
				(0.07)	(0.07)	
Savings				0.86[c]	–0.07	0.14
				(0.07)	(0.07)	
Time deposit (st)				0.85[c]	–0.16[b]	0.32
				(0.07)	(0.07)	
Time deposit (lt)				0.83[c]	–0.043	0.17
				(0.07)	(0.07)	
Bank soundness	–0.379			–2.434		
	(0.13)			(1.03)		
Credit risk	0.832			2.109		
	(2.53)			(1.22)		
Interest rate risk	–0.379[b]			–0.204		
	(0.13)			(1.12)		
Competition	0.982			–0.442		
	(2.37)			(0.39)		
Efficiency	1.181			0.21		
	(1.98)			(0.20)		
Observations	576			1,728		
Wald statistic	58.32[c]			42.57[c]		
R^2	0.49			0.29		

NOTES: Models were estimated using fixed-effects across banks. Standard errors are in parenthesis. [a, b, c] indicates significance at the 10%, 5% and 1% level, respectively. The column *pass-through* reports the share of changes in bank rates after two quarters to the change in the policy rate.

TABLE A.4. ESTIMATION RESULTS: BASELINE MODEL MERCHANT BANKS

	Model R1 *No effects plus SUR*			*Model R2* *No effects plus SUR*		
	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*
Policy rate						
Loans	-1.1264^{c} (0.17)	1.1109^{c} (0.16)	0.98			
Deposits	0.8059^{c} (0.08)	-0.799^{c} (0.08)	0.99			
Personal				-0.993^{c} (0.16)	0.97^{c} (0.15)	0.97
Commercial				-1.00^{c} (0.15)	0.99^{c} (0.14)	0.99
Time deposit (st)				0.69^{c} (0.07)	-0.68^{c} (0.07)	0.98
Time deposit (lt)				0.71^{c} (0.09)	-0.72^{c} (0.08)	1.01
Bank soundness	–0.53 (1.17)			0.147 (1.15)		
Credit risk	–1.1174 (1.70)			–2.638 (1.67)		
Interest rate risk	0.26^{b} (0.14)			0.283^{b} (0.14)		
Competition	0.5446 (0.93)			0.473 (0.91)		
Efficiency	–0.2307 (1.03)			–0.954 (1.01)		
Observations	392			784		
Wald statistic	16.74^{c}			18.41^{c}		
R^2	0.28			0.23		

NOTES: Models were estimated using fixed-effects across banks. Standard errors are in parenthesis. $^{a, b, c}$ indicates significance at the 10%, 5% and 1% level, respectively. The column *pass-through* reports the share of changes in bank rates after two quarters to the change in the policy rate.

TABLE A.5. ESTIMATION RESULTS: BASELINE MODEL BUILDING SCOCIETIES

	Model R1 No effects plus SUR			*Model R2 No effects plus SUR*		
	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*
Policy rate						
Loans	–0.8546[c]	–0.126[b]	0.02			
	(0.16)	(0.14)				
Deposits	0.8794[c]	0.0162	0.12			
	(0.16)	(0.14)				
Mortgage				–0.923[c]	0.00	0.08
				(0.08)	(0.07)	
Savings				0.861[c]	0.09[a]	0.05
				(0.05)	(0.05)	
Time deposit (st)				0.85[c]	–0.04	0.15
				(0.05)	(0.04)	
Time deposit (lt)				0.83[c]	0.06	0.17
				(0.06)	(0.06)	
Bank soundness	–1.70			0.656		
	(3.95)			(2.41)		
Credit risk	–1.70			–2.277		
	(3.95)			(4.78)		
Interest rate risk	–0.834			–0.02[a]		
	(0.15)			(0.18)		
Competition	–21.916[b]			–0.218		
	(9.50)			(0.74)		
Efficiency	–0.4339			0.203[a]		
	(2.86)			(0.59)		
Observations	384			784		
Wald statistic	57.75[c]			47.69[c]		
R^2	0.58			0.45		

NOTES: Models were estimated using fixed-effects across banks. Standard errors are in parenthesis. [a, b, c] indicates significance at the 10%, 5% and 1% level, respectively. The column *pass-through* reports the share of changes in bank rates after two quarters to the change in the policy rate.

TABLE A.6. ESTIMATION RESULTS: BASELINE MODEL MERCHANT BANKS

	Model 3 Asymmetry			*Model 4 Asymmetry*		
	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*
Loans						
Up	–0.2323[c]	0.026[c]	0.77			
	(0.05)	(0.05)				
Down	0.1199[a]	–0.28[c]	0.72			
	(0.09)	(0.09)				
Deposits						
Up	0.113	–0.026	0.91			
	(0.05)	(0.05)				
Down	–0.2109[b]	0.2744[b]	0.94			
	(0.09)	(0.09)				
Personal						
Up				–0.227	0.059[b]	1.06
				(0.04)	(0.04)	
Down				0.071	–0.205	1.07
				(0.08)	(0.08)	
Commercial						
Up				–0.204[c]	0.02	0.80
				(0.04)	(0.04)	
Down				0.089	–0.12	1.09
				(0.08)	(0.08)	
Time deposit (st)						
Up				0.104[b]	–0.019	0.90
				(0.04)	(0.04)	
Down				–0.23[b]	–0.23[b]	1.23
				(0.08)	(0.08)	
Time deposit (lt)						
Up				0.096[b]	–0.037	0.90
				(0.04)	(0.04)	
Down				–0.162[a]	–0.162[b]	1.16
				(0.08)	(0.08)	
Observations	392			784		
Wald statistic	16.74[c]			18.41[c]		
R^2	0.28			0.23		

NOTES: Models were estimated using fixed-effects across banks. Standard errors are in parenthesis. [a, b, c] indicates significance at the 10%, 5% and 1% level, respectively. The column *pass-through* reports the share of changes in bank rates after two quarters to the change in the policy rate.

TABLE A.7. ESTIMATION RESULTS: BASELINE MODEL BUILDING SOCIETIES

	Model 3 Asymmetry			*Model 4 Asymmetry*		
	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*	$PR_{(t)}$	$PR_{(t-1)}$	*Pass-through*
Loans						
Up	–0.1446[c]	–0.029	0.86			
	(0.03)	(0.03)				
Down	0.1981[c]	–0.261[c]	0.94			
	(0.05)	(0.05)				
Deposits						
Up	0.1658[c]	–0.025	0.83			
	(0.03)	(0.03)				
Down	–0.1265[b]	0.2318[c]	0.89			
	(0.05)	(0.05)				
Mortgage						
Up				–0.128[c]	–0.006	0.87
				(0.03)	(0.04)	
Down				–0.011	–0.389[c]	0.61
				(0.05)	(0.05)	
Savings						
Up				0.17[c]	0.021	0.83
				(0.03)	(0.03)	
Down				–0.18[b]	0.31[c]	0.87
				(0.05)	(0.05)	
Time deposit (st)						
Up				0.15[c]	–0.029	0.85
				(0.03)	(0.03)	
Down				–0.106[a]	0.27[c]	1.11
				(0.05)	(0.05)	
Time deposit (lt)						
Up				0.154[c]	0.005	0.85
				(0.03)	(0.03)	
Down				–0.068	0.246[c]	1.07
				(0.05)	(0.05)	
Observations	384			784		
Wald statistic	57.75[c]			47.69[c]		
R^2	0.58			0.45		

NOTES: Models were estimated using fixed-effects across banks. Standard errors are in parenthesis. [a, b, c] indicates significance at the 10%, 5% and 1% level, respectively. The column *pass-through* reports the share of changes in bank rates after two quarters to the change in the policy rate.

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Pablo Pincheira

A real time evaluation of the central bank of Chile GDP growth forecasts

1. INTRODUCTION

In this article we evaluate the Central Bank of Chile's annual GDP growth forecasts during the period 1991-2009. We compare the Central Bank of Chile' forecasts (CBCh) with those from the Survey of Professional Forecasters (SPF), the Consensus Forecasts, and also with those obtained using simple time-series models. We evaluate a number of different forecast properties, including forecasts accuracy and efficiency. In particular we place our attention on root mean squared

Paper prepared by P. Pincheira, Central Bank of Chile. I am grateful to Gustavo Leyva and Cristián Muñoz for excellent research assistantship and to Michael Pedersen for sharing the real time database. I also thank Felipe Stanger, María Pilar Pozo and Andrea Contreras for their valuable contribution in an alternative real time database. Helpful and insightful comments from José De Gregorio, Manuel Marfán, Sebastián Claro, Rodrigo Vergara, Pablo García, Ricardo Vicuña, Claudio Soto, Juan Pablo Medina, Andrea Bentancor and Luis Felipe Céspedes have enriched this paper. All errors are responsibility of the author. The views expressed in this paper do not necessarily represent those of the Central Bank of Chile or its Board members. *Correspondence: Agustinas 1180. Santiago-Chile. Tel: 56-2-670-2874, Fax: 56-2-670-2836. E-mail: ⟨ppinchei@bcentral.cl⟩.*

prediction errors (RMSPE) and autocorrelation in forecast errors.

This is not the first written article comparing the CBCh annual GDP growth forecasts. Nevertheless, the main contribution of our article is significant. This is because we make use of a real time database, already used in Pedersen (2009), containing both quasi-final and first releases of annual GDP growth for Chile.[1] For a given date in the past, this database contains the last revisions of GDP growth observations that were actually available at that moment in time.[2] This allows us to properly generate real time forecasts with simple time series models and make a fair comparison of these forecasts with those of the CBCh. This is important because, otherwise, times series forecasts based upon revised data would count with the benefit of revisions, which of course, were not available at the moment of prediction.

It is important to emphasize that during the sample period official GDP growth observations were released in four different reference years. These reference years are 1977, 1986, 1996 and 2003. These multiple changes in reference years induce a missing observations problem in the computation of the CBCh's forecast errors. This happens because, in a few occasions, the CBCh released their forecasts at a moment in time in which the new methodology associated with the new reference year was not yet released.[3] Nevertheless, when official data of that particular year was released, it was expressed in the new methodology corresponding to the new reference year. As a consequence, the CBCh never issued a forecast for that particular year in that particular reference year, so we rather prefer to treat that figure as a missing observation.

[1] Quasi-final releases correspond to the last revision available for a given reference year. Quasi-final releases may or may not coincide with final releases. Actually they are the same with the exception of the following years: 1991, 1992, 1999, 2000, 2004 and 2005. Quasi-final and final data differ in that final observations are never revised in the future. Quasi-final observations may be revised in the future, but if they are revised the revision is expressed in a new reference year.

[2] Other articles dealing with real-time-data for activity measures in Chile are Chumacero and Gallego (2002), Morandé and Tejada (2008) and Pincheira and Rubio (2009).

[3] We will assume that this ignorance about the future reference year methodology also holds true for every single forecaster.

Besides this missing observations problem, we face the additional limitation of an extremely small sample. We have 19 observations of what we call *one-step-ahead* forecasts (OSA forecasts) from the CBCh and 20 observations of what we call *two-step-ahead forecasts.* (TSA forecasts). The small sample problem is even worse when we turn to private forecasters. Just to give an example, there are only eight OSA forecasts from Consensus Forecasts and ten OSA forecasts from the SPF.

Despite this small sample issue, we think that a work like ours is extremely important both from a policy and academic point of view. From a policy point of view, we need to recall that the Central Bank of Chile follows a flexible inflation targeting regime. In this particular monetary regime, inflation and output forecasts are the building blocks of monetary policy decisions. Furthermore, good forecasts not only help policy makers to make appropriate decisions, they also play a major role in the construction of a central bank key asset: credibility. This is so, because forecasts provide a solid and objective measure of the ability that a central bank may have to understand the economy. Good forecasts may help to strengthen the credibility of a central bank and therefore the efficiency of monetary policy. Bad forecasts may well work in the opposite direction.

From the academic point of view this paper is also appealing. The Central Bank of Chile works with many state of the art models to characterize the economy. A forecast evaluation of the type we make here, could be indirectly indicating the usefulness of those models either to provide good forecasts or, at least, a good understanding of the economy that enable policy makers to make well informed judgmental forecasts. Of course, in this article we are not evaluating these state of the art models directly. We are evaluating the final output of a long decision making process in which these models may play a role.

We show results covering a wide variety of issues. We compare the accuracy of the CBCh's forecasts using both first vintages and revised GDP growth data. We also analyze whether the forecasts by the CBCh are optimistic or pessimistic when compared with private analysts' forecasts. We also analyze forecast efficiency and whether forecasts have been more accurate in the recent years or in the distant past. Finally, we analyze the empirical coverage of the forecasting intervals reported by the CBCh.

We report mixed results in terms of root mean squared prediction errors. Depending on the benchmark, the sample period, the forecast horizon, and the vintage used in the analysis, forecasts from the Central Bank of Chile may outperform the benchmarks or may be outperformed by them. Despite these mixed results, differences in root mean squared prediction errors are, in general, moderate and with no statistical significance, with only one exception favoring forecasts from the Central Bank of Chile. Nevertheless, our efficiency analysis, in addition to the fact that in some periods the forecasts produced by the Central Bank of Chile have been slightly outperformed by alternative forecasts, opens the question about the room for improvement in the accuracy of the Central Bank of Chile forecasts. While the room for improvement seems to be small for point forecasts, it seems larger for interval forecasts.

The rest of the document is organized as follows: in section 2 we present a literature review. In the third section we describe the methodology we use to compute forecasts errors and to compare the CBCh forecasts errors with those from private analysts and simple time series models. In section 4 we deliver the main results of this paper, and in section 5 we provide conclusions and a brief summary of our results.

2. LITERATURE REVIEW

Most of the forecasting literature relies on statistical measures of accuracy to compare different forecasts. Actually, the most commonly used statistical measure of forecast accuracy is the mean squared prediction error (MSPE) or its squared root denoted by RMSPE.[4] Another branch of the literature focuses

[4] Although most of the literature uses error measures drawn from statistics, McCulloch and Rossi (1990), Leitch and Tanner (1991) and West et al. (1993) use economic-based measures. This is the case of evaluations where the loss functions are associated with economic criteria such as profits or measures of welfare. This kind of evaluation goes beyond the scope of this paper. In addition, McCracken and West (2002) provide an interesting dis-

on measures of forecast efficiency. This approach aims at detecting whether a particular forecast has or has not been able to properly use all the available information at the moment when forecasts were made. This idea is of old vintage and many papers have derived either tests or theoretical results on forecasts efficiency under different assumptions, see for instance Elliot and Timmermann (2008). In particular, under quadratic loss, efficient forecast errors should be unbiased and uncorrelated with variables in the information set used for the construction of forecasts. If past forecast errors belong to this information set, then optimality implies a finite autocorrelation structure for these errors.[5]

Interestingly, recent literature shows that when more general loss functions are considered, efficient forecasts errors could present bias and autocorrelation. As a matter of fact, the presence of bias and autocorrelation in forecast errors might be the result of an optimal strategy when agents face asymmetric loss functions. See, for instance, Patton and Timmermann (2007), Elliott, Komunjer and Timmermann (2008), Capistrán (2007) and Capistrán and Timmermann (2008).

On empirical grounds it is usual to find articles evaluating the accuracy and efficiency of private and public forecasters. For instance, Joutz and Stekler (2000) take GDP and inflation forecasts from the Federal Reserve of the United States of America to find that they show systematic errors and similar properties and problems than private forecasts. In particular, they show that the Federal Reserve forecasts are not statistically better than simple ARIMA forecasts or those provided by surveys.

More recently, Capistrán (2007) shows that the US Federal Reserve inflation forecasts underpredicted effective inflation in a given sample period, and that over predicted effective

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cussion about the variety of metrics available in the literature to evaluate forecasts.

[5] It is important to mention the contribution of Mincer and Zarnowitz (1969) in this regard. They propose simple methods to evaluate bias and forecast efficiency. Similarly, Granger and Ramanathan (1984) and Chong and Hendry (1986) propose encompassing test to evaluate if the information embedded in a particular series of forecasts is able to explain, at least in part, another forecasting method prediction errors.

inflation in the rest of the sample. Furthermore, he also showed evidence indicating that the US Federal Reserve forecasts may have not used all the information available in private analyst forecasts.

Groen, Kapetanios and Price (2009) is another paper evaluating the forecasting performance of a central bank. As in our paper, they compare forecasts to different GDP vintages. They show that the Bank of England's inflation forecasts outperform a variety of time series benchmarks, whereas GDP growth rates forecasts are generally less accurate than traditional univariate and multivariate benchmarks. In particular they show that the traditional random walk model generates GDP forecasts that are more accurate than those of the Bank of England when prediction is made one quarter ahead. This result is robust to the different vintages used in the evaluation.

Another article evaluating forecasts by a central bank is due to Andersson et al. (2007). In this paper, the authors evaluate the relative performance of the central bank of Sweden's inflation forecasts. In general, they find that the Swedish central bank's forecasts are more accurate than forecasts provided by the National Institute of Economic Research, but the difference is not statistically significant. Moreover, their results suggest that the Swedish central bank performs quite well compared to Consensus Forecasts.

For the Euro zone, Bowls et al. (2007) analyze private analyst forecasts. Among other things, they find that private analyst have shown a tendency to underestimate effective inflation and to overestimate GDP growth. The sample period in their evaluation goes from the first quarter in 1999 to the last quarter in 2006.

In another article, Loungani (2001) evaluates GDP growth prediction errors from Consensus Forecasts in several developed and developing countries for the period 1989-1998. She finds some evidence of inefficiency and bias. She also detects a high correlation between the forecasts of international institutions (World Bank, IMF and OECD).

More recently Loungani and Rodriguez (2008) focus on the speed of adjustment in the revisions of GDP growth private forecasts in 14 countries. They show that private analyst forecasts are smoother than optimal forecasts under quadratic loss. In other words, forecasts do not seem to incorporate news properly. They change slowly, which may be a very unpleasant feature at the brink of a recession.

Romer and Romer (2008) show, in a striking paper, that US monetary policymakers have no advantage over their staff to generate better forecasts. Furthermore, they show that policymakers are not using the available information optimally. They conclude that a simple citizen looking for good GDP and inflation forecast should disregard forecasts built by policymakers and use forecasts built by the Board of Governors staff. Ellison and Sargent (2009) explain that the striking results in Romer and Romer (2008) are consistent with US monetary policymakers being rational but caring about a worst-case scenario. In their view, US monetary policymakers may be seen as using efficiently the information provided by their staff in a context of model uncertainty, in which they have doubts about the specification and limitations of the models.

Some research in the topic of forecasts evaluation has also been carried out in Chile. Chumacero (2001), for instance, analyses private forecasters' estimates of GDP growth rates during the period 1986-1998. His results show that forecasters systematically underestimate the true growth rate of the economy.

More recently Bentancor and Pincheira (2010) shows that inflation forecasts from the SPF in Chile display a significant downward bias and excess of autocorrelation in the second half of their sample period. By correcting this autocorrelation in an out-of-sample exercise, the authors achieve significant reduction in MSPE and bias.

The paper by Albagli et al. (2003) is probably the closest to ours. These authors evaluate the Central Bank of Chile inflation and GDP growth forecasts errors. They run a horse race between the Central Bank of Chile and private analysts. They also make a comparison against foreign central banks. They conclude that in the last years of their sample period (1991-2002) there has been an improvement in the accuracy of the Central Bank of Chile forecasts. They also show that GDP forecasts errors from the Inflation Report of the Central Bank of Chile are marginally larger than those of private analysts but that the performance of inflation forecasts from the Central Bank is significantly better than that of private analysts. They also mention that the performance of the Central Bank of Chile forecasts is similar to the performance of others Central Banks in the world.

This brief and selective review of the literature shows two interesting facts: most public and private forecasts display some degree of inefficiency when traditional metrics of predictive ability are used. Secondly, simple univariate and multivariate benchmarks may be competitive and even more accurate for some variables and horizons than forecasts produced by central banks or private analysts. In the following sections we will see how our evaluation of the Central Bank of Chile GDP growth forecasts fits in with the existing literature.

3. DATA AND METHODOLOGY

We aim at evaluating annual real seasonally unadjusted GDP growth forecasts from the Central Bank of Chile over the sample period between 1991 and 2009. We obtained these forecasts from two sources: the Central Bank of Chile's Inflation Report and the Central Bank of Chile's report to the Congress. We use these two sources to get as many forecasts as we can. Reports to the Congress are available from 1991 to 2000. In these reports, released on September of each year, the Central Bank of Chile provided annual GDP growth rate forecasts for the current year as well as for the following year. Since 2001, the Central Bank of Chile publishes Inflation Reports three times a year, in January, May and September.[6] As in the previous Reports to the Congress, the September issue of the Inflation Report includes annual GDP growth rate forecasts for the current and following years. We focus on these forecasts released on September to carry out our analysis. Focusing on the September report has two advantages. First, we include a larger number of observations by considering both forecasts from the report to the Congress and the Inflation

[6] Since 2009 the usual policy of writing three Inflation Reports per year was changed to a policy of writing four Inflation Reports. These four reports are to be released in March, June, September and December.

Reports. Second, we can evaluate forecasts at two different horizons, the end of the current year, which we call one-step-ahead forecasts (OSA forecasts) and forecasts made for the end of the subsequent year, which we call two-step-ahead forecasts (TSA forecasts).

Since 2002, inflation reports contain two slightly different forecasts. These reports provide an explicit interval forecast and also an implicit point forecast that can be inferred from the domestic demand, exports and imports forecasts available in the reports. For the sake of simplicity we will consider two series of point forecasts, those corresponding to the center of the interval explicitly released, and those implicitly obtained from the domestic demand, imports and exports forecasts. These two series of forecasts are labeled Central Bank of Chile's forecasts 1 and 2 (CB1, CB2). These two series of forecasts are similar. They are displayed in table 1.

We compare the Central Bank of Chile's GDP growth forecasts with different benchmarks. We use predictions for the

TABLE 1. CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS, 1991-2010 (in percentages)

	CB1		*CB2*	
	One step ahead	*Two steps ahead*	*One step ahead*	*Two steps ahead*
1991	5.0	5.0	5.0	5.0
1992	7.5	5.0	7.5	5.0
1993	5.6	5.5	5.6	5.5
1994	4.0	4.5	4.0	4.5
1995	7.0	5.3	7.0	5.3
1996	6.8	6.0	6.8	6.0
1997	5.8	5.8	5.8	5.8
1998	5.0	6.8	5.0	6.8
1999	0.1	3.8	0.1	3.8
2000	5.6	5.0	5.6	5.0
2001	3.7	5.7	3.7	5.7
2002	2.3	5.0	2.2	5.0
2003	3.3	4.0	3.1	4.0
2004	5.3	4.5	5.1	4.5
2005	6.3	5.0	6.1	5.3
2006	5.0	5.8	4.7	5.7
2007	6.0	5.8	6.0	5.6
2008	4.8	5.5	4.7	5.5
2009	−1.8	4.0	−1.7	3.8
2010		5.0		5.0

NOTES: One-step-ahead forecasts are released on September of the current year. Two-step ahead-forecasts are released on September of the previous year.

current and subsequent year collected from the September issue of Consensus Forecasts for the 2001-2009 period. We also use the information in the Survey of Professional Forecasters (SPF) carried out periodically by the Central Bank of Chile. Forecasts for the current year can be deduced from their quarterly forecasts for the period 2000-2002. We gave the SPF a little advantage over the Central Bank, by looking at the survey carried out in October, which actually has GDP predictions for the last two quarters of the corresponding year. From 2003 until now, the survey provides explicit forecasts for the current and subsequent year.[7] Finally, we take advantage of the real time database of quarterly GDP containing, for a given date, the most recent quarterly GDP series available at that moment. We use this database to generate real time forecasts using a number of univariate time series models. In particular we consider the following models: AR(1), AR(1), a driftless rando||||m walk for the level of quarterly GDP, a driftless random walk for the quarterly growth GDP rate, one version of the *airline model* proposed by Box and Jenkins (1970), and the average of all these forecasts.[8] We also combined models for data in different frequencies. To do so, we use forecasts for the current year (one-step-ahead forecasts) from the Central Bank of Chile, from a random walk process in levels and from a version of the *airline model.* Then we plug each of these forecasts as an additional observation to the available annual series and then we estimate an ARMA(1,1) with annual GDP growth rates observations. With this last model we generate two-step-ahead forecasts which are also used as a benchmark for the two-step-ahead-forecasts released by the Central Bank of Chile.

We use the RMSPE as a measure of predictive accuracy. This measure corresponds to the squared root of the MSPE, which is defined as follows:

$$MSPE(e) = E(e^2)\ ,$$

[7] From Consensus we use the average of all the surveyed analysts. In the case of the Survey of Professional Forecasters, the Central Bank of Chile releases only the median of the predictions.

[8] We consider the following version of the *airline model:* $g_t = g_{t-1}+\varepsilon_t+\theta\varepsilon_{t-4}$, where g_t represents the accumulated GDP growth in the last four quarters and ε_t represents a white noise.

where e denotes the prediction error, defined as the actual value minus the predicted value.

While RMSPE is one of the leading metrics to evaluate predictions in the forecasting literature, some other metrics may be equally useful. In particular, we will also show results using Mean Absolute Prediction Errors (MAPE), which is defined as follows:

$$MAPE(e) = E\left|e\right| ,$$

where e denotes the prediction error defined as before.

We work with quasi-finals errors as well as with first vintage errors. Quasi-final errors are defined as the last version of actual GDP growth in a given reference year, minus the forecast. First vintage errors are defined as the first released GDP growth observation minus the forecast. As we already mentioned in the introduction, GDP growth figures undergo several rounds of revisions, so typically the first release is different from the quasi-final release, and sometimes this difference is sizeable as it is shown in table A1 in the appendix.

We also mentioned in the introduction that our analysis faces the challenge of an extremely small sample size with missing observations. The reason for the missing observations problem relies on the fact that during the sample period four different reference years for the actual calculation of real GDP were used. This is not a simple problem. GDP growth figures expressed in a given reference year are not easily translated into a different reference year. This is because different reference years may be using a different methodology to measure sectoral GDP and in general they use different weights to weight up the different sectors of the Chilean economy. Tables 2 and 3 next illustrate this problem showing the annual GDP growth rates figures in each of the four reference years covering our sample period.

There are a few years in which the Central Bank of Chile actually computed annual GDP growth rates using two different reference years. Sometimes the difference in the figures is small, but sometimes is fairly large. These differences suggest that comparing forecasts built upon information of a given reference year, with figures expressed in another reference year may be misleading, because those errors would

TABLE 2. GDP ANNUAL GROWTH RATES FOR CHILE QUASI-FINAL RELEASE, 1991-2009 (in percentages)

	Reference year			
Year	*1977*	*1986*	*1996*	*2003*
1991	6.0	8.0		
1992	10.4	12.3		
1993		7.0		
1994		5.7		
1995		10.6		
1996		7.4		
1997		7.4	6.6	
1998		3.9	3.2	
1999		–1.1	–0.8	
2000		5.4	4.5	
2001			3.4	
2002			2.2	
2003			3.9	
2004			6.2	6.0
2005			6.3	5.6
2006				4.6
2007				NYA
2008				NYA
2009				NYA

NOTES: NYA stands for *not yet available.* Quasi-final releases correspond to the last vintage for a given reference year.

TABLE 3. GDP ANNUAL GROWTH RATES FOR CHILE FIRST VINTAGE, 1991-2009 (in percentages)

	Reference Year			
Year	*1977*	*1986*	*1996*	*2003*
1991	6.0	6.1		
1992	10.4	10.3		
1993		6.0		
1994		4.2		
1995		8.5		
1996		7.2		
1997		7.1	6.6	
1998		3.4	3.2	
1999		–1.1	–1.0	
2000		5.4	4.4	
2001			2.8	
2002			2.1	
2003			3.3	
2004			6.1	6.0
2005			6.3	5.7
2006				4.0
2007				5.1
2008				3.2
2009				

TABLE 4. CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECAST ERRORS QUASI–FINAL RELEASE, 1991–2009

	CB1		*CB2*	
	One step ahead	*Two steps ahead*	*One step ahead*	*Two steps ahead*
1991	100	100	100	100
1992	290	540	290	540
1993	140	MO	140	MO
1994	170	120	170	120
1995	360	530	360	530
1996	60	140	60	140
1997	160	160	160	160
1998	–110	–290	–110	–290
1999	–120	–490	–120	–490
2000	–20	40	–20	40
2001	MO	MO	MO	MO
2002	–10	MO	0	MO
2003	60	–10	80	–10
2004	90	170	110	170
2005	0	130	20	100
2006	MO	MO	MO	MO
2007	NYA	MO	NYA	MO
2008	NYA	NYA	NYA	NYA
2009	NYA	NYA	NYA	NYA
Average full sample	84	95	89	93
Average 2001–2009	35	97	53	87

NOTES: NYA stands for *not yet available*. Quasi–final releases correspond to the last vintage for a given reference year. MO stands for *missing observations*.

correspond to the sum of the forecasts errors plus the error due to the change in reference year. Unfortunately, sometimes the Central Bank of Chile released figures expressed in only one reference year. Every time that reference year is different to the reference year on which forecasts were originally built we treat those forecasts errors as missing observations. We do this to avoid an unfair evaluation of forecast ability when errors may be affected by changes in reference years.[9]

[9] The missing observations problem arises when a change in the reference year is about to take place but the future methodology for computing GDP is not yet released. In these occasions, we assume forecasters provide GDP forecasts expressed in the *old* reference year. According to our records, the methodology associated to the 1986 reference year was released in October 1992. Therefore, we assume that by September of that year the new methodology was unknown for private and public forecasters, which implies that one and two-step-ahead forecasts were made in the 1977 reference year. Because there is no GDP growth observations for the year 1993

Tables 4 and 5 next show forecasts errors from the Central Bank of Chile. We find missing observations in the following years 1993, 2001, 2002, 2006 and 2007.

Due to the very small sample we are working with, we will use critical values from a t(n-1) distribution when showing results of the Giacomini and White (2006) test (which actually

TABLE 5. CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECAST ERRORS FIRST VINTAGE, 1991–2009

	CB1		*CB2*	
	One step ahead	*Two steps ahead*	*One step ahead*	*Two steps ahead*
1991	100	100	100	100
1992	290	540	290	540
1993	40	MO	40	MO
1994	20	–30	20	–30
1995	150	320	150	320
1996	40	120	40	120
1997	130	130	130	130
1998	–160	–340	–160	–340
1999	–120	–490	–120	–490
2000	–20	40	–20	40
2001	MO	MO	MO	MO
2002	–20	MO	–10	MO
2003	0	–70	20	–70
2004	80	160	100	160
2005	0	130	20	100
2006	MO	MO	MO	MO
2007	–90	MO	–90	MO
2008	–160	–230	–150	–230
2009	NYA	NYA	NYA	NYA
Average full sample	–8	–24	–2	–26
Average 2001–2009	–32	–3	–18	–10

NOTES: NYA stands for *not yet available.* MO stands for *missing observations.*

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expressed in the 1977 reference year, we have a missing observation for the two-step-ahead forecast error in the year 1993. Similarly, the methodology associated with the 1996 reference year was released in September 2001. We make the assumption that forecasts built in that month were based on the *old* reference year. Because there are no GDP growth observations for the years 2001 ad 2002 expressed in the 1986 reference year, we have missing observations for the two-step-ahead forecast error in the years 2001 and 2002 and for the one-step-ahead forecast error in 2001. Finally, the last reference year was released in November 2006. We assume that forecasts generated by September 2006 were made in the 1996 reference year which explains the missing observations in tables 4 and 5 corresponding to years 2006 and 2007.

coincides in this setting with the test by Diebold and Mariano; 1995 and West; 1996). Harvey, Leybourne and Newbold (1997) show via simulations that these critical values improve the size of the test in small samples.

4. MAIN RESULTS

In this section we present the main results of our analysis. First we show comparisons of RMSPE. Second, we show results concerning efficiency of the forecasts. In the third subsection we provide additional results regarding the behavior of the forecasts under consideration. Finally in the fourth section we show simple results regarding the coverage of the interval forecasts.

4. 1. Forecast accuracy

We first compare the accuracy of the Central Bank of Chile forecasts with the accuracy of forecasts from Consensus Forecasts and the Survey of Professional Forecasters. Tables 6 and 7 show our RMSPE results.

TABLE 6. ROOT MSPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS COMPARISON AGAINST CONSENSUS FORECASTS

	Quasi-final release		*First vintage*		*First vintage in restricted sample*	
	One step ahead	*Two step ahead*	*One step ahead*	*Two step ahead*	*One step ahead*	*Two step ahead*
CB1	54	124	82	158	41	126
CB2	69	114	83	153	52	116
Consensus	67	137	67	152	51	133
Sample Size	4	3	6	4	4	3

Table 6 shows RMSPE for the Central Bank and Consensus Forecasts. Figures in the first three rows are expressed in basis points. The last row shows the number of observations in the analysis. It is remarkable how low this number is, which makes us to be very cautious when analyzing our results.

We focus on two targets: quasi-final GDP growth releases and first GDP growth vintages. The first two columns in table 6

TABLE 7. ROOT MSPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS COMPARISON WITH THE SURVEY OF PROFESSIONAL FORECASTERS

	Quasi-final release		*First vintage*		*First vintage in restricted sample*	
	One step ahead	*Two step ahead*	*One step ahead*	*Two step ahead*	*One step ahead*	*Two step ahead*
CB1	49	151	76	178	38	146
CB2	62	139	77	172	48	133
SPF	67	169	70	180	56	163
Sample Size	5	2	7	3	5	2

indicate that when forecasts are compared to quasi-final GDP releases, forecasts from the Central Bank of Chile labeled as CB1 have been more accurate than Consensus' forecasts at both horizons. Central Bank of Chile's forecasts labeled as CB2 have been slightly less accurate than Consensus when predictions are made one-step-ahead. For the two-step–ahead forecasts, the Central Bank of Chile forecasts CB2 have been also more accurate than those of Consensus. The third and fourth column in table 6 show results when forecasts are compared to the first vintage of GDP growth. Now, forecasts from the Central Bank of Chile CB1 and CB2 are less accurate than those of Consensus no matter what predictive horizon we consider.

The last two columns in table 6 shows results for first vintages when the sample is restricted to the same years included in columns one and two. The reason why we include these columns will be clearer in the following paragraph. In these two columns forecasts from the CBCh are more accurate than Consensus' forecasts. In summary, in this horse race between the CBCh and Consensus forecasts there is no a clear winner. Depending on the vintage under consideration, the forecast horizon and the sample period, we can either have the CBCh or Consensus Forecasts as a winner. Furthermore, maybe the most interesting result is that differences in RMSPE between the CBCh's forecasts and Consensus' forecasts are rather small.

We are also interested in determining whether forecasts are more accurate when compared with quasi-final releases or first vintages. With this in mind we could proceed by comparing the results in the first two columns with those in the third and fourth columns in table 6. Nevertheless, results from these columns are not directly comparable. The reason for this is that there are more vintages than quasi-final releases, so more observations are included in the computation of the first vintage RMSPE. To overcome this problem, table 6 includes two additional columns (five and six) presenting RMSPE using first vintages but restricting the sample to the same years included in the results displayed in columns one and two. This enables us to make a fair comparison between quasi-final and first vintage RMSPE using exactly the same sample period. This is important, because in small samples, the addition of one extra observation in only one of the two statistics we are computing may introduce an unpleasant noise.

When comparing results in column one with those in column five and those in column two with the results in column six, we find a clear pattern for one-step-ahead forecasts: predictions seem more accurate when compared to first vintages. For two-step-ahead forecasts there is no clear pattern. Besides, differences in RMSPE at this forecasting horizon are very small.

Table 7 has the same structure showed in table 6 but now RMSPE are reported for the Central Bank of Chile's forecasts and for the SPF's forecasts. Notice that RMSPE shown for the Central Bank of Chile's forecasts need not to be the same to those in table 6. This is because the sample period is slightly different in both tables. Differing from the previous analysis, now we see that the two-step-ahead forecasts from the CBCh are more accurate than two-step-ahead forecasts from the SPF. This result is robust to the sample period and the vintage under consideration. For one-step-ahead forecasts we have mixed results: the first column indicates that one-step-ahead forecasts from the Central Bank of Chile have been more accurate than those of the SPF when forecasts are compared with quasi-final releases. When predictions are compared to first vintages, column three shows that one-step-ahead forecasts from the Central Bank of Chile have been outperformed by those of the SPF. Finally, column five shows that comparing forecasts with first vintages during the same years used in column one produces the same output as in table 6: The Central Bank of Chile does a better job than the SPF. Therefore, in this horse race between the CBCh and SPF's forecasts, the CBCh is a clear winner for predictions two-step-ahead. There is no a clear winner, however, when considering one-step-ahead forecasts. Again we see that differences in RMSPE between forecasts from the CBCh and the SPF are rather small.

Table 7 also displays a clear pattern regarding the accuracy of forecasts when compared to first and quasi-final vintages: predictions are more accurate when compared to first vintages. (see table 7 columns 1,5 and 2,6).

To complement our analysis, we also compare the Central Bank of Chile's forecasts with forecasts from simple time-series models. We use several specifications of ARMA(p,q) models estimated with recursive windows over a real time sample at quarterly frequency. Tables 8 and 9 show RMSPE results using the same structure previously shown in table 6.

TABLE 8. ROOT MSPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS COMPARISON WITH TIME SERIES MODELS, FULL SAMPLE PERIOD

	Quasi-final release		*First vintage*		*First vintage in restricted sample*	
	One step ahead	*Two step ahead*	*One step ahead*	*Two step ahead*	*One step ahead*	*Two step ahead*
CB1	156	291	117	262	115	265
CB2	157	290	117	261	116	264
R. Walk Level	130	674	121	607	91	628
R. Walk Growth Rate	165	461	134	436	140	446
Airline Model*	154	390	119	366	117	375
AR(1)	156	378	135	364	139	372
AR(2)	208	337	134	347	137	355
Average	163	448	119	354	125	435
Sample Size	14	12	16	13	14	12

The most remarkable result in tables 8 and 9 is the overwhelming good performance of the Central Bank of Chile two-step-ahead forecasts compared to forecasts from time-series models. At times ARMA forecasts display RMSPE that are about twice as big as those from the Central Bank of Chile. We will go back to this point later.

One-step-ahead forecasts from time series models are more competitive than their two-stepahead counterparts. The first two columns in table 8 indicate that when forecasts are compared to quasi-final GDP releases, one-step-ahead forecasts from the Central Bank of Chile are outperformed by a random walk in levels and by the variation of the *airline model* we are working with. The third and fourth columns in table 8 show results when forecasts are compared to GDP growth first vintages. Now, one-step-ahead forecasts from the Central Bank of Chile are slightly more accurate than the best one-step-ahead forecasts of the time series models. Column five indicates that the CBCh is only outperformed by the best time-series strategy when the sample is restricted to the same years used in column one and predictions are compared to first vintages.

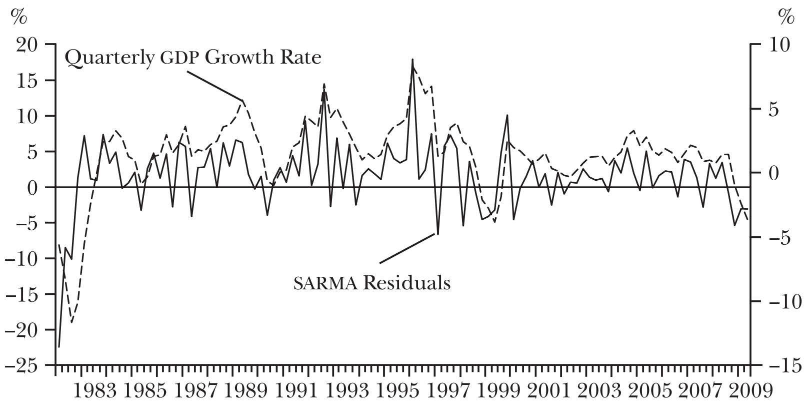
TABLE 9. ROOT MSPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS COMPARISON WITH TIME SERIES MODELS, PERIOD 2001-2008

	Quasi-final release		*First vintage*		*First vintage in restricted sample*	
	One step ahead	*Two step ahead*	*One step ahead*	*Two step ahead*	*One step ahead*	*Two step ahead*
CB1	54	124	82	158	41	126
CB2	69	114	83	153	52	116
R. Walk Level	38	524	147	458	55	509
R. Walk Growth Rate	81	249	71	245	61	228
Airline Model*	90	271	98	251	70	257
AR(1)	93	262	84	247	77	248
AR(2)	108	285	97	263	91	274
Average	90	315	90	272	72	301
Sample Size	4	3	6	4	4	3

We also notice that in all cases but one, forecasts are more accurate when compared to first vintages than when compared to quasi-final releases (see columns 1,5 and 2,6 in table 8).

Table 9 is similar to table 8. The only difference relies in the sample period. Table 9 shows results when the sample is restricted to the period 2001-2008. We do this because of the structural change in the volatility of GDP growth already reported in the literature (see Calani, Fuentes and García, 2009; and Betancour, De Gregorio and Medina, 2006). Figure 1 shows quarterly GDP growth rates for the Chilean Economy, as well of the residuals of a SARMA(1,0,1)x(0,0,1) process for the same variable. This figure shows clearly that from some point near to 2001, the Chilean economy experienced a reduction in GDP growth volatility. To give some numbers, the standard deviation of GDP growth rates fell from 6.3% in the period 1982Q1-1999Q4 to 2.4% in the period 2001Q1-2009Q2. This reduction holds true even if we do not consider the first observations which could be considered as outliers given the magnitude of the 1982 crisis. When we discard observations corresponding to years 1982-1984, the reduction in volatility is still significant, falling from 4.2% to 2.4% in the last period.

FIGURE 1. QUARTERLY GDP GROWTH RATES AND RESIDUALS OF SARMA SPECIFICATION, 1982-2009



Results in table 9 confirm the excellent relative performance of the Central Bank of Chile GDP growth two-step-ahead forecasts. On the other hand, we get mixed results for the Central Bank of Chile one-step-ahead forecasts: sometimes they are the best but sometimes they are outperformed by the best time-series forecasts. Finally, in all cases but one RMSPE are much lower in table 9 than in table 8, indicating a strong reduction in the size of forecasts errors. It is important to remark that this reduction holds true for all forecasts: CB1, CB2 and those from time series models. It is clearly very difficult to correctly identify the sources behind this increment in forecast accuracy, so we leave this problem for future research.[10]

We also carry out an additional exercise aimed at producing better two-step-ahead forecasts than those from ARMA(p,q) models. We take one-step-ahead forecasts from two time series

[10] See Betancour, De Gregorio and Medina (2006) for possible explanations of the Chilean moderation.

models and from the Central Bank of Chile (CB1), and we consider them as an additional true observation of annual GDP growth. Then we estimate an ARMA(1,1) model at annual frequencies to generate a two-step ahead forecast according to our terminology. RMSPE of this exercise are shown in tables 10 and 11.

TABLE 10. ROOT MSPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS. COMPARISON WITH CONCATENATION METHODS, FULL SAMPLE PERIOD

	Quasi-final release	*First vintage*	*First vintage in restricted sample*
	Two steps ahead	*Two steps ahead*	*Two steps ahead*
CB1	302	271	275
CB2	301	270	274
CB1-ARMA(1,1)	315	299	299
R. Walk Level-ARMA(1,1)	272	294	292
Airline Model*-ARMA(1,1)	341	315	316
Sample Size	11	12	11

TABLE 11. ROOT MSPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS. COMPARISON WITH CONCATENATION METHODS, PERIOD 2001-2008

	Final release	*First vintage*	*First vintage in restricted sample*
	Two steps ahead	*Two steps ahead*	*Two steps ahead*
CB1	124	158	126
CB2	114	153	116
CB1-ARMA(1,1)	70	165	91
R. Walk Level-ARMA(1,1)	119	208	151
Airline Model*-ARMA(1,1)	75	170	89
Sample Size	3	4	3

The good news arising from tables 10 and 11 is that these concatenating strategies generate relatively accurate forecasts that are competitive with those of the Central Bank of Chile and private analysts. Anyway, Central Bank of Chile forecasts outperform these concatenating strategies when forecasts are compared with first vintages in the longest available sample. This result, however, is overturned in different sub samples when compared with either first vintages or quasi-final releases, so its robustness is still questionable.

Beyond these mixed results, differences in root mean squared prediction errors are in general either small or moderate and with no statistical significance.[11] Table A2 in the appendix show the magnitude in the difference of RMSPE for selected forecasting methods. Just in one occasion there is a statistically significant difference and it favors forecasts produced by the Central Bank of Chile. It is worth mentioning that in another occasion the difference is almost significant at the 10% significance level favoring the CBCh as well.

Tables A3 to A8 in the appendix are the analogs of tables 6 to 11 but now constructed using MAPE. These tables show in general similar results to those obtained from tables 6 to 11, but at least one interesting fact is worth of mention: MAPE are lower than RMSPE. This is because a quadratic form imposes a higher penalty to large errors. For instance, in terms of an absolute loss function, two errors of fifty basis points are the same as two errors of five and 95 basis points (MAPE of fifty basis points). In the case of a quadratic loss function these two sets of errors yield different outcomes (50 and 67.3 respectively).

4. 2. Efficiency

The last two rows in tables 4 and 5 show average Central Bank of Chile forecast errors. Given the fact that all the averages in table 4 are positive, and that all the averages in table 5 are negative, it is tempting to conclude that on average the Central Bank of Chile has under predicted GDP growth when forecasts are compared with quasi-final releases and has over predicted GDP growth when compared with first vintages. Nevertheless, neither of these averages is statistically different from zero nor stable along time. Furthermore, we think it is more relevant to emphasize the autocorrelation of one-step-ahead forecasts errors. Tables 4 and 5 shows persistence in the sign of forecasts errors which means that they look like a sequence of nonnegative errors followed by another sequence of positive errors. In other words, once the Bank under predicts GDP growth it is likely to repeat that under prediction in the following year. In fact, the estimated probability

[11] 54 basis points is the biggest difference in RMSPE. The second biggest difference is 31 basis points.

of making a mistake in the same direction next year is 82% when considering a comparison with first vintages. Two-step-ahead forecast errors show much lower autocorrelation. In fact the probability of making an error next year in the same direction is 52%, much similar to a fair coin toss, although the number of observations is really low to make a reliable case.

Table 12 below complements this analysis showing the first order autocorrelation coefficient for one-step-ahead errors, and the second order autocorrelation coefficient in the case of two step-ahead forecast errors. Our results confirm the presence of autocorrelation in one-stepahead forecasts errors in the sense that in three out of four evaluations, the autocorrelation coefficient is statistically significant at the 90% confidence level. This is traditionally considered an indication of inefficiency. This is found despite the fact that we are working with an extremely small sample including missing observations. We recall that missing observations is a serious problem that may generate a bias towards not detecting existing autocorrelation, so we think this result is important. Interestingly, at longer horizons no evidence of autocorrelation is found.

TABLE 12. CENTRAL BANK OF CHILE GROWTH RATE FORECASTS ERRORS AUTOCORRELATION ANALYSIS, FULL SAMPLE

	Coefficient	*Std. Error*	*t-Statistic*	*P-Value.*
CB1 Quasi-final release OSA	0.37	0.22	1.82	0.10
CB1 First vintage OSA	0.34	0.17	1.97	0.07
CB2 Quasi-final release OSA	0.36	0.23	1.64	0.13
CB2 First vintage OSA	0.34	0.18	1.93	0.08
CB1 Quasi-final release TSA	0.13	0.20	0.64	0.55
CB1 First vintage TSA	–0.13	0.15	–0.86	0.43
CB2 Quasi-final release TSA	0.14	0.21	0.68	0.52
CB2 First vintage TSA	–0.12	0.15	–0.79	0.46

Note: P-Value computed according to a t(n-2) distribution.

4. 3. Are the forecasts really different?

As we mentioned in earlier sections, working with a small sample with missing observations is a serious problem when applying traditional inference methods. This problem may be by- passed, at least partially, if we focus on analyzing the forecasts rather than forecasts errors. Forecasts do not suffer from missing observations and also we can count with forecasts made for years 2009 and 2010, increasing the number of observations we can work with.

We explore the relationship between forecasts in two dimensions. First, we look for optimistic and pessimistic agents. Second, we take a look at the correlation between these forecasts.

TABLE 13. CENTRAL BANK OF CHILE GROWTH RATE FORECASTS MINUS BENCHMARK FORECASTS, 2001-2010

	CB1 OSA	*CB2 OSA*	*CB1 TSA*	*CB2 TSA*
Consensus OSA	17[a]	6		
SPF OSA	13[b]	3		
TS OSA	39[a]	11		
Consensus TSA			34[a]	32[a]
SPF TSA			37[b]	34[b]
TS TSA			(65[a])	(67[a])

NOTES: Information from the Survey of Professional Forecasters corresponds to the period 2000-2010. [a] Significance at 1%. [b] Significance at 5%.

Table 13 shows the difference between forecasts from the Central Bank of Chile and three different benchmarks: those from Consensus, from the SPF and the average of a number of time series models. We consider the period 2001-2010. Interestingly, all figures comparing forecasts from the Central Bank of Chile and private analysts are positive, indicating that the Central Bank of Chile has a relatively optimistic view regarding the Chilean growth process. This is true, irrespective of the horizon and forecast of the Central Bank we consider. Besides, this difference is statistically significant in six out of the eight relevant comparisons. From the economic point of view, some of the figures in table 13 are negligible, but some others might be relevant in terms of monetary policy.

Interest rate setting implications of this difference are interesting as well. Let us recall that a traditional equation characterizing the decisions of a Central Bank is called *Taylor rule*. A traditional version of this rule usually incorporates a contemporary output-gap term. This term suggests a raise in interest rates whenever current output is higher than potential output and a decrease in interest rates whenever current output is below potential output. Nevertheless, current output gap is never observed. This happens because potential output is by definition an unobservable variable, and because output observations are released with some lag. Under these conditions, central banks need to build nowcasts of current output gap. If there is no difference in the views of the Central Bank of Chile and private analysts regarding potential output, and short term GDP forecasts are proper proxies of GDP nowcasts, then it is reasonable to expect, based on a Taylor rule type of equation, that private analysts would have rather set lower monetary policy rates than those that have actually been set by the Central Bank of Chile.

It is important to remark that this optimistic relative behavior of the Central Bank of Chile needs not to be a problem in terms of forecasts accuracy. This is because we are not talking about bias in the forecasts. In principle there is no way we can label this optimistic behavior as good or bad news in terms of accuracy.

Finally let us take a look at the correlation of forecasts. Table 14 shows the correlation between one-step-ahead GDP growth forecasts from the Central Bank, a random walk model in levels, in growth and for an average of a number of time series predictions. Table 15 includes also correlations between all these forecasts and those from Consensus and the SPF for the period 2001-2009.

TABLE 14. CORRELATION OF ONE-STEP-AHEAD FORECASTS, 1991-2009

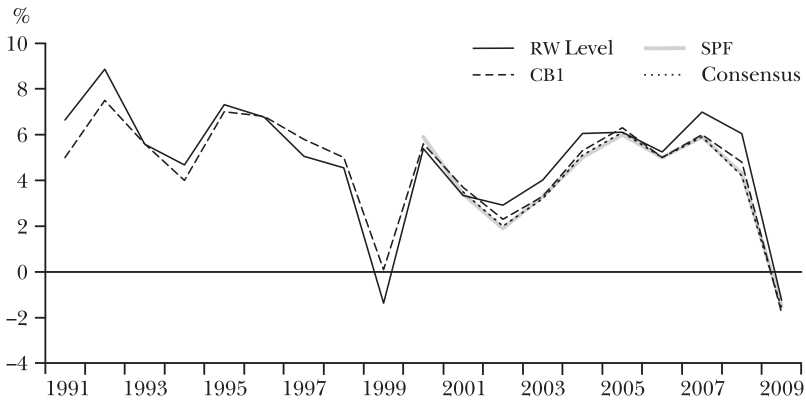
	CB1	*CB2*	*TS Average*	*RW Level*	*RW Growth*
CB1	1.000	0.999	0.997	0.978	0.997
CB2	0.999	1.000	0.997	0.978	0.995
TS Average	0.997	0.997	1.000	0.976	0.997
RW Level	0.978	0.978	0.976	1.000	0.966
RW Growth	0.997	0.995	0.997	0.966	1.000

TABLE 15. CORRELATION OF ONE-STEP-AHEAD FORECASTS, 2001-2009

	CB1	*CB2*	*TS Average*	*RW Level*	*RW Growth*	*Consensus*	*SPF*
CB1	1.000	0.999	0.997	0.978	0.997	0.997	0.996
CB2	0.999	1.000	0.997	0.978	0.995	0.995	0.995
TS Average	0.997	0.997	1.000	0.976	0.997	0.992	0.992
RW Level	0.978	0.978	0.976	1.000	0.966	0.969	0.973
RW Growth	0.997	0.995	0.997	0.966	1.000	0.995	0.993
Consensus	0.997	0.995	0.992	0.969	0.995	1.000	0.999
SPF	0.996	0.995	0.992	0.973	0.993	0.999	1.000

Results are striking. The lowest correlation is 0.966 and most of them are around 0.99. These results suggest that one-step-ahead forecasts are very similar. Figure 2 confirms this conclusion.

FIGURE 2. SEVERAL ANNUAL GDP GROWTH ONE-STEP-AHEAD FORECASTS, 1991-2009



Actually, forecasts from Consensus, the SPF and the Central Bank of Chile look almost the same. It is also intriguing that all these forecasts follow quite close the behavior of a random walk in levels. This is especially evident when considering the forecasts first difference. Just to emphasize this, let us mention that the direction in which forecasts move it is exactly the same for all forecasts. This means that if one particular agent has a forecast that is higher that the forecast he or she had in the previous year, then, most likely, the rest of the forecasts will similarly move up compared with the previous year forecast.

Two-step-ahead forecast show a different picture. Correlations are lower, and sometimes much lower, indicating that two-step-ahead forecasts seem to be significantly different. In particular Central Bank of Chile forecasts are only mildly correlated with time-series forecasts. On the opposite side of the coin, we still see an important correlation of the CBCh's forecasts with those of private analyst. This is confirmed in figure 3 which shows the evolution of a number of two-step-ahead forecasts including those of the Central Bank of Chile, Consensus and the SPF. Regarding the direction in which forecasts move, again it is exactly the same between consensus and CB2, and similar but not equal to that predicted by the SPF. Time series forecasts are quite similar in their direction to those of the Central Bank of Chile during the first half of the sample, but in the second half this link is a little weaker.

FIGURE 3. SEVERAL ANNUAL GDP GROWTH TWO-STEP-AHEAD FORECASTS, 1991-2010

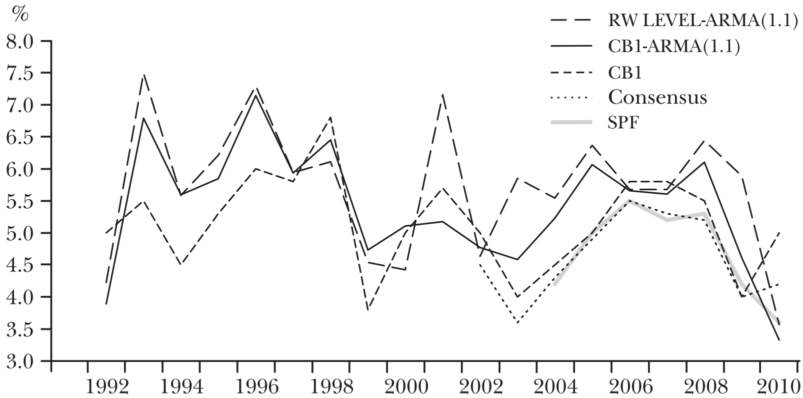


TABLE 16. CORRELATION OF TWO-STEP-AHEAD FORECASTS FULL SAMPLE

	(1)	*(2)*	*(3)*	*(4)*	*(5)*
(1) CB1	1.000	0.992	0.591	0.412	0.555
(2) CB2	0.992	1.000	0.608	0.420	0.569
(3) CB1-ARMA(1,1)	0.591	0.608	1.000	0.813	0.824
(4) R. Walk Level-ARMA(1,1)	0.412	0.420	0.813	1.000	0.850
(5) Airline Model*-ARMA(1,1)	0.555	0.569	0.824	0.850	1.000

TABLE 17. CORRELATION OF TWO-STEP-AHEAD FORECASTS, 2001-2009

	(1)	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*	*(7)*
(1) CB1	1.000	0.980	0.454	0.180	0.555	0.937	0.722
(2) CB2	0.980	1.000	0.493	0.204	0.585	0.922	0.710
(3) CB1-ARMA(1,1)	0.454	0.493	1.000	0.725	0.817	0.710	0.883
(4) R. Walk Level-ARMA(1,1)	0.180	0.204	0.725	1.000	0.892	0.294	0.721
(5) Airline Model*-ARMA(1,1)	0.555	0.585	0.817	0.892	1.000	0.708	0.884
(6) Consensus	0.937	0.922	0.710	0.294	0.708	1.000	0.934
(7) SPF	0.722	0.710	0.883	0.721	0.884	0.934	1.000

4. 4. Coverage of interval forecasts

As we already mentioned, since 2002, inflation reports contain two slightly different forecasts. These reports provide an explicit interval forecast and also an implicit point forecast that can be inferred from the domestic demand, exports and imports forecasts available in the reports. In this subsection we show the intervals displayed in the inflation reports and also their empirical coverage, which is nothing but the percentage of times that actual releases are contained within these intervals. Tables 18 and 19 show coverage results when forecasts are compared with quasi-final and first GDP growth releases. Successful forecasts have been remarked in shaded cells. In the last row of each table we show coverage results. Coverage is defined as the ratio between the number of successful interval forecasts and the total number of forecasts. An interval forecast is successful when the actual observation belongs to the respective interval. For coverage calculation we rule out years in which the CBCh made predictions based upon a given reference year and actual GDP growth observations were released in a different reference year, for the same reasons explained in previous sections. Results in the tables show that coverage is 0.5 for one-step-ahead forecasts and either

TABLE 18. COVERAGE OF INTERVAL FORECASTS QUASI-FINAL RELEASES, 2002-2006

	Interval forecasts		*Quasi-final GDP growth vintage Reference year*	
	One step ahead	*Two steps ahead*	*1996 (%)*	*2003 (%)*
2002	[2.0, 2.5]		2.2	
2003	[3.0, 3.5]	[3.5, 4.5]	3.9	
2004	[5.0, 5.5]	[4.0, 5.0]	6.2	
2005	[6.0, 6.5]	[4.5, 5.5]	6.3	
2006	[4.75, 5.25]	[5.25, 6.25]		4.6
Rate of success	0.5	0.33		

TABLE 19. COVERAGE OF INTERVAL FORECASTS FIRST VINTAGE, 2002-2008

	Interval forecasts		*Quasi-final GDP growth vintage Reference year*	
	One step ahead	*Two steps ahead*	*1996 (%)*	*2003 (%)*
2002	[2.0, 2.5]		2.1	
2003	[3.0, 3.5]	[3.5, 4.5]	3.3	
2004	[5.0, 5.5]	[4.0, 5.0]	6.1	
2005	[6.0, 6.5]	[4.5, 5.5]	6.3	
2006	[4.75, 5.25]	[5.25, 6.25]		4.0
2007	[5.75, 6.25]	[5.25, 6.25]		5.1
2008	[4.5, 5.0]	[5.0, 6.0]		3.2
Rate of success	0.5	0		

0.33 or 0 for two-steps-ahead forecasts, depending on the actual vintage we are comparing with. Even though we are not using other models to derive a sort of *benchmark coverage,* our coverage results seem rather low.

5. SUMMARY AND CONCLUSIONS

In this article we evaluate the Central Bank of Chile's annual GDP growth forecasts during the period 1991-2009. We compare the Central Bank of Chile' forecasts with those from the Survey of Professional Forecasters, Consensus Forecasts, and also with those obtained using simple time-series models. We evaluate a number of different forecast properties, including forecast accuracy and efficiency. In particular we place our attention on root mean squared prediction errors and autocorrelation of forecast errors. We compare the accuracy of the CBCh's forecasts using both first vintages and revised GDP growth data. We also analyze whether the forecasts by the CBCh are optimistic or pessimistic when compared with private analysts' forecasts. Furthermore, we analyze if forecasts have been more accurate in the recent years or in the distant past. Finally we analyze the empirical coverage of the forecasting intervals reported by the CBCh. Our main results follow next.

- First, our comparison of the CBCh's forecasts with those of private analysts indicates that in terms of forecast accuracy they are similar. In fact table A2 in the appendix shows that there is no statistical significance in the difference of RMSPE between one-step-ahead forecasts. The same table shows one statistically significant result favoring the Central Bank of Chile twostep-ahead forecasts, but this result is not robust to the vintage we use to compute forecast errors. Despite these findings, probably the most important conclusion is that differences in accuracy are rather small or moderate.

- Second, our analysis indicates that the CBCh's forecasts are comparable to those coming from the best time-series strategies we used. It is intriguing, however, that in some of our comparisons, simple models as the random walk in levels tends to outperform the CBCh's onestep-ahead forecasts, in a very similar result to that shown by Groen et al. (2009). Even when used in a concatenation strategy, some simple time-series models are able to outperform the CBCh's two-step-ahead forecasts. Nevertheless, we cannot identify one single superior time-series model consistently outperforming the CBCh in our different comparisons. In other words, the best time-series forecasts usually come from different models in the different exercises we carried out. For this reason it is difficult to claim superiority of time-series forecasts over forecasts from the CBCh.

- Third, we also see that the CBCh, Consensus and the SPF's one-step-ahead forecasts are more accurate when compared to first vintages than to quasi-final releases. Since 2001, the difference between the comparison with first vintages and quasi-final releases ranges between 11 and 17 basis points for one-step-ahead forecasts (favoring first vintages). In the case of two-step-ahead forecasts this clear pattern also holds true for Consensus and the SPF, but not for the CBCh.

- Fourth, despite the fact that forecasts from the CBCh are competitive when compared with private analysts and time series models, they display inefficiency in the form of excess of autocorrelation. This happens mainly in one-step-ahead-forecasts. At longer horizons forecasts seems efficient from this point of view. It is worth mentioning that this finding is consistent with a bulk of literature reporting different sources of inefficiency in private as well as public forecasts.

- Fifth, since 2001, forecasts from the CBCh and also from the majority of time series models have been on average, more accurate than in the first section of the sample (1991-2000). This is coincident with the reduction in GDP growth volatility reported in previous articles. We tend to think that this reduction of volatility is one of the basic pillars associated with the increase in forecast accuracy. The reasons explaining the growth volatility reduction, however, are beyond the scope of this article.

- Sixth, CBCh's forecasts have been, on average, more optimistic than those of Consensus and the SPF. In particular, since 2001, one of the CBCh series of forecasts (CB1) has been 17 basis points higher than that of Consensus, for one-step-ahead forecasts, and 34 basis points higher in twostep-head forecasts. Whereas this optimism has moderate size, it is systematic and statistically significant.

- Seventh, when analyzing how different several one-step-ahead forecasts are, we realize that most of them are alike. As a matter of fact, correlations are always above 0.96 and the information they contain regarding the direction of change in GDP growth is basically the same. Regarding two-step-ahead forecasts, we detect important differences between private analysts and time series forecasts. Nevertheless, private forecasts are still highly correlated to those of the Central Bank of Chile.

- Finally, we also report coverage results for the CBCh interval forecasts. Despite the small number of observations, some results are striking. For instance, irrespective of the vintage against which we compare the CBCh forecasts, only half of the times quasi-final GDP growth has fallen within the forecasting interval. For two-step-ahead forecasts the coverage is lower. It is a third when compared to quasi-final releases and zero when compared to first vintages.

In summary, and with the big caveat of having a really low number of observations, our results suggest that the CBCh's forecasts are similar to those of Consensus and the SPF. Despite these findings, our efficiency analysis, in addition to the fact that in some periods the forecasts produced by the Central Bank of Chile have been outperformed by alternative forecasts, opens the question about the room for improvement in the accuracy of the Central Bank of Chile forecasts. While the room for improvement may actually exist, according to the different benchmarks we consider in this article, this room seems to be small for point forecasts but larger for interval forecasts.

Finally, let us conclude mentioning that the tendency of greater accuracy of the CBCh's onestep-ahead forecasts when GDP is measured with first vintages poses the question about the final target of the CBCh's forecast. Should the Bank target first vintages, final revisions or both? From the point of view of building credibility, the target should be closer to first vintages that are the first numbers released to the public. On the other hand, if we think that quasi-final revisions are a better estimate of the effective GDP growth of the economy, then the target should be quasi-final revisions because they represent a better appraisal of the *true* state of the economy. From this point of view, a subject for future research should be the construction of more accurate forecasts, especially for quasi-final releases, or the construction of a unique series of GDP growth forecasts displaying robust accuracy when forecasts are compared to first and quasi-final releases.

Appendix

TABLE A1. GDP ANNUAL GROWTH RATES AND REVISIONS (in percentages)

Year	*GDP growth first vintage*	*GDP growth quasi-final release*	*Revision*
1991	6.1	8.0	1.9
1992	10.3	12.3	2.0
1993	6.0	7.0	1.0
1994	4.2	5.7	1.5
1995	8.5	10.6	2.1
1996	7.2	7.4	0.2
1997	7.1	7.4	0.3
1998	3.4	3.9	0.5
1999	−1.1	−1.1	0.0
2000	5.4	5.4	0.0
2001	2.8	3.4	0.6
2002	2.1	2.2	0.1
2003	3.3	3.9	0.6
2004	6.1	6.2	0.1
2005	6.3	6.3	0.0
2006	4.0	4.6	0.6
2007	5.1		
2008	3.2		
Average	5.0	5.8	0.7
Correlation	0.98		

NOTES: Different tones of gray represent different reference years. The darkest represents figures expressed in the 2003 reference year. The lightest represents figures expressed in the 1986 reference year. The middle zone shows figures expressed in the 1996 reference year. Quasi-final releases correspond to the last vintage for a given reference year.

TABLE A2. INFERENCE ON PREDICTIVE ABILITY, LONGEST AVAILABLE SAMPLE

	Quasi-final release		*First vintage*	
	One step ahead	*Two steps ahead*	*One step ahead*	*Two steps ahead*
CB1-Consensus	–13	–13[a]	15	6
CB2-Consensus	2	–23	16	1
CB1-EEE	–18	–18	6	–2
CB2-EEE	–5	–30	7	–8
CB1-Random Walk	26	30	–4	–23
CB2-Random Walk	27	29	–4	–24

NOTES: [a] Represents statistical significance of the Diebold-Mariano-West test, at the 10% significance level. A negative figure favors forecasts produced by the Central Bank of Chile. We carry out inference comparing MSPEs. Nevertheless, to make the interpretation easier, we show in this table the difference in RMSPEs. For the two-step-ahead comparisons against the random walk, we used the concatenating strategy using the random walk and an ARMA(1,1) model.

TABLE A3. MAPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS. COMPARISON AGAINST CONSENSUS FORECASTS

	Quasi-final release		*First vintage*		*First vintage in restricted sample*	
	One step ahead	*Two steps ahead*	*One step ahead*	*Two steps ahead*	*One step ahead*	*Two steps ahead*
CB1	40	103	58	148	25	120
CB2	53	93	65	140	38	110
Consensus	55	120	53	138	35	117
Sample Size	4	3	6	4	4	3

TABLE A4. MAPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS. COMPARISON WITH THE SURVEY OF PROFESSIONAL FORECASTERS

	Quasi-final release		*First vintage*		*First vintage in restricted sample*	
	One step ahead	*Two steps ahead*	*One step ahead*	*Two steps ahead*	*One step ahead*	*Two steps ahead*
CB1	36	150	53	173	24	145
CB2	46	135	59	163	34	130
SPF	58	165	57	177	42	160
Sample Size	5	2	7	3	5	2

TABLE A5. MAPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS. COMPARISON WITH TIME SERIES MODELS, FULL SAMPLE PERIOD

	Quasi-final release		*First vintage*		*First vintage in restricted sample*	
	One step ahead	*Two steps ahead*	*One step ahead*	*Two steps ahead*	*One step ahead*	*Two steps ahead*
CB1	121	227	89	208	84	206
CB2	124	224	91	205	87	203
R. Walk Level	93	619	92	550	71	575
R. Walk Growth Rate	140	377	112	364	115	370
Airline Model*	123	300	96	276	90	280
AR(1)	129	314	111	299	113	303
AR(2)	161	274	107	272	108	276
Average	138	316	103	352	106	307
Sample Size	14	12	16	13	14	12

TABLE A6. MAPE OH THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS. COMPARISON WITH TIME SERIES MODELS, PERIOD, 2001-2008

	Quasi-final release		*First vintage*		*First vintage in restricted sample*	
	One step ahead	*Two steps ahead*	*One step ahead*	*Two steps ahead*	*One step ahead*	*Two steps ahead*
CB1	40	103	58	148	25	120
CB2	53	93	65	140	38	110
R. Walk Level	29	510	109	428	44	487
R. Walk Growth Rate	70	230	63	227	50	207
Airline Model*	84	261	88	237	64	238
AR(1)	80	252	71	232	60	228
AR(2)	93	264	82	236	73	241
Average	71	303	83	272	58	280
Sample Size	4	3	6	4	4	3

TABLE A7. MAPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS. COMPARISON WITH CONCATENATION METHODS, FULL SAMPLE PERIOD

	Quasi-final release	*First vintage*	*First vintage in restricted sample*
	Two steps ahead	*Two steps ahead*	*Two steps ahead*
CB1	238	217	215
CB2	235	214	213
CB1-ARMA(1,1)	215	219	212
R. Walk Level-ARMA(1,1)	216	224	215
Airline Model*-ARMA(1,1)	234	241	235
Sample Size	11	12	11

TABLE A8. MAPE OF THE CENTRAL BANK OF CHILE GDP ANNUAL GROWTH RATE FORECASTS. COMPARISON WITH CONCATENATION METHODS, PERIOD 2001-2008

	Quasi-final release	*First vintage*	*First vintage in restricted sample*
	Two steps ahead	*Two steps ahead*	*Two steps ahead*
CB1	103	148	120
CB2	93	140	110
CB1-ARMA(1,1)	63	132	80
R. Walk Level-ARMA(1,1)	89	160	106
Airline Model*-ARMA(1,1)	66	138	82
Sample Size	3	4	3

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A larger slice of a growing pie: the role of emerging Asia in forecasting commodity prices

1. INTRODUCTION

Commodity prices are a key determinant of Canada's terms of trade, inflation and the exchange rate. As such, the monitoring and forecasting of commodity prices are a critical component of the monetary policy process. Since 2005, commodity prices have been subject to considerable volatility: the increase, decline, and subsequent rebound in prices have been dramatic. While the volatility may be linked to fluctuations in the business cycle, the underlying increase in commodity prices seems to be driven by the increase in demand from emerging markets. In particular China and other non-

Paper prepared by O. Gervais, I. Kolet and R. Lalonde, International Economic Analysis Department, Bank of Canada. Thies paper was presented at the XIV Central bank Researchers Network of the Americas Meeting, in Salvador, Bahia, November 11-13, 2010. The authors would like to thank Ron Alguist, Sharon Kozicki, Robert Lafrance, Robert Lavigne, Eric Santor, and other colleagues in the International Economic Analysis Department for their helpful comments and suggestions. Please do not cite without author's permission. Correspondence author: ⟨ogervais@bankofcanada.ca⟩.

OECD countries have played a role in driving commodity demand, and prices, higher in recent years.[1] Based on simulations using the Bank of Canada's large DSGE model (BoC-GEM), most of the increase in commodity prices observed over the 2000-2007 period can be explained by strong output growth in emerging Asia combined with an increase of commodity intensity in this region (Elekdag, Lalonde, Laxton, Muir and Pesenti, 2008). More recently the financial crisis and great recession have played an important role in the behavior of the commodity prices.

Naturally, forecasting commodity prices presents a significant empirical challenge. Previous Bank of Canada commodity price projection models explained non-energy commodity price movements primarily using a measure of the world output gap, and as such, ignored the impact of strong productivity growth in Asia or even in the United States.[2] These models were also unable to answer questions related to the increase in commodity intensity and the effects of the industrialisation process among important emerging market economies.

This paper describes a new forecasting model for both energy and non-energy commodity prices that for the first time incorporates the role of the increasing importance of emerging Asian commodity demand and the rise in commodity intensity in these regions. We find an empirical link between commodity prices and both the share of China in world GDP (a proxy for commodity intensity), and the world GDP growth rate (a proxy for global demand). This model also has the benefit of linking the projection of commodity prices to the other large policy analysis models used at the Bank of Canada, and as such is able to produce a forecast that is consistent with the Bank of Canada's global macroeconomic outlook.[3] The dynamic response of this model to a variety of shocks (when introduced as part of the Bank of Canada's large US/rest of the world projection model –MUSE/GPM), confirms that disturbances from emerging Asia are important in explaining and forecasting both energy and non-energy commodity prices.

[1] Importantly, this trend is unlikely to reverse in the near term.
[2] See Lalonde, Zhu and Demers (2003).
[3] See Bailliu. J, Blagrave, P. and Rossiter, J. (2010).

The paper proceeds as follows. Section 2 describes the literature related to price formation and forecasting in world commodity markets, and the role of emerging market economies on commodity prices. Section 3 discusses the energy and non-energy commodity price projection models. Section 4 describes the data used and the estimation techniques, while section 5 presents the estimation and out of sample forecast results. Section 6 includes an examination of the model responses to a variety of temporary and permanent shocks, and section 7 concludes.

2. LITERATURE REVIEW

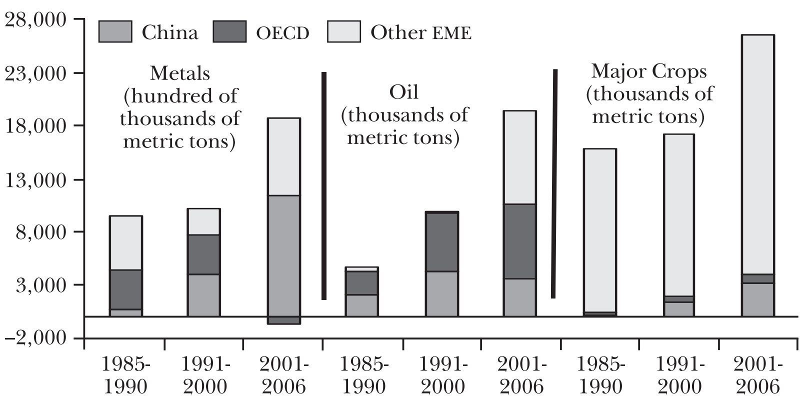
This section discusses two important areas of the literature, namely recent economic research relating to price formation and forecasting in world commodity markets, and the effect of emerging market economies on commodity demand and prices.

A central theme in commodity prices forecasting has been the attempt to disentangle commodity price movements into cyclical and long-term movements. Following a study by Reinhart and Wickham (1994), Borensztein and Reinhart (1994) adopt a structural model to identify the fundamentals behind commodity prices, and conclude that both permanent and transitory shocks contribute substantially to the variation of commodity prices. Cashin, Liang, and McDermott (2000) examine the persistence of shocks to commodity prices. Using International Monetary Fund (IMF) data on 60 individual commodity prices, they find that shocks to most commodity prices are long-lasting and the variability of the persistence is fairly large. Cashin and McDermott (2001) uses much longer sample periods and conclude that there has been a downward trend in real commodity prices over the last 140 years because of relatively faster productivity growth in commodity sectors and a structural change in supply conditions.

Drawing from this literature, Lalonde, Zhu and Demers (2003) develop a model for the projection of non-energy commodity prices at the Bank of Canada. Using structural vector autoregressions to disentangle the permanent from the transitory components, the authors were able to capture the declining trend in real non-energy commodity prices. Their model, however, was not able to capture the intensity effect of emerging market economies.

In fact, this intensity effect is the other important area of the literature. The IMF estimates that annual increases in the global consumption of major commodity groups between 2001 and 2007 were larger than during the 1980s and 1990s, and that a combination of strong income growth, globalization, and rapid population growth have all contributed to the rapid increase in demand (figure 1). Other simple facts also points to the importance of emerging market economies in commodity price movements. For example, China's share in global oil demand increased from 2.7% in 1980 to 9.4% in 2008, while OECD demand share declined from 66% to 56% over the same period. Figures 1 and 2 show that global demand for non-energy and energy commodities was mostly been driven by emerging market economies, and especially China between 2001 and 2008.

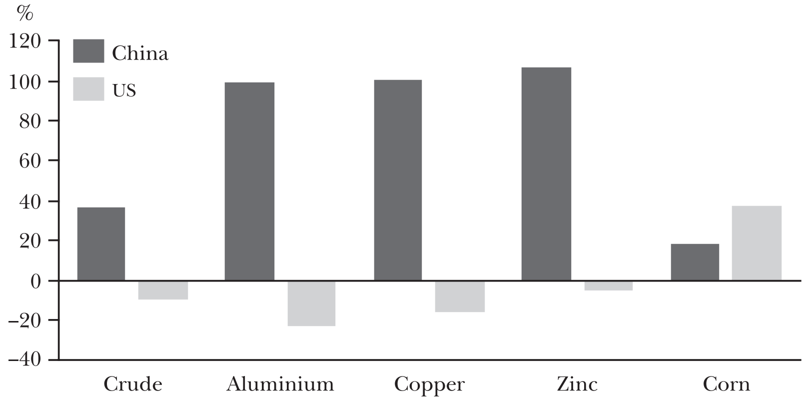
FIGURE 1. COMMODITY DEMAND SINCE 1985, 1985-2006



SOURCE: IMF.

Cheung and Morin (2007), note that China has accounted for a large share of the increase in energy and non-energy commodity demand, which resulted in higher commodity prices. They find that although oil and metals prices have historically moved with the business cycle in developed economies, this relationship broke down around mid-1997. Since then, industrial activity in emerging Asia appears to have

FIGURE 2. CONTRIBUTION TO DEMAND GROWTH FROM 2003 TO 2008
(Selected commodities)



SOURCE: Morgan Stanley.

become a more dominant determinant of oil price movements and the rise in commodity intensity, especially in metals, played an important role in explaining price increases. They also show that commodity consumption per unit of GDP is higher in China, which is consistent with its industrializing and urbanizing development phase. Within a DSGE framework, Elekdag, Lalonde, Laxton, Muir and Pesenti (2007) show that most of the rise in oil prices between 2000 and 2007 can be explained by a combination of permanent productivity and oil intensity shocks in emerging Asia.[4]

The projection models presented in this paper incorporates these two areas of the literature by using an error correction model that includes an explicit role for China's commodity demand and intensity of use while at the same time capturing the impact of global GDP growth on aggregate commodity demand.

3. THE MODEL

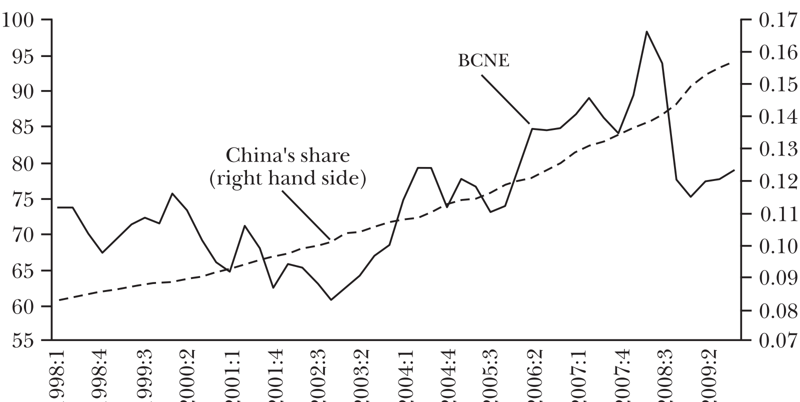
Both the non-energy and energy models are built using an error correction framework (ECM) which relies, in the long run, on the share of China's GDP in the level of world GDP while

[4] A shock on oil intensity imply that the country will use more oil to produce one unit of GDP

short term dynamics are modelled using the potential growth rate of the world and the level of the world output gap. This implies that growth in emerging Asia will affect commodity prices more over the medium and long run than growth in the rest of the World, which can be interpreted as an intensity effect or may represent the transition from an agricultural based economy to a more industrialized one. Given its level of development and composition of GDP, China consumes a greater amount of commodities per unit of GDP than advanced economies.

It's important to note that this effect could not be taken into account by simply including Chinese growth in the dynamic equation and leaving aside the error correction term. Actually, such representation would not capture a permanent increase in commodity demand with constant growth (intensity of use), and therefore would not be able to reproduce the upward sloping trend observed in the data since 1998.[5] As shown in figure 3 and 4, the share of China's economy seems to capture the increasing trend in both energy and non-energy commodity prices.

FIGURE 3. SHARE OF CHINA AND NON-ENERGY COMMODITIES PRICES, 1998-2009



Finally, the strength of the ECM approach is that it allows for transitory shocks coming from economic growth, while the

[5] Moreover, Cheug and Morin (2007) argue that Chinese industrial activity alone cannot explain the increase in metal prices post-1997. One needs to take into account the intensity effect to replicate the data.

FIGURE 4. SHARE OF CHINA AND WTI PRICES, 1998-2009

long-run equilibrium remains anchored by commodity intensity in emerging Asia. (Model 1: equation below):

$$\Delta Non\_energy_t = \alpha(Non\_energy_{t-1} - \beta GDP\_ch\_share_{t-1} - \delta) + \\ + \lambda \Delta Non\_energy_{t-1} + \varphi \Delta Pot\_wld_t + \theta Gap\_wld_t + \varepsilon_t,$$

(1)

$$\Delta WTI_t = \alpha(WTI_{t-1} - \beta GDP\_ch\_share_{t-1} - \delta) + \lambda \Delta WTI_{t-1} \\ + \varphi \Delta Pot\_wld_t + \theta Gap\_wld_t + \varepsilon_t,$$

where *Non-energy* is the relatice price of the Bank of Canada's non-energy commodity price index, *WTI* is the relative price of the West Texas Intermediate price of crude oil, and *GDP_ch_share* is the proportion of China in world GDP (on a PPP basis). *Pot_wld* is the log of the global potential GDP, and *Gap_wld* is the global output gap.[6] Therefore both potential world growth and the world output gap affect oil or non-energy commodity prices with potentially different elasticities. Having the level of the world output gap in the short term dynamic equation also implies that the level of demand, and not the growth of demand, drives commodity prices. Therefore, introducing the level of the output gap in the dynamic equation imply that the response of commodity prices to a temporary shock on world GDP will be positive and

[6] In equation one, the emerging Asian share of global growth is expressed in logs, the world potential is also expressed in logs and is non-stationary, but stationary when differenced. The world output gap is expressed in logs level.

temporary, consistent with properties of a structural model, like BoC-GEM.[7]

On the other hand, one could argue that the output gap and potential growth is not observable and inferred from another model, which could bias estimation. For this reason we also estimate a version of the model that uses global GDP growth (Model 2):

$$\begin{aligned}\Delta Non\_energy_t &= \alpha(Non\_energy_{t-1} - \beta GDP\_ch\_share_{t-1} - \delta) + \\ &+ \lambda \Delta Non\_energy_{t-1} + \varphi \Delta GDP\_wld_t + \varepsilon_t\end{aligned},$$

(2)

$$\begin{aligned}\Delta WTI_t &= \alpha(WTI_{t-1} - \beta GDP\_ch\_share_{t-1} - \delta) + \\ &+ \lambda \Delta WTI_{t-1} + \varphi \Delta GDP\_wld_t + \varepsilon_t\end{aligned},$$

where *GDP_wld* is world GDP growth. This specification, therefore, does not rely on the world potential's estimation by MUSE/GPM, and is thus be more stable over time. However, having GDP growth in the dynamic equation means that a temporary shock on the level of GDP will create a commodity prices response that will increase in the short term, but subsequently decline (given that growth will be negative as the shock fades), which is less consistent with theoretical priors.

While both models give similar results and forecasts, we prefer Model 1 for two reasons. First, the separation of demand factors is essential in order to assess the different elasticities of commodity prices to permanent and temporary shocks to world GDP (see Lalonde and Muir, 2007). Second, this specification is more consistent with the properties of a large global DSGE model.

4. ESTIMATION AND DESCRIPTION OF THE DATA

The macroeconomic data used in the regression models for commodity prices includes data from the Bank of Canada's projection model for the international economy (MUSE/GPM). Variables such as world potential GDP, the world output gap, and the share of China's GDP, are drawn directly

[7] For a more detailed discussion please see the Bank of Canada Technical Report on BoC-GEM (2007).

from internal Bank of Canada MUSE/GPM databases. The dependent variable under investigation is either the real price of crude oil or the real Bank of Canada non-energy commodity price index.[8] Limited data availability only allows for estimation beginning in 1996 using quarterly data. Ideally, when estimating a cointegrating relationship one would use a longer sample. It is important to note, however, that Cheung and Morin (2007), find that the influence of emerging Asian economies on commodity demand only began in the late-nineties, while they use a longer sample.

4.1. Estimation approach

In order to address the potential endogeneity of using both the world potential growth rate and the world output gap contemporaneously, we estimate equation (1) using the generalized method of moments technique (GMM).[9] Given that GMM estimation can potentially suffer from a small sample problem, we also estimated the forecasting model using a non-linear least squares (NLLS) technique. These estimation results show that the error correction term is still significant, with roughly the same estimated coefficient as the GMM estimate.[10] Furthermore, the coefficients on the output gap and potential output growth have the correct signs, are still statistically significant, and are not statistically different from the GMM estimates.[11] However, t-statistic values are lower when the model is estimated with NLLS, thus we focus on the model estimated using GMM.

[8] To obtain real commodity prices, we divide the nominal price of West Texas intermediate (WTI) oil by the US GDP deflator and the Bank of Canada's non-energy commodity price index by the US PPI finished goods index. Also, China's share is computed using PPP exchange rates.

[9] We use lags of the change in either the non-energy commodity price index or the price of oil, world potential growth, and the world output gap as instruments. Note that if the price of oil is used instead of lags of the non-energy index as instruments (for the non-energy model), we find that the results are almost identical.

[10] Model 2, specified using the change in world GDP, is slightly more robust to the estimation approach.

[11] At a significance level between 5 and 10% depending if we include the financial crisis from the sample.

5. ESTIMATION AND FORECAST RESULTS

This section discusses the estimation results for both specifications of the non-energy and energy models. We also present the properties of these models and out of sample forecasting results.

Stock and Watson estimation of the cointegration vector reveals that in the non-energy model the coefficient on the Chinese GDP share is 0.41 with a t-statistic of 5.7, which is adjusted for the long run variance. In the energy model this coefficient is 2.4 with a t-statistic of 8.7 (see table 1). This suggests that oil prices are almost six times more affected by the China's share than non-energy prices, consistent with the fact that the variance of the real price of oil is substantially larger than the variance of the real non-energy commodity prices. In a context of a structural model with fully endogeneous world commodity market like BoC-GEM, this increased volatility can be explained by larger supply and demand rigidities in the oil sector than in the non-energy commodity sector. *Ceteris paribus*, this would imply that the price of oil would be more responsive then non-energy prices to any type of shocks.

TABLE 1. ESTIMATES OF THE ERROR CORRECTION VECTOR[a]

	BCNE		*Oil*		*IMF index*	
	Coef.	*t-stat*	*Coef.*	*t-stat*	*Coef.*	*t-stat*
Constant	6.11	19.0	8.62	13.5	5.91	10.4
Chinese Share	0.41	2.8	2.42	8.6	0.63	2.5

[a] The t-stat presented in that table are using standard errors adjusted for long-run variance.

Tables 2 and 3 show the estimation results for both specifications of the non-energy and oil models.[12] For the first non-energy specification, aside from the third auto-regressive coefficient, the coefficients are statistically different from zero and have the expected sign. In the second model, the inclusion of the world GDP growth rate controls for some of the autocorrelation, which explains why the first and third lags

[12] As expected, the constant of the dynamic equation is statistically insignificant for models 1 and 2, since the trend is captured by the cointegrating vector.

are not significant. All of the coefficients in the energy model are significant and have the expected sign.

TABLE 2. PARAMETER ESTIMATES FOR NON-ENERGY COMMODITY PRICES
Dependent Variable = ΔLog BCNE

Explanatory variable	*Model one*		*Model two*	
	Coef.	*p value*	*Coef.*	*p value*
Speed of Adjustment	–0.29	0.00	–0.13	0.01
ΔLog BCNE (t-1)	0.12	0.17	0.08	0.40
Δ Log BCNE (t-2)	–0.40	0.00	–0.33	0.00
World GDP growth	–		3.87	0.00
Growth of Potential Output	11.88	0.00	–	
Output Gap	1.21	0.01	–	
R^2	0.43		0.41	

TABLE 3. PARAMETER ESTIMATES FOR THE PRICE OF OIL
Dependent Variable = ΔLog (WTI)

Explanatory variable	*Model one*		*Model two*	
	Coef.	*p value*	*Coef.*	*p value*
Speed of Adjustment	–0.394	0.000	–0.321	0.000
ΔLog WTI (t-1)	0.382	0.000	0.321	0.000
ΔLog WTI (t-2)	–0.241	0.032	–	
World GDP growth	–		–	
Growth of Potential Output	24.664	0.013	11.399	0.000
Output Gap	4.865	0.024	–	
R^2	0.505		0.597	

Estimation results also reveal that the error correction terms in both models are highly significant with a speed of adjustment of 0.22 in the non-energy model and 0.39 in the energy model. While these may seem high, it is important to note that the persistence of the output gap slows the adjustment process of the model as a whole. This is confirmed when looking at the results of the second specification in which the speed of adjustment is lower as a result of a lower degree of persistence in GDP growth. Furthermore, autocorrelation tests show that the lag structures for both specifications in the non-energy and oil models are well defined. Finally, an R-squared of 43% and 51% for the first difference of non-energy commodity prices and oil prices respectively, shows that the models seem to fit the data reasonably well.

Tables 4 and 5 show the results from the alternative specifications, which include the US real federal funds rate and the rate of change in the US real effective exchange rate. One could argue that interest rates may affect commodity prices given that it is a proxy for capital costs, and that exchange rate movements could affect prices because most commodities are expressed in US dollar. However, these variables are not significantly different from zero, and do not improve the fit of the BCNE model but may add some information in the oil model.

TABLE 4. ALTERNATIVE SPECIFICATIONS WITH BCNE

Dependent Variable = ΔLog BCNE

Variable	*Model one*		*Model two*	
	Coef.	*p value*	*Coef.*	*p value*
Speed of Adjustment	–0.27	0.01	–0.16	0.00
ΔLog BCNE (t-1)	0.11	0.21	0.02	0.89
ΔLog BCNE (t-2)	–0.42	0.00	–0.30	0.00
World GDP Growth	–		4.81	0.00
Growth of Potential Output	11.71	0.00	–	
Output Gap	1.28	0.00	–	
Real US Federal Funds Rate	–0.14	0.51	–0.15	0.54
Δ US Real Effective Exchange Rate	–0.12	0.68	0.55	0.16
R^2	0.44		0.39	

TABLE 5. ALTERNATIVE SPECIFICATION WITH OIL PRICES

Dependent Variable = ΔLog (WTI)

Variable	*Model one*		*Model two*	
	Coef.	*p value*	*Coef.*	*p value*
Speed of Adjustment	–0.30	0.00	–0.30	0.00
ΔLog WTI (t-1)	0.32	0.00	0.30	0.00
ΔLog WTI (t-2)	–0.33	0.00	–	
World GDP growth	–		13.10	0.00
Growth of Potential Output	45.62	0.00	–	
Output Gap	2.23	0.11	–	
Real US Federal Funds Rate	1.05	0.04	–0.14	0.76
Δ US Real Effective Exchange Rate	0.44	0.41	1.58	0.02
R^2	0.48		0.51	

5.1. Out-of-sample forecasting results and dynamic forecasts

Although the fit of both models are reasonable, particularly given their parsimonious specifications, we also conduct dynamic simulations and out-of-sample forecasting exercises. We compute dynamic forecasts and compare them to a random walk. We estimate the dynamic equations from 1996 to 2005 and compute rolling forecasts from 2005 to 2009. Note that for every sub-period we re-estimate the specification using the latest available information. Tables 6 and 7 report the root mean squared errors (RMSE).

TABLE 6. OUT OF SAMPLE ROOT MEAN SQUARED ERRORS – BCNE

	T+1	*T+2*	*T+4*	*T+8*
Model 1	0.045	0.066	0.065	0.082
Model 2	0.042	0.075	0.101	0.136
Random Walk	0.061	0.100	0.115	0.125

TABLE 7. OIL MODEL OUT OF SAMPLE FORECAST RESULTS

Model 1	0.254	0.341	0.282	0.228
Model 2	0.189	0.330	0.270	0.297
Random Walk	0.231	0.366	0.431	0.306

Despite the short sample, the energy and non-energy models seem to be stable enough to beat or match a random walk.[13] This shows those equations can provide some structural framework to analyze and forecast movements in commodity prices, without hurting the forecasting performance when compared to a naïve forecast.

When we estimate the first specification (that uses the level of the world output gap and the growth of world potential GDP) with the entire sample and begin the forecasting exercise at different points in time, we see that the models do a reasonably good job of tracking commodity price movements over the last few years (figures 3 and 4), even at the eight quarter ahead horizon. It is interesting to note, however, that

[13] It should be noted that the first specification does seem to perform better than the second one as the forecast horizon increases, for both non-energy and oil prices.

even when estimated using the full data set, the models cannot perfectly replicate the large increases in commodity prices observed in 2008. That said, the model is generally able to track the broad increase and subsequent decline in prices observed recently.

Finally, as a check for robustness, we conducted a variety of alternative estimations and found that:

1. If we exclude the recent financial crisis period (2008Q3-2009Q2), the estimation results remain relatively unchanged.

2. The results are robust when we compare estimates from specification 1 (which uses global potential output and the level of the global output gap) to estimates from specification 2 (which includes the change in global output).

3. The estimation results are not statistically different when estimated with the GMM technique or an instrumental nonlinear least squares approach.

4. Finally, the estimation results are robust to the use of the International Monetary Fund's measure of non-energy commodity prices instead of the Bank of Canada's non-energy commodity price index (see table 8). Thus, the role of emerging Asia is important across alternative commodity weighting schemes, and for both energy and non-energy commodities.

TABLE 8. ESTIMATION RESULTS USING THE IMF NON-ENERGY COMMODITY PRICE INDEX

Dependent Variable = ΔLog IMF index

	Model one		*Model two*	
Variable	*Coef.*	*p value*	*Coef.*	*p value*
Speed of Adjustment	–0.14	0.00	–0.08	0.09
IMF index (t-1)	0.14	0.34	0.14	0.14
IMF index (t-2)	–0.51	0.00	–0.25	0.00
IMF index (t-3)	–		–0.02	0.87
World GDP growth	–		5.83	0.00
Growth of Potential Output	11.19	0.00	–	
Output Gap	3.12	0.00	–	
R^2	0.49		0.56	

6. SHOCK ANALYSIS

To summarize the dynamic properties of the two commodity projection models (oil and non-energy) we introduce the commodity projection models in the Bank of Canada's large multi-region projection model (MUSE/GPM) and report the response of both oil and non-energy commodity prices to various shocks in specific economic regions. Several simulations help to illustrate the dynamic behavior of the models. The purpose of this exercise is twofold. First, it allows us to describe the dynamic responses of commodity prices given a specific shock, which is informative in the context of projection. One can therefore assess the length and the magnitude of commodity prices responses to demand shocks. Second, it can compute elasticities from shocks arising from specific regions that are not directly included in the equations. For example, we cannot simply shock China's potential in our model because we need to capture the world economy's reaction in order to assess its effect on commodity prices. By including our equations in MUSE/GPM, which takes several regions into account, we can estimate those elasticities, and draw simple rule of thumb from them.

We only report impulse responses from the first specification, which is our preferred one. We focus on the effects of the following two shocks scaled equivalently for comparison purposes:

1. *Temporary* output shocks in the US and China equivalent to 1% of global GDP.

2. *A permanent* increase in output in the US and China equivalent to 1% of global GDP.

6.1. Shock 1: A temporary demand shock

In order to examine the importance of different levels of commodity intensity between regions, we simulate a temporary demand shock (i.e., a shock on the output gap) equivalent to one percent of global GDP originating from either the US or China (results are summarized in tables 9 and 10 and figures 5 and 6).

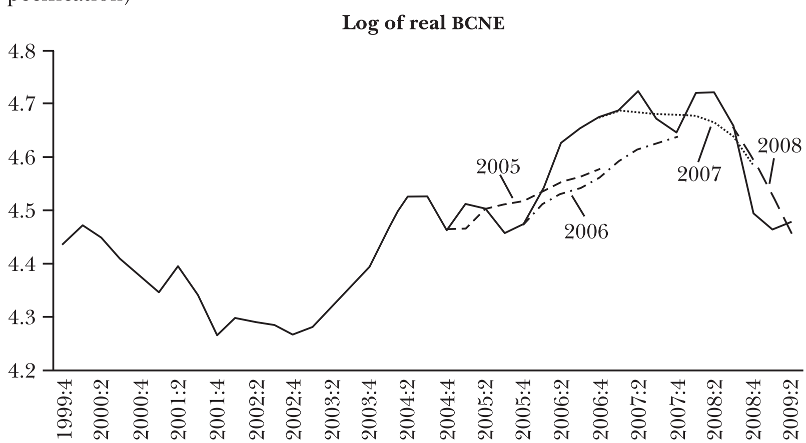
TABLE 9. PEAK RESPONSE OF COMMODITY PRICES TO A TEMPORARY DEMAND SHOCK

When the shock originates in...	*Non-energy (%)*	*Oil (%)*
China	3	20
US	2.8	10

TABLE 10. PEAK RESPONSE OF COMMODITY PRICES TO A PERMANENT DEMAND SHOCK

When the shock originates in...	*Non-Energy (%)*	*Oil (%)*
China	8	27
US	4.5	10

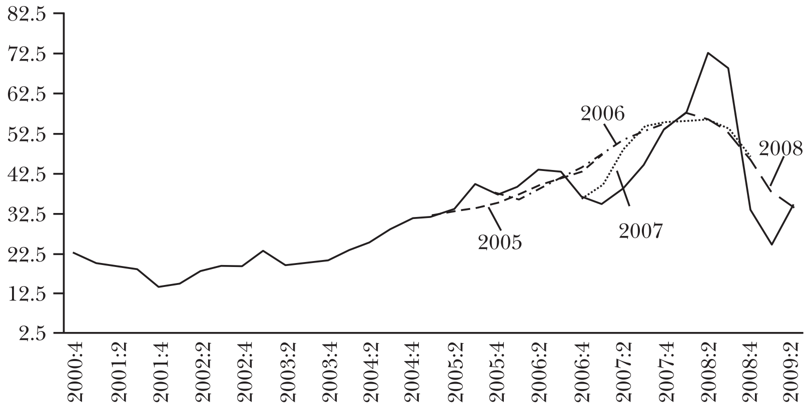
FIGURE 5. NON-ENERGY MODEL DYNAMIC FORECASTS, 1999-2009 (First especification)



We observe a much larger effect on oil prices when the demand shock originates in China, whereas the magnitude of the shock on non-energy commodity prices is similar if the shock originates in either the US or China. The large difference between the two oil price responses is due to the fact that the model not only controls for the demand for commodities, but also the intensity of use. Since China uses more energy per unit of GDP, a temporary shock on Chinese growth will have more impact on prices. As a result, the response from the same shock originating in the US creates in a smaller but longer-lasting increase in commodity prices. This is explained by the fact that in MUSE/GPM the US output gap is more persistent than the Chinese output gap. Furthermore,

FIGURE 6. OIL MODEL DYNAMIC FORECASTS, 2000-09 (First especification)

Real oil price



the difference in response between oil and non-energy commodities is explained by the speed of adjustment parameters in the two ECMs. This parameter is much larger in the oil price model, implying that a deviation of oil prices away from their long-run desired level (which is determined by the share of China in world GDP), is corrected faster as compared to a deviation in the price of non-energy commodity prices. The oscillation of both commodity prices, when the shock originates in China, is driven by the higher degree of variability in China's economy relative to other countries as a result of its fixed exchange rate regime.[14]

6.2. Shock 2: A permanent demand shock

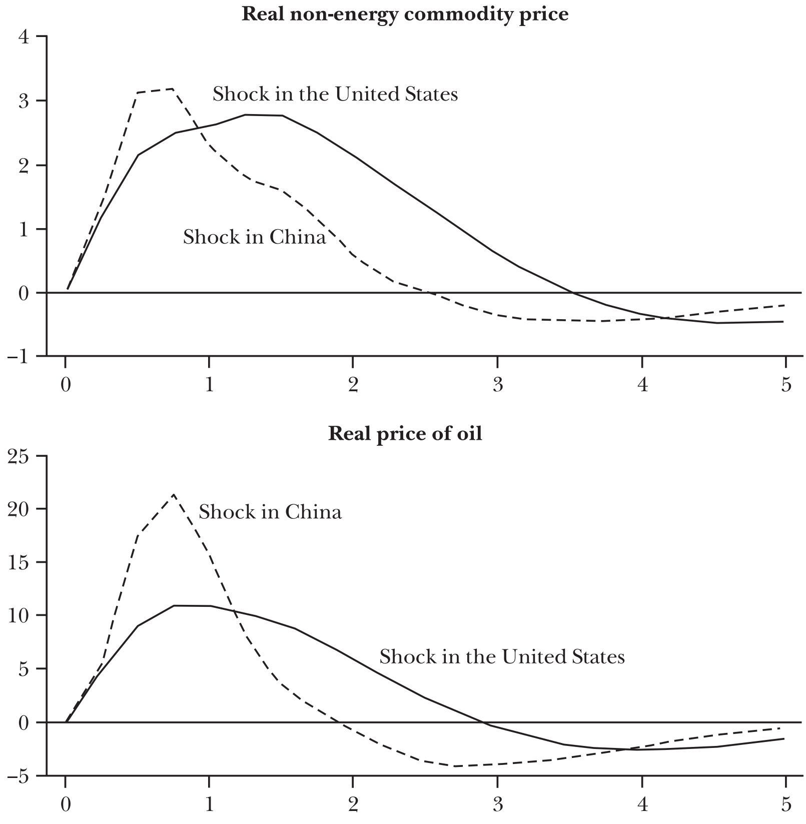
A permanent increase in potential output has a larger effect on oil and non-energy commodity prices when the shock originates in China because of the higher commodity intensity of its production (table 2). The hump-shaped response can be explained by the slow and gradual adaptation of supply to the shock.[15] Real rigidities limit the ability for production to adjust immediately to the permanent demand shock. As supply gradually adjusts, prices tend to fall, though remaining

[14] This result comes from the response of MUSE/GPM to a temporary demand shock.

[15] Note that this hump-shape is also present after a temporary demand shock, but this is due to the fall in demand as the shock fades.

permanently higher than prior to the shock. Note, however, the absence of a positive permanent effect when the shock originates in the US. This steams from the decline in China's share in the error correction vector in the long run. This result highlights a shortcoming of the model.

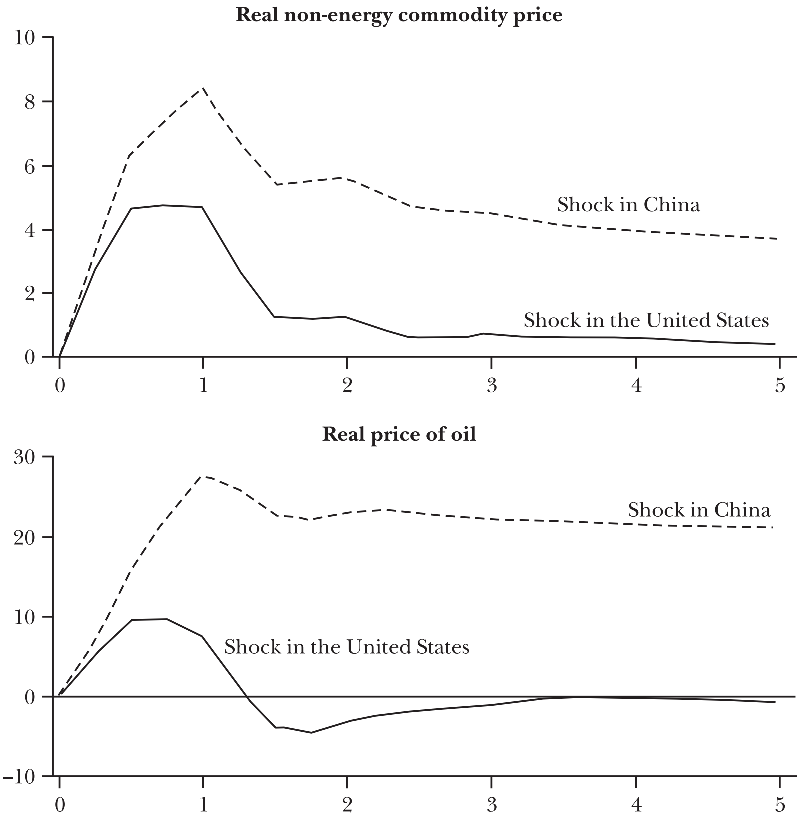
FIGURE 7. TEMPORARY SHOCK TO THE OUTPUT GAP, EQUIVALENT TO ONE PERCENT OF GLOBAL GDP, (Shock minus control, in percent, years)



7. CONCLUSIONS

The models presented in this paper provide a framework that links the projection of commodity prices with the Bank of Canada's global forecast. We also describe the properties of the model by examining a series of dynamic responses when the model is introduced into a larger model (MUSE/GPM). We assume that the factors behind the movements of commodity prices since the mid 1990s, namely strong growth in commodity intensive regions like China, will also be the key factor driving commodity prices in the coming years.

FIGURE 8. PERMANENT SHOCK TO POTENTIAL OUTPUT EQUIVALENT TO ONE PERCENT OF GLOBAL GDP, (Shock minus control, in percent, years)



Because of data limitations and because emerging Asian countries only started to play a significant role in commodity markets in the later part of the 1990s, we are constrained to use a small sample size. This can reduce the power of estimation in a cointegration model.

Another limitation of the model is the lack of explicit supply variables. While it would be ideal to control for those factors inherent in many commodity markets, the non-energy commodity index includes a diverse set of 20 commodities, and thus 20 supply-side factors. Thus, modeling the supply side would be extremely difficult. For oil, modeling supply conditions is challenging as full disclosure of supply is not normal among OPEC members. Therefore, modeling oil supply is made more difficult by the challenge of modeling OPEC decisions.

Also, there may be other omitted variables that could affect commodity prices, such as financial factors (i.e., investor interest). In this regard, future work could complement the model with a common factor approach that would attempt to disentangle demand factors from commodity specific supply factors.

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ISSN: 0187-7615