CENTRO DE ESTUDIOS MONETARIOS LATINOAMERICANOS

2003

MONEY AFFAIRS

VOLUME XVI, NUMBER 1, JANUARY-JUNE

MONEY AFFAIRS is a bi-yearly publication of the Centre for Latin American Monetary Studies (CEMLA), Durango nº 54, Mexico City, D. F., 06700. ISSN-0187-7615.

MONEY AFFAIRS is regularly listed in the International Current Awareness Service: Economics. Selected material is indexed in the International Bibliography of Economics.

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Forecasting and analyzing world commodity prices

1. INTRODUCTION

The resource sector has traditionally played an important role in the Canadian economy, especially in the area of foreign trade. Over the past decade, total exports of commodities represent, on average, about forty one per cent of Canada's exports of goods[1] and fifteen per cent of Canada's gross domestic product (GDP). Consequently, changes in world commodity prices have histori-

[1] This ratio is defined as the share of nominal commodity exports in total nominal exports over the period 1990 to 2001.

Paper prepared by R. Lalonde, Principal Researcher International Department, Z. Zhu and F. Demers, Economists Research Department, Bank of Canada. The authors would like to thank seminar participants at the Bank of Canada for useful comments. Special thanks also to Donald Coletti, Allan Crawford, Simon van Norden and Jaime Marquez. The earlier version of this paper was also presented at the VII Meeting of the Central Banks Research Network of the Americas hosted by the Central Bank of Guatemala, November 22-24, 2002. The views expressed in this paper are those of the authors. No responsibility for them should be attributed to the Bank of Canada. *Corresponding author. Contact address: 234 Wellington St. Ottawa, ON, K1A 0G9. E-mail: rlalonde@bank-banque-canada.ca.*

cally been a key determinant of Canada's terms of trade, which in turn have affected the real income of Canadians.

The staff at the Bank of Canada (BOC) have designed the Bank of Canada Commodity Price Index (BCPI) to track the prices paid for key Canadian commodities. The BCPI is a fixed- weighted index of the spot or transaction prices of 23 commodities produced in Canada and sold in world markets.[2] All components of the BCPI are priced in U.S. dollars. The choice of commodities is determined by their importance in Canadian production, subject to limitations imposed by data availability. For the purpose of this paper, we only focus on two subcomponents of the BCPI: non-energy commodity prices (BCNE) and the West Texas Intermediate crude oil price (WTI). To obtain real commodity prices, we divide by the U.S. GDP deflator.

In this paper, we employ two different empirical approaches to model commodity prices. For real BCNE prices, we use an approach that combines a structural vector autoregressive model (SVAR) with a single equation model. The SVAR is used to give us a historical decomposition of movements in real BCNE prices, and to project the permanent (or long-run) component of prices, while the transitory (or short-run) component of real BCNE prices is forecasted with a single equation model. We find that this approach successfully captures the strong linkage of real BCNE prices with the world economic activity and the real U.S. effective exchange rate in the short run.[3] A 1% positive shock to world economic activity leads to an approximately 6% peak response of real BCNE prices, while the response to the real U.S. dollar effective exchange rate shock is small but it is statistically significant and exhibits the expected sign.[4] We also find that the variance of the transitory component of real BCNE prices accounts for approximately 60% of total real BCNE price variance. This result is consistent with numerous other studies of commod-

[2] See Appendix A for a description of the BCPI and its components. The weight of each commodity in the index is based on the average value of Canadian production of the commodity over the 1982-90 period.

[3] The U.S. dollar effective exchange rate is defined as a U.S. export weighted average of the exchange rate of U.S. dollar relative to the currencies of Japan (17.59%), U.K. (8.22%), Mexico (14.52%), Canada (35.40%) and Euro zone (24.27%).

[4] We use the world output gap as a proxy for the overall world economic activity, where the output gap is generated using the Hodrick-Prescott (HP) filter.

ity prices.[5] In terms of out-of-sample forecasting, our approach outperforms a VAR model and an autoregressive (AR) model.

For real crude oil prices, we use a statistical multiple structural-break approach to identify significant shifts in OPEC behavior. After controlling for these mean shifts, we find a very strong role for world economic activity in the determination of oil prices. We estimate that a 1% positive shock to world economic activity leads to an approximately 12% peak response of real crude oil prices with a lag of two to three quarters. In terms of out-of-sample forecasts, the real oil price model outperforms an AR model and a random walk (RW) model.

The remainder of the paper is organized as follows. Section 2 gives a brief overview of recent literature on commodity prices. In Section 3, we present the methodology as well as results for the real BCNE price model. The methodologies and results for the real crude oil price model are discussed in Section 4. Section 5 concludes.

2. LITERATURE REVIEW

This section gives a brief overview of some of the recent economic literature pertaining to price formation in world commodity markets. A common theme in many of these studies has been an attempt to disentangle commodity price movements into a cyclical and a long-term movement. This distinction is important for forecasting commodity prices both in the short-run and long-run.

Various methodologies have been used in order to disentangle the trend movement of world commodity prices from the cycle. Reinhart and Wickham (1994) apply two different approaches, namely the Beveridge and Nelson (1981) technique and the Harvey (1985) approach. The first approach is a pure reduced form time-series technique used for the decomposition of a time-series variable, while the second one is a structural time-series approach using the Kalman filter. Each of these two approaches has its own strengths and weaknesses. The pure mechanical filters can easily split a time series into cyclical and permanent components, but lack economic fundamentals. Although the Kalman filter contains certain economic information, it often does not perform very well in practice if the assumptions of normal distributions for disturbances and the initial state vector are violated. When the normal-

[5] For example, Borensztein and Reinhart (1994) obtain a similar result.

ity assumption is dropped, there is no longer a guarantee that the Kalman filter will give the conditional mean of the state vector, i.e. the estimates of the state vector could be conditionally biased.[6] Moreover, it becomes more cumbersome to calculate the likelihood function without the normality assumption.

Following the study of Reinhart and Wickham (1994), Borensztein and Reinhart (1994) adopt a structural model to identify the key fundamentals behind commodity prices, and more importantly to quantify the relative contributions of demand and supply shocks. On the demand side, they find that the real U.S. dollar effective exchange rate and the state of the business cycle in industrial countries are closely linked to the cyclical movement of world commodity prices. On the supply side, strong productivity growth of commodity sectors relative to the rest of the economy and the increased commodity supply relative to the rest of the economy are the primary causes of the downward trend of commodity prices. Using a variance decomposition, the authors conclude that both types of shocks contribute to the total variation of commodity prices in the near term and around 60% of the variation is caused by demand shocks.

Cashin, Liang and McDermott (2000) examine the persistence of shocks to commodity prices. They use a median-unbiased estimation procedure proposed by Andrews (1993) instead of a unit root test to check the persistence of shocks. Using IMF data on sixty individual commodity prices, they find that shocks to most commodity prices are long-lasting (reflected by the high value of the half-life of a unit shock), and the variability of the persistence is fairly large. Cashin and McDermott (2001) use much longer sample periods and examine whether the long-run behavior of commodity prices has changed. In particular, they look at the trend of most commodity prices, the duration of price booms and slumps, and also the volatility of price movements. They apply various statistical tests and compare the patterns of commodity price movements across different sample periods. The authors come to the conclusion that there has been an apparent downward trend in real commodity prices over the last 140 years because of relative productivity growth in commodity sectors and a structural change in supply conditions.[7] Moreover, the short-term volatility is highly related to the business cycle.

[6] See Harvey (1989) for details.

[7] Coletti (1992) examines a small set of non-energy commodities that mainly include industrial materials (e.g. metals, minerals and forest products) over the

In practice, numerous methodologies have been employed to disentangle transitory and permanent movements in commodity prices. Though convenient to apply, pure time-series filters suffer from a lack of structural economic fundamentals. In contrast, although structural models are constructed based on the economic theory, they are often costly and time-consuming to develop and maintain. For instance, it would be very costly to develop and maintain models for 23 individual components of the BCPI. Therefore, as a compromise, we combine basic time series approaches with simple economic theory to develop econometric models for the two major BOC commodity price indexes.

3. THE REAL BCNE PRICE MODEL

This section consists of three subsections. The first two parts describe the methodology used to identify and to forecast the transitory and permanent components of real BCNE prices. The last subsection presents the results.

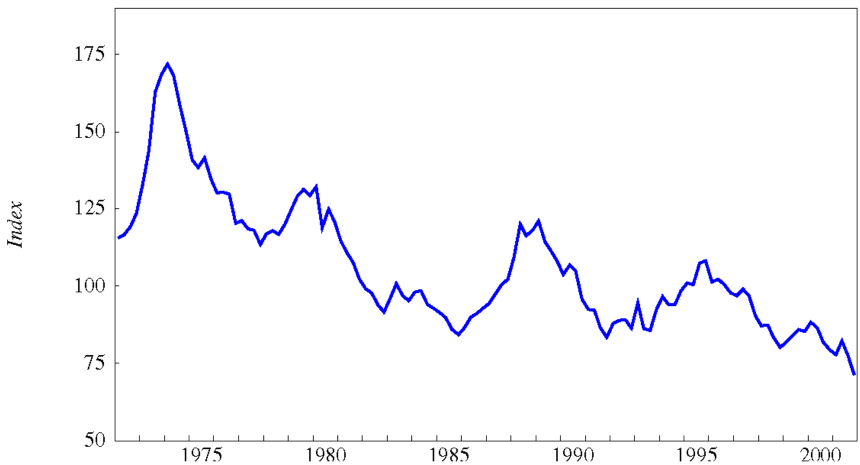
3.1 Identifying the transitory and permanent components of real BCNE prices

Figure 1 plots the evolution of real BCNE prices over the sample period. It is evident that the non-energy commodity price index has experienced a downward trend, a consistent finding with other studies on primary commodity prices mentioned earlier. This characteristic could be largely explained by the rise in relative productivity in commodity-producing sectors as well as the increase in supply conditions of most primary commodities over the past three decades. Furthermore, we also notice that the series exhibits substantial variations around trend over the course of business cycles. We use a SVAR approach to decompose historical BCNE prices into transitory and permanent components. Under this approach, a number of economic restrictions are imposed on the long-run effects of different types of shocks. The main strength of the SVAR methodology is that one does not have to impose a fully specified theoretical structure and the data are allowed to speak. The only assumptions are that the variable

1900-91 period. He finds no obvious secular decline in relative prices of those commodities.

of interest (i.e. real BCNE prices) can be decomposed into one or more permanent components and one or more transitory components, and that the transitory shocks are uncorrelated with the permanent shocks. However, the SVAR methodology has its own weakness. Notably, the results are often sensitive to the choice of variables included in the estimation. Also, results can be affected by the number of lags chosen in the reduced form, assumptions on the order of integration of variables, and the presence of cointegrating relationships among variables.

FIGURE 1: REAL BCNE PRICES, 1975-2000



In our model, variable selection is based on economic theory and the findings of previous studies. To capture the information about transitory shifts in real BCNE prices arising from changes in world economic conditions, we use the G7 output gap as a proxy for the world economic activity.[8] The G7 inflation rate,[9] a proxy for the global inflation rate, is added to capture the importance of

[8] The G7 output gap is generated using the SVAR methodology for the U.S. (see Lalonde (1998)) and the HP filter for the rest of G7 countries. We take the sum of individual output gaps weighted by each country's share in the composition of the G7 output evaluated at purchasing power parity. We use the term "world output gap" through the rest of the paper.

[9] The G7 inflation rate is generated by taking the sum of individual inflation rates weighted by each country's share in the composition of the G7 output evaluated at purchasing power parity. We use the term "global inflation rate" through the rest of the paper.

having a nominal anchor in the model as suggested by the SVAR literature.[10] In light of the empirical studies on world commodity prices in the previous section, we include two additional demand indicators - the real U.S. long-term interest rate as a proxy for the real world interest rate, and the real U.S. dollar effective exchange rate - to identify the cyclical component of real BCNE prices.[11] In addition, we have attempted to include some supply-side determinants of the permanent component of real BCNE prices. However, given the fact that the real BCNE price is an aggregate price index, it is hard to find a proper measure of productivity.[12]

The final SVAR contains the following five variables: real BCNE prices (*Rbcne*), the world output gap (*Wygap*), the global inflation rate ($W\pi$), the real U.S. long term interest rate (*RRus*) and the real U.S. dollar effective exchange rate (*Erus*). We assume that the real U.S. interest rate is stationary in levels.[13] ADF tests show that the world output gap is stationary in levels and the rest of variables are first difference stationary. Furthermore, a Johansen co-integration test shows that there is no co-integrating relationship between *Rbcne*, $W\pi$ and *Erus*. The technical details on the SVAR methodology are presented in Appendix B. We estimate the model over the period of 1972-2001 using quarterly data.[14]

[10] If monetary policy has a neutral effect across different sectors of the economy, both in the short-run and long run, the presence of the global inflation rate in the model may not be important. However, because monetary policy may not affect all sectors in the same manner in the short-run, it can have a transitory effect on relative prices. Consequently, using real BCNE prices alone may not be sufficient to purge the effects of monetary policy. Out-of-sample forecasting performance of the model including the global inflation rate is slightly better than the one which excludes it. Furthermore, results show that real BCNE prices do react, in the short-run, to a shock affecting the trend inflation rate.

[11] Since world commodities are all priced in U.S. dollars, movements in the real U.S. dollar effective exchange rate will affect the demand for commodities by countries other than the U.S. This in turn will affect prices.

[12] If the productivity growth only happens in a particular sector, this tends to lower production cost in this sector relative to the rest of the economy. Consequently, this causes lower prices of goods produced in this sector relative to the aggregate level (i.e. lower relative prices).

[13] The augmented Dickey-Fuller (ADF) test provides ambiguous evidence regarding the stationarity of the real U.S. long-term interest rate. However, we assume here that it is stationary. The results are robust to this assumption.

[14] We estimate the same model over the sample of 1972-95 and we find that the transitory component of real BCNE prices is almost identical to the one estimated over the full sample period.

3.2 Forecasting the transitory and permanent component of real BCNE prices

The second step of the approach consists of finding the best way to produce forecasts of both the temporary and permanent components that are not only tractable but also consistent with projections of the rest of the world economy. We use the VAR to forecast the permanent component of real BCNE prices and develop a single equation model to forecast the transitory component. This single equation links the transitory component of real BCNE prices to the world output gap as well as the real U.S. effective exchange rate gap.[15] This equation is defined as:

Regression equation:

$$Rbcnegap_t = A(L)Rbcnegap_{t-1} + B(L)Wygap_t + C(L)Ergap_t,^{16} \qquad (1)$$

where *Rbcnegap* is the transitory component of real BCNE prices (i.e. real BCNE price minus the SVAR estimates of its permanent component), *Wygap* is the world output gap and *Ergap* is the real U.S. dollar effective exchange rate gap. This equation has the advantage of relying on a small number of estimated parameters, which helps to reduce out-of-sample forecasting errors. In addition, it clearly quantifies the impact of the world output gap as well as the real U.S. dollar exchange rate gap on the change in the transitory component of real BCNE prices.

3.3 Results of real BCNE price model

This section presents results of the real BCNE price model. First, we use variance decomposition to quantify the relative importance of supply and demand shocks. Second, we discuss the link between the world output gap and the transitory component of real BCNE prices. The last part evaluates the forecasting performance of the model.

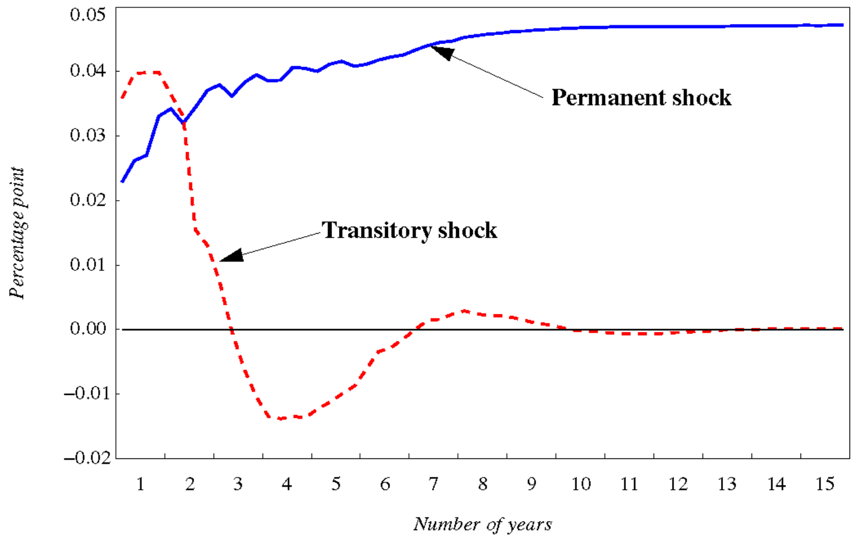
3.3.1 The relative importance of supply and demand shocks

Table 1 reports the variance decomposition of real BCNE prices

[15] We use the HP filter to generate the real U.S. dollar exchange rate gap.

[16] It is worth noting here that the real U.S. interest rate is excluded in the equation due primarily to its strong collinearity with the world output gap, but it can still indirectly affect the forecast of real BCNE prices via its impact on the forecast of the world output gap (i.e. through the forecast of the U.S. output gap).

FIGURE 2: IMPULSE RESPONSE OF REAL BCNE PRICES



at different time horizons. After the first year (step = 4), the transitory shocks (i.e. demand shocks) explain almost 60% of the total variance of real BCNE prices. After two years (step = 8), however, the contribution of demand shocks falls dramatically and accounts for only 10% of the total. The model shows a significant contribution of demand shocks to real BCNE prices in the short term, and this is consistent with other studies mentioned earlier. Figure 2 plots the corresponding impulse responses of real BCNE prices to a positive one standard deviation *total* transitory and permanent shock.[17] Real BCNE prices exhibit a small hump-shaped response to the total transitory shock while the response to the permanent shock appears to be more gradual.

3.3.2 The world output gap and the transitory component of real BCNE prices[18]

Figure 3 plots the evolution of both the world output gap and the transitory component of real BCNE prices over the historical

[17] The total transitory shock is defined as the aggregation of four individual transitory shocks presented in the SVAR model. See appendix A for details.

[18] As a robustness checking, we also try the U.S. output gap. We find that models with the world output gap outperform those with the U.S. output gap in most cases. All results for the U.S. output gap model are available upon request.

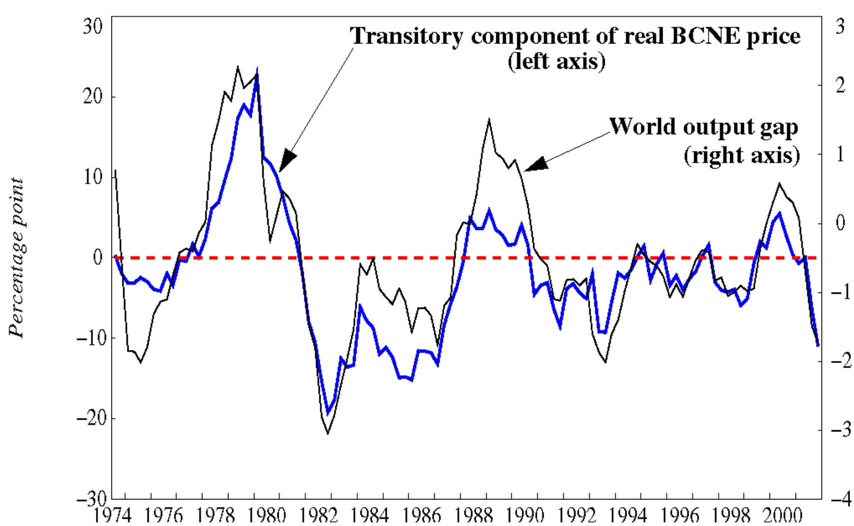
TABLE 1: VARIANCE DECOMPOSITION OF THE REAL BCNE PRICE (in percent)

Step	*Supply shock*	*Demand shock*
1	29%	71%
4	41%	59%
8	89%	11%
16	90%	10%
∞	100%	0%

period. There is a strong positive relationship between the two variables. The world output gap tracks most of the important cyclical movements of real BCNE prices since the mid 1970s.

Table 2 reports the parameter estimates of equation (1). The Hausman test fails to reject the null hypothesis of exogeneity of the world output gap, and hence we use instrumental variable estimation (IVE). The instruments used for the estimation are four lags of the world output gap, the transitory component of real BCNE prices, the real U.S. dollar exchange rate gap and the real U.S. long-term interest rate. The standard errors of the estimated parameters are modified using an 8-lag Newey-West correction. We start with eight lags for each regressor, and then remove the most insignificant estimates one by one until all the remaining coefficients are statistically significant at the 5% significance level.

FIGURE 3: WORLD OUTPUT GAP AND THE TRANSITORY COMPONENT OF REAL BCNE PRICES, 1974-2000



All the coefficient estimates have the expected signs and are statistically significant in the final model. The transitory component of real BCNE prices itself is fairly persistent with a root of about 0.71, and both the world output gap and the real U.S. dollar effective exchange rate gap contribute significantly to transitory movements of real BCNE prices. Furthermore, we calculate the relative contributions of a positive one standard deviation shock to each regressor in our model to the total response of the real BCNE transitory component. We find that around 80% of the total response comes from shocks to the world output gap (72%) and the real U.S. dollar exchange rate gap (8%). In other words, only a small fraction (20%) of the response is left unexplained by our model.

TABLE 2: FORECASTING EQUATION OF THE TRANSITORY COMPONENT REAL BCNE PRICES, R*BCNEGAP* (IV ESTIMATION OF EQUATION (1))

Variable	*Coefficient*
Rbcnegapt-1	0.711 (15.79)
Wygapt	0.059 (7.34)
Wygapt-1	-0.041 (-5.24)
Ergapt	-0.002 (-2.54)

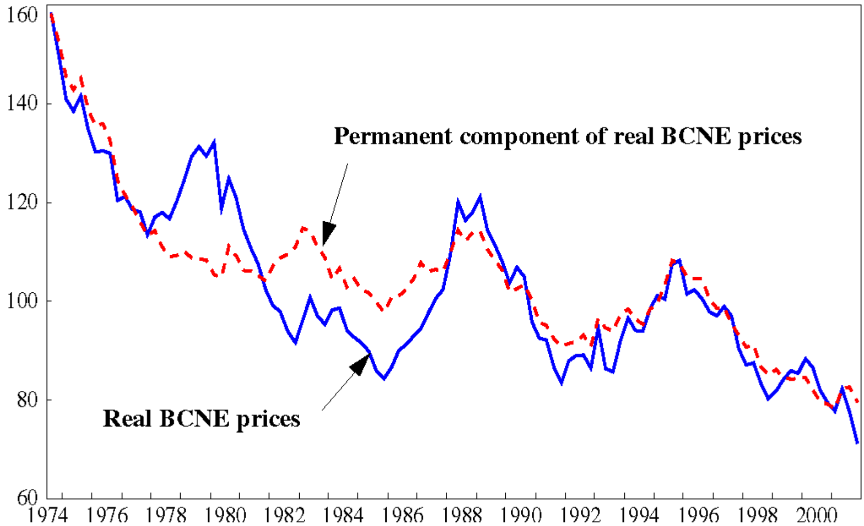
NOTE: *t*-statistics are in parentheses.

Figure 4 plots the real BCNE price vis-a-vis its permanent component over history. The implied cyclical movements of real BCNE prices are consistent with our priors. Figures 5 and 6 show the responses of real BCNE prices to a 1% positive shock to the world output gap and the real U.S. dollar effective exchange rate gap respectively. The peak response of real BCNE prices to the world output gap shock is about 6% and occurs almost contemporaneously with the peak of the world output gap itself. In comparison, the response to a shock to the real U.S. dollar effective exchange rate gap is much smaller, with a peak of about -0.35%, but it exhibits the expected sign.

3.3.3 Out-of-sample forecast of real BCNE prices

World commodity price shocks have a peak impact on the core

FIGURE 4: REAL BCNE PRICES AND ITS PERMANENT COMPONENT, 1974-2000



CPI inflation rate with a lag of two to four quarters in the Canadian economic projection model used at the Bank of Canada. Since monetary policy tends to have its full impact on inflation with a lag of six to eight quarters, the monetary authority will be most interested in forecasts of world commodity prices two to four quarters ahead.

FIGURE 5: SIMULATION OF A ONE PERCENTAGE POINT POSITIVE SHOCK TO WORLD OUTPUT GAP

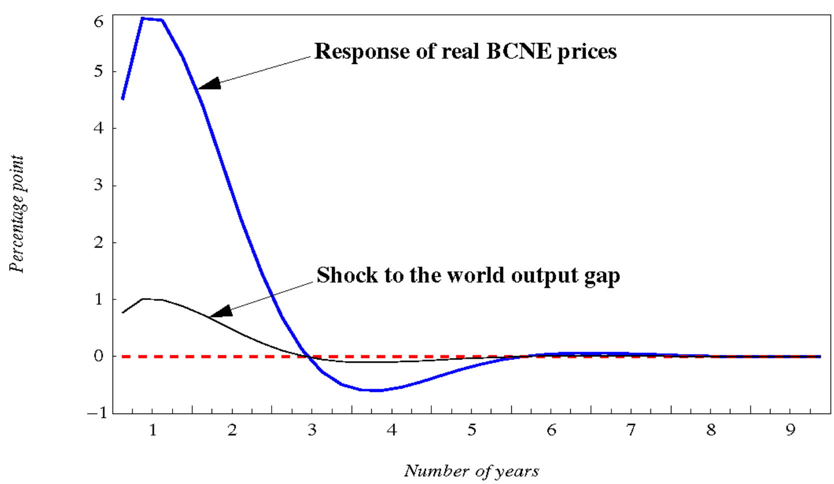
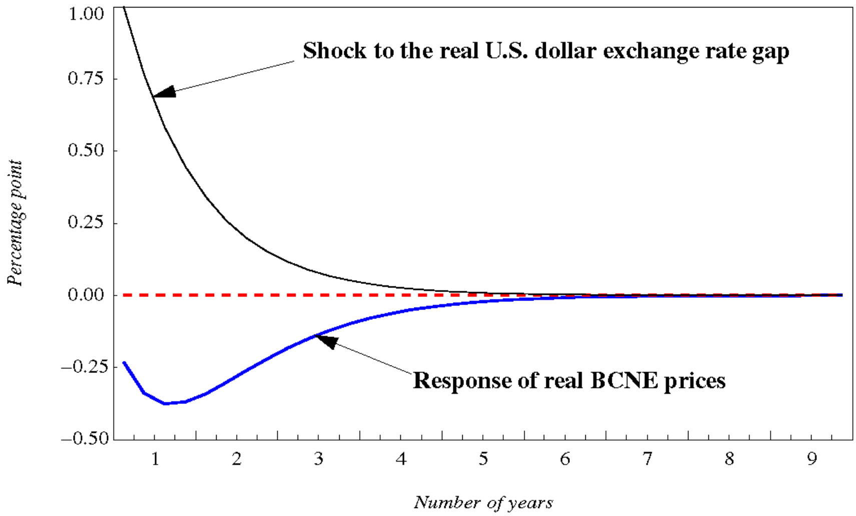


FIGURE 6: SIMULATION OF A ONE PERCENTAGE POINT POSITIVE SHOCK TO THE REAL U.S. DOLLAR EFFECTIVE EXCHANGE RATE GAP



We evaluate our model's out-of-sample forecasting performance by comparing it with forecasts from two benchmark models: VAR model and AR(1) model. Our model forecasts of real BCNE prices combine SVAR forecasts of the permanent component and single equation forecasts of the transitory component. As mentioned earlier, we focus on the forecasting horizon which is of interest to the monetary authority, namely two to four-quarters ahead. According to the RMSE of out-of-sample forecasts from 1992q1 to 2001q4,[19] the combined approach uniformly outperforms the two benchmark models regardless of the forecasting horizons (see Table 3) according to smaller values of RMSE.[20]

[19] We use a rolling sample regression to generate out-of-sample forecasts for a given time horizon.

[20] The fact that the combined approach outperforms the VAR model could be explained by the following arguments. The choice of the variables included in the SVAR were not made on the basis of their ability to forecast real BCNE prices but on their ability to give information pertinent to the identification of the permanent and transitory components of real BCNE prices. Second, SVAR literature shows that it is important to include a large number of lags in the SVAR in order to identify properly the transitory component of a variable. With a small sample, this strategy is clearly not optimal in terms of out-of-sample forecast performance because it relies on many estimated parameters. The combined model attempts to address those issues.

TABLE 3: OUT-OF-SAMPLE FORECASTS OF REAL BCNE PRICES (FORECASTING PERIOD: 1992Q1 - 2001Q4)

	VAR	*AR(1) model*	*Combined approach*
Forecasting horizon	*RMSE*	*RMSE*	*RMSE*
2	0.0633	0.0649	0.0530
3	0.0737	0.0671	0.0564
4	0.0815	0.0748	0.0640

Tables 4 to 6 report the *p*-values of the forecast encompassing test statistic, which was originally devised by Chong and Hendry (1986) to compare two competing models based on the out- of-sample forecasting errors. We compare our approach with two benchmark cases: AR(1) model and RW model.[21] The encompassing test results support the use of the combined approach. The results indicate that we can not reject the null hypothesis of "A encompasses B", which implies that forecasts from either of two benchmark models (model B) are unlikely to improve the forecasting performance of the combined approach (model A) for any forecasting horizon. On the other hand, we can always reject the null hypothesis of "B encompasses A" at the 5% significance level except for one case when we compare four-quarter ahead forecasts with the RW model. This implies that our combined approach improves the forecasting performance of two benchmark models.

TABLE 4: FORECAST ENCOMPASSING TESTS, 2-STEP AHEAD FORECASTS (*p-value*)

Encompassing tests for A = combined approach and B = benchmark models		
Null hypothesis	*Combined approach vs AR(1) model*	*Combined approach vs RW model*
A encompasses B	0.288	0.994
B encompasses A	0.009	0.015

4. THE REAL CRUDE OIL PRICE MODEL

This section consists of two subsections. The first part describes the methodology that we use to identify and forecast the transi-

[21] Given that the results in table 3 have shown that the AR(1) model performs better than the VAR model, we do not need to report here the encompassing test results between our approach and VAR model.

tory and permanent components of real WTI crude oil prices. The second subsection presents main results of the model.

TABLE 5: FORECAST ENCOMPASSING TESTS, 3-STEP AHEAD FORECASTS (*p-value*)

Encompassing tests for A = combined approach and B = benchmark models		
Null hypothesis	*Combined approach vs AR(1) model*	*Combined approach vs RW model*
A encompasses B	0.685	0.597
B encompasses A	0.008	0.039

TABLE 6: FORECAST ENCOMPASSING TESTS, 4-STEP AHEAD FORECASTS (*p-value*)

Encompassing tests for A = combined approach and B = benchmark models		
Null hypothesis	*Combined approach vs AR(1) model*	*Combined approach vs RW model*
A encompasses B	0.727	0.781
B encompasses A	0.025	0.061

4.1 Methodologies

Figure 7 plots the series of real crude oil prices over the sample period. As seen, crude oil prices have experienced a few large permanent shifts over history, most notably in 1979-80 and 1985-1986.[22] To test for structural breaks in the data (under the assumption that the time and the number of breaks are unknown), we use the methodology proposed by Bai and Perron (1998) (hereafter BP).[23] The strength of the BP methodology is that we can estimate the time and the number of structural breaks endogenously with allowance for varying parameters across regimes. Given the fact that world oil prices are spot prices, we also allow the model to capture the contemporaneous effect of the world output gap on prices.

4.2 Results of the real crude oil price model

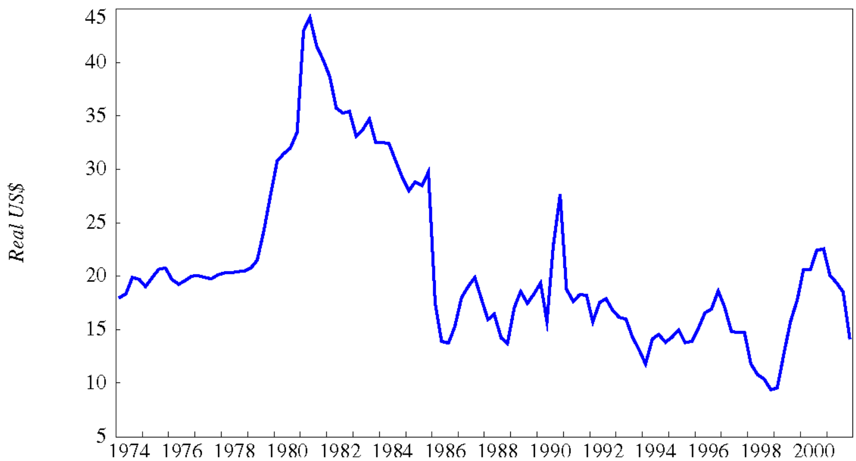
This section is divided into three parts. First, we examine the

[22] These large movements in price are related to specific developments in the market, particularly with changes in the behaviour of the OPEC cartel.

[23] See Appendix C for a brief discussion of the BP methodology.

estimation results for the model. Second, we use the estimated model to identify transitory and permanent components of crude oil prices. Third, we evaluate the model's out-of-sample forecasting performance.

FIGURE 7: REAL CRUDE OIL PRICES, 1974-2000



4.2.1 Estimation of real WTI crude oil model

We first consider real crude oil prices over the whole sample period. An ADF test cannot reject the presence of a unit root. However, when allowing for structural changes, we can reject the hypothesis that the series has a unit root.[24]

Using the procedure proposed by BP, we estimate a single equation model as equation (2) with allowance for up to three structural changes. The sample period is from 1974q2 to 2001q4. In this framework, all model parameters are allowed to shift at the structural break point. At the 5% level, the test detects two breaks in 1979q3 and 1985q4.[25] It is interesting to note that these two break points well match two large oil price shocks over the history. The first oil price shock in the late 1970s began with the Iranian revolution and the accompanying disruption of its petroleum exports. Moreover, the outbreak of the war between Iran

[24] The sum of AR coefficients is 0.58 and the t-statistic is -5.3, which compares to a 2.5% critical value of –5.3 (see Zivot et al. (1991)).

[25] The test statistics for the sup$F(1|0)$ and sup$F(2|1)$ are 30.7 and 27.4, respectively. This compares to the 5% critical values of 20.1 and 22.1.

and Iraq in 1980 shook the oil market as well. The second oil shock in the mid-1980s was primarily caused by the collapse of OPEC discipline. From 1982 to 1985, OPEC attempted to set production quotas low enough to stabilize prices. These attempts met with repeated failure as various members of OPEC would produce beyond their quotas. During most of this period, Saudi Arabia acted as the swing producer cutting its production to stem the free falling prices. In late 1985, Saudi Arabia stopped doing that and increased its production, and this eventually caused the oil price plunge in 1986. The first experiment we do is to use three dummy variables to capture the three BP regimes separated by two breaks.

Regression equation:

$$Rwti_t = D1{*}\,Dumm1 + D2{*}Dumm2 + D3{*}Dumm3 + \\ + C(L)Rwti_{t-1} + DD(L)Wygap_t + E(L)Ergap_t \qquad (2)$$

TABLE 7: THE REAL CRUDE OIL PRICE MODEL, FULL SAMPLE: 1974Q2-2001Q4 (OLS ESTIMATION OF EQUATION (2) WITH THREE DUMMY VARIABLES)

Variable	*Model 1 (with real U.S. dollar effective exchange rate gap)*	*Model 2 (without real U.S. dollar effective exchange rate gap)*
Dumm1	1.013 (6.00)	0.976 (5.81)
Dumm2	1.188 (6.04)	1.144 (5.85)
Dumm3	0.927 (5.94)	0.898 (5.76)
Rwtit-1	0.912 (9.57)	0.935 (9.88)
Rwtit-2	-0.245 (-2.77)	-0.255 (-2.89)
Wygapt-1	0.017 (1.69)	0.023 (2.42)
Ergapt-2	-0.005 (-1.48)	-
$\bar{R}^2$	0.85	0.83

NOTE: *t*-statistics are in parentheses.

Table 7 reports the OLS estimation results of equation (2) without allowing for varying coefficients across regimes. The purpose of this experiment is to examine the impact of world economic conditions on crude oil prices over the full sample. We report the model parameter estimates for the two cases with and without the real U.S. dollar exchange rate gap. As seen in Table 7, the results for both cases are almost identical. Real crude oil prices are fairly persistent over the full sample with an AR root of 0.67 (the sum of the two autoregressive coefficients) and the estimated coefficient associated with the lagged world output gap is about 0.02. The real U.S. dollar effective exchange rate gap is not statistically significant over the whole sample period.

Tables 8 to 10 report the BP procedure results for three regimes with allowance for varying coefficients. As seen, the estimate of the lagged world output gap changes considerably across regimes. Although it is not statistically significant in the first regime, the estimate has the correct sign. In the second regime, it exhibits the wrong sign, but it is statistically insignificant. In contrast, its magnitude increases substantially in the most recent regime with a value of about 6%, which is almost three times the average value over the entire sample as reported in Table 7.[26]

Furthermore, since WTI crude oil prices are spot prices, we would expect crude oil prices to respond immediately to the world

TABLE 8: THE REAL CRUDE OIL PRICE MODEL (REGIME 1: 1974Q2 - 1979Q3)

Variable	*Coefficient*
Dumm1	1.986 (17.15)
Rwtit-1	0.170 (1.97)
Rwtit-2	0.170 (2.10)
Wygapt-1	0.007 (1.41)
Ergapt-2	-0.186 (-1.23)
$\bar{R}^2$	0.92

NOTE*:* *t*-statistics are in parentheses.

[26] However, the real U.S. dollar effective exchange rate gap is not statistically significant, and excluding it increases the magnitude and improves the significance of the estimated elasticity of real crude oil price with respect to the world output gap. The result of an alternative model excluding the real U.S. exchange rate gap is reported in table 11.

TABLE 9: THE REAL CRUDE OIL PRICE MODEL (REGIME 2: 1979Q4 - 1985Q4)

Variable	*Coefficient*
Dumm2	1.048 (5.50)
Rwtit-1	0.868 (7.34)
Rwtit-2	-0.165 (-1.61)
Wygapt-1	-0.007 (-0.74)
Ergapt-2	-0.778 (-1.89)
	0.92

NOTE: *t*-statistics are in parentheses.

output gap shock. The Hausman test indicates that the null hypothesis of exogeneity of the world output gap is rejected at 5% level of significance. This implies that applying the OLS to the BP procedure cannot produce consistent estimates, and we should instead use the IVE. The instruments used are four lags of all explanatory variables in the model. However, given the small number of observations in the first two regimes, applying IVE to the BP

TABLE 10: THE REAL CRUDE OIL PRICE MODEL (REGIME 3: 1986Q1 - 2001Q4)

Variable	*Coefficient*
Dumm2	1.103 (4.01)
Rwtit-1	0.873 (8.96)
Rwtit-2	-0.261 (-3.74)
Wygapt-1	0.055 (2.64)
Ergapt-2	-0.390 (0.61)
$\bar{R}^2$	0.92

NOTE: *t*-statistics are in parentheses.

procedure tends to give us very biased results.[27] Hence, we use IVE to re-estimate equation (2) only for the third regime, which has a relatively sufficient number of observations.[28] Table 11 reports the IVE estimates of all parameters. The world output gap remains statistically significant at the 5% level of significance, and the estimated coefficient associated with the world output gap is around 4.5%. In our final model, we also add the change in crude oil inventories, another key indicator of crude oil prices. As shown in the table, the third lag of the change in crude oil inventories is statistically significant and exhibits the expected sign.[29]

TABLE 11: THE REAL CRUDE OIL PRICE MODEL, REGIME 3: 1986Q1 - 2001Q4 (IVE ESTIMATION OF REGIME 3 WITH CONTEMPORANEOUS WORLD OUTPUT GAP AND OIL INVENTORIES)

Variable	*Coefficient*
Dumm3	1.071 (3.98)
Rwtit-1	0.965 (11.31)
Rwtit-2	-0.344 (-3.26)
Wygapt	0.045 (3.07)
Inventoryt-3	-0.035 (-4.34)

NOTE: *t*-statistics are in parentheses.

4.2.2 The transitory and permanent components of real crude oil prices

Given the nature of the model, the permanent component consists of three different means caused by two structural breaks. Figure 8 plots the world output gap vis-a-vis the transitory com-

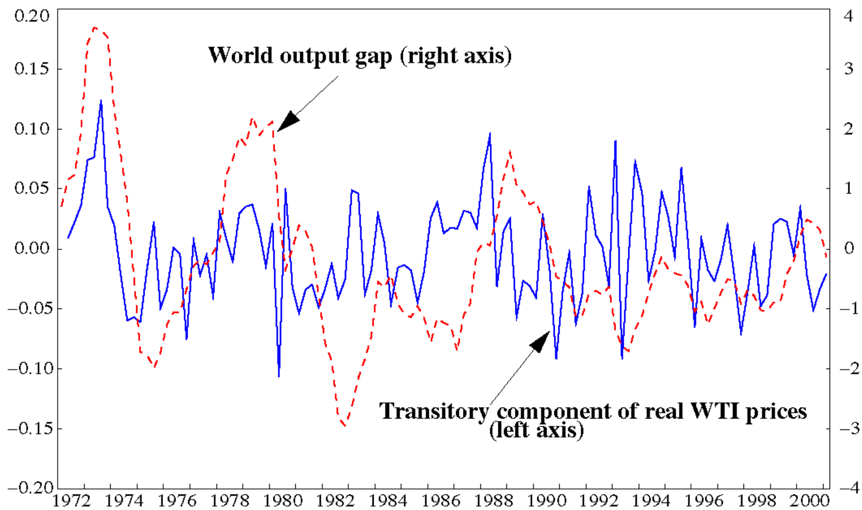
[27] The bias problem becomes severe in a small sample. See Davidson and Mackinnon (1993) for details.

[28] As Tables 8 to 10 have shown that the strongest link between the world output gap and real crude oil prices is in the third regime, we are more interested in estimating model parameters for this regime. Also, it is the current regime and therefore it has particular relevance for current forecasts.

[29] Because only the third lag of the change of crude oil inventory enters our model, we do not need a model to forecast this variable in order to forecast real oil prices over very short time horizons.

ponent of real WTI crude oil prices across the three regimes. The closest link between the two variables appears in the most recent regime. Figure 9 shows that a 1% positive shock to the world output gap leads to an approximate 12% peak response of real WTI crude oil prices with a lag of two to three quarters.

FIGURE 8: THE WORLD OUTPUT GAP AND THE TRANSITORY COMPONENT OF REAL WTI CRUDE OIL PRICES, 1972-2000 (in per cent)



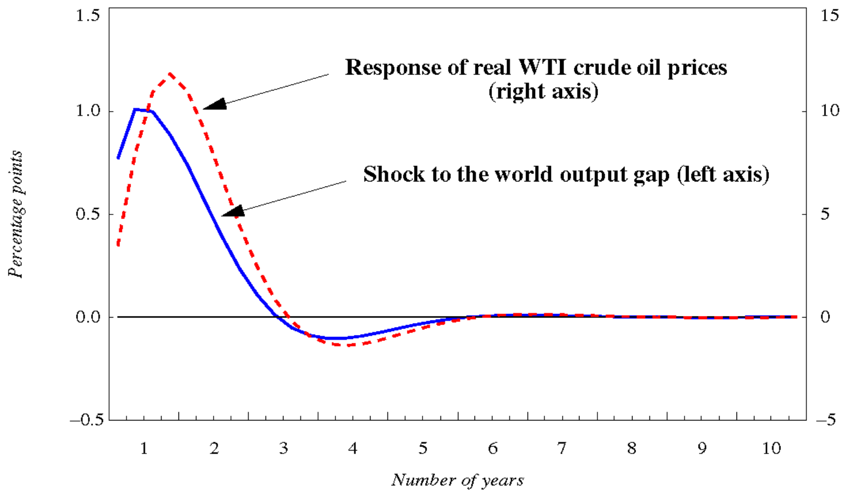
4.2.3 Out-of-sample forecast of the real crude oil price model

We use the estimated single equation model from Table 11 to forecast real crude oil prices. We compare the two- to four-step ahead forecasting performance of our model with two benchmark models from 1992q1 to 2001q4. Table 12 compares the RMSE of two to four-step ahead out-of-sample forecasts of our model with two benchmark models: AR(1) model and RW model. It is evident that regardless of the forecasting horizons concerned, our

TABLE 12: THE REAL CRUDE OIL PRICE MODEL: RMSE OF OUT-OF-SAMPLE FORECASTS (1992Q1 - 2001Q4)

Forecasting horizon (quarters)	*Our model*	*RW model*	*AR(1) model*
2	0.134	0.176	0.208
3	0.170	0.225	0.239
4	0.202	0.276	0.246

FIGURE 9: SIMULATION OF A ONE PER CENT POSITIVE SHOCK TO THE WORLD OUTPUT GAP (REGIME 3)



model uniformly outperforms the other two as reflected by smaller values of RMSE.

We have also estimated an alternative specification excluding the oil inventory measure. We find models excluding the inventory measure always perform worse than those including it. Furthermore, for near-term forecasting (two-quarter ahead), they are even worse than naive forecasts using a RW model.

TABLE 13: FORECAST ENCOMPASSING TESTS, 2-STEP AHEAD FORECASTS (*p-value*)

Encompassing tests for A = our model and B = benchmark models		
Null hypothesis	*Our model vs AR(1) model*	*Our model vs RW model*
A encompasses B	0.157	0.499
B encompasses A	0.016	0.033

TABLE 14: FORECAST ENCOMPASSING TESTS, 3-STEP AHEAD FORECASTS (*p-value*)

Encompassing tests for A = our model and B = benchmark models		
Null hypothesis	*Our model vs AR(1) model*	*Our model vs RW model*
A encompasses B	0.356	0.654
B encompasses A	0.041	0.086

Tables 13 to 15 report the *p*-values of forecast encompassing test statistics. The test results again support the use of our model. The results show that we cannot reject the null hypothesis of "A encompasses B", which implies that it is impossible to improve the forecasting capability of our model (model A) with the help of forecasts from either of two benchmark models (model B) for any forecast horizon. On the other hand, we can always reject the null hypothesis of "B encompasses A" at the 10% significance level, which implies that our model can provide useful information to improve the forecasting performance of two benchmark models.

TABLE 15: FORECAST ENCOMPASSING TESTS, 4-STEP AHEAD FORECASTS (*p-value*)

Encompassing tests for A = our model and B = the benchmark models		
Null hypothesis	*Our model vs AR(1) model*	*Our model vs RW model*
A encompasses B	0.660	0.411
B encompasses A	0.046	0.077

5. CONCLUSION

To summarize, the variance decomposition shows that about 60% of the total variation in real BCNE prices is attributed to demand shocks. This is consistent with other studies on non-energy commodity prices in the literature. We also find a very close link between world economic conditions and transitory movements in real BCNE prices. For real WTI crude oil prices, a multiple structural-break test identifies two structural breaks over the sample period. We use the exogenous mean shifts of real WTI crude oil prices across three different regimes as a measure of the permanent component of real crude oil prices. The corresponding forecasting model shows the strongest link between the cyclical component of real WTI crude oil prices and world economic conditions occurs in the most recent history.

In terms of the forecasting performance, we compare two- to four-step ahead forecasts of our models with benchmark models: a VAR model, an AR(1) model and a RW model. These tests show that our models uniformly outperform the baseline models,

All results in this paper suggest that we can provide better short-term forecasts of world commodity prices relative to benchmark models. However, there are several extensions to this paper that may be worth pursuing. A potential avenue for future work is to put more effort into developing the supply side of the model, and to explore key supply indicators such as productivity growth in commodity producing sectors that can reasonably explain the long-term behavior of world commodity prices. The models with richer structures can be further developed in order to better analyze and forecast the commodity prices both in the short-run and long-run.

Appendix A:

WEIGHTS OF COMMODITIES AND SUB-INDEXES IN THE BCPI

Item	*Weight*
Total BCPI	100.0
1.0 Energy	34.9
Crude Oil (WTI)	21.7
Natural Gas	10.4
Coal	2.7
2.0 Total BCPI excluding Energy (BCNE)	65.1
2.1 Food	18.8
2.1.1 Grains and Oilseeds	8.8
Barley	1.2
Canola	1.3
Corn	0.8
Wheat	5.5
2.1.2 Livestock	9.2
Cattle	6.1
Hogs	3.2
2.1.3 Fish	0.7
Cod	0.04
Lobster	0.34
Salmon	0.36
2.2 Industrial Materials	46.3
2.2.1 Metals	14.4
Gold	2.8
Silver	0.6
Aluminum	3.0
Copper	2.9
Nickel	2.4
Zinc	2.7
2.2.2 Minerals	2.3
Potash	1.3
Sulphur	1.0
2.2.3 Forest Products	29.6
Lumber	9.0
Newsprint	8.3
Pulp	12.3

Appendix B:

THE BLANCHARD-QUAH (1989) DECOMPOSITION AND THE LINK BETWEEN THE STRUCTURAL FORM AND THE REDUCED FORM OF THE MODEL

The shocks and the variables in the SVAR for real BCNE prices can be defined as follows:

$$\varepsilon_t = \begin{bmatrix} \varepsilon_s \\ \varepsilon_{d1} \\ \varepsilon_{d2} \\ \varepsilon_{d3} \\ \varepsilon_{d4} \end{bmatrix} \text{ and } Z_t = \begin{bmatrix} \Delta Rbcne \\ Wgap \\ \Delta W\pi \\ RRus \\ \Delta Erus \end{bmatrix}, \tag{1}$$

where is the only type of shock that will have a permanent effect on real BCNE prices and the other four shocks are restricted to have only transitory effects on real BCNE prices. Given that we are only interested in the decomposition of real BCNE prices into a permanent component and a *total* transitory component, we treat four transitory shocks as a single aggregate demand shock term.

The moving average representation of the structural model is defined as follow:

$$Z_t = \Gamma(0)\varepsilon_t + \Gamma_1\varepsilon_{t-1} + \Gamma_2\varepsilon_{t-2} + \ldots = \Gamma(L)\varepsilon_t, \tag{2}$$

and the corresponding long-run effect matrix of the structural shocks is:

$$\Gamma(1) = \Gamma(0) + \Gamma_1 + \Gamma_2 + \ldots + \Gamma_\infty , \tag{3}$$

where, $E(\varepsilon_t \varepsilon_t) = I$. The diagonal elements are normalized to 1's only for the purpose of simplification.

In order to identify the structural model, we first estimate the reduced form of the model (i.e. VAR):

$$Z_t = \sum_{i=1}^{p} \Pi_i Z_{t-i} + e_t, \tag{4}$$

where p is the number of lags[30] and e_t is the vector of the reduced form shocks, where $E(e_t e_t)\Sigma$.

Given that the stochastic process is stationary, the moving average representation of equation (4) is defined by the following relationship:

$$Z_t = e_t + C_1 e_{t-1} + C_2 e_{t-2} + \ldots = C(L) e_t , \tag{5}$$

and the long-run effect matrix of the reduced-form shocks is:

$$C(1) = 1 + C_1 + C_2 + \ldots + C\propto . \tag{6}$$

Given equations (2) and (5), the reduced-form residuals are linked to the structural residuals in the following way:

$$e_t = \Gamma(0)\varepsilon_t . \tag{7}$$

Consequently,

$$E(e_t e_t) = \Gamma(0)\Gamma(0)' \quad \text{because } E(\varepsilon_t \varepsilon_t) = I. \tag{8}$$

In addition, the long-run effect matrix of the reduced-form shocks, $C(1)$, is linked to the equivalent matrix of the structural shocks ($\Gamma(1)$) and,

$$\Gamma(1) = C(1)\Gamma(0). \tag{9}$$

In order to identify the structural model, we need to impose a sufficient number of restrictions on the system of equations formed by equations (8) and (9). The fifty elements of the structural form matrices Γ(0) and Γ(1) are unknown and the elements of C(1) and $E(e_t e_t)$ are known from the estimation of the reduced form model. Given that Σ is a symmetric matrix, equations (8) and (9) contain forty different relations. Therefore, we have to impose ten restrictions on the elements of Γ(0) and Γ(1). The Blanchard and Quah decomposition consists of imposing restrictions on the long-run effect matrix of the structural shocks. (i.e. Γ(1)) instead of imposing a predetermined structure on the variables by the restrictions on the Γ(0) matrix.) We achieve this by imposing that Γ(1) is triangular. Given these restrictions, the system of equations formed by equations (8) and (9) is solvable, and therefore the structural model is identified. The following equation shows the restrictions imposed on the long-run effect matrix of the structural shocks with, for presentation purposes, the

[30] The reduced-form model includes eight lags. We have estimated a model which includes six lags, and results are almost identical.

shocks of the structural model on the horizontal axis and the variables of the model, in levels, on the vertical axis:

$$\begin{array}{c} \\ \begin{bmatrix} Rbcne_? \\ \int Wygap_? \\ W\pi_? \\ \int RRus_? \\ Erus_? \end{bmatrix} \end{array} \begin{array}{c} \begin{bmatrix} \varepsilon_{s_?} & \varepsilon_{d1} & \varepsilon_{d2} & \varepsilon_{d3} & \varepsilon_{d4} \end{bmatrix} \\ \begin{bmatrix} r_{11} & 0 & 0 & 0 & 0 \\ r_{21} & r_{22} & 0 & 0 & 0 \\ r_{31} & r_{32} & r_{33} & 0 & 0 \\ r_{41} & r_{42} & r_{43} & r_{44} & 0 \\ r_{51} & r_{52} & r_{53} & r_{54} & r_{55} \end{bmatrix} \end{array} = \Gamma(1) \qquad (10)$$

Therefore, we impose that is the only type of shock that has a permanent effect on real BCNE prices. This gives four restrictions. The other six restrictions are just required to decompose the total transitory component into its four subcomponents. Consequently, they are irrelevant for decomposing real BCNE prices into a permanent and a *total* transitory component. In other words, results concerning the decomposition of real BCNE prices are unaffected by the assumption regarding the ordering of the four last variables. This simply reflects the fact that, in the long-run, the model is recursive from top to bottom.

Appendix C:

TECHNICAL DETAILS ON BAI AND PERRON (1995) METHODOLOGY

We consider a multiple linear regression with m breaks ($m+1$ regimes). The equation of real crude oil prices is specified in a compact matrix notation as:

$$Y = X\beta + \varepsilon$$

where Y is the observed dependent variable at time t, X is the matrix of covariates, which is partitioned according to the break points T_B, β is the corresponding vector of coefficients, and ε is the disturbance term. The break points $(T_1,\ldots,T_m)$ are explicitly treated as unknown. The purpose is therefore to estimate the regression coefficients and the break points simultaneously when T observations of Y and X are available. The estimation method considered is based on the least squares principle. For each m-partition $(T1,...,Tm)$, the associated least squares estimates of β is obtained by minimizing the sum of squared residuals, denoted here as ST. The estimated coefficients and break points are such that

$$(T_1, \ldots, T_m) = \operatorname{arg}min_{\{T_1, \ldots, T_m\}} S_T(T_1, \ldots, T_m)$$

where the minimization is taken over all partitions $(T_1,...,T_m)$, so the break-point estimators are global minimizers of the objective function.

BP proposed a test based on the supremum of the F-statistic, which is called the sup$F$ test, to detect the multiple breaks. This test is labelled as the sup$F(l+1\,|\,l)$. The method amounts to the application of $l+1$ tests of the null hypothesis of no structural change versus the alternative hypothesis of l changes. The test is applied to each segment containing the observation T_{m-1} to T_m with $m = 1,...,l+1$. We reject the null hypothesis in favour of a model with $l+1$ breaks if the overall minimal value of squared residuals is sufficiently smaller than the sum of squared residuals from the l-break model. The break date is selected as the one associated with this overall minimum. The asymptotic distribution of the test statistic depends on the selected minimal length of the segments which is a function of a trimming parameter.[31] To apply the test, we use a trimming of fifteen per cent. Hence, given our sample period of 1974-2001, no more than six breaks are allowed while each regime must have at least sixteen observations.

REFERENCES

Andrews, D. W. K. 1993. "Tests for Parameter Instability and Structural Changes with Unknown Change Points." *Econometrica* 61: 821-25.

Bai, J. and P. Perron. 1998. "Computation and Analysis of Multiple Structural-Change Models." *Econometrica* 66: 47 - 78.

[31] We need to trim the sample by some fraction since the test statistic diverges to infinity (see Andrews (1993a) for details). For this reason, we cannot test for the presence of a structural break in the first/last four years of the sample or the one very close to another break point.

Beveridge, S. and C. Nelson. 1981. "A New Approach to Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to the Measurement of the Business Cycle." *Journal of Monetary Economics* 7: 151-74.

Blanchard, O. J. and D. Quah. 1989. "The Dynamic Effect of Aggregate Demand and Supply Disturbances." *American Economic Review* 79: 655-73.

Borensztein, E. and M. Reinhart. 1994. "The Macroeconomic Determinants of Commodity Prices." Staff Paper No. 41: 236 - 61, International Monetary Fund.

Cashin, P. and C. McDermott. 2001. "The Long-Run Behavior of Commodity Prices: Small Trends and Big Variability." Staff Paper No. WP/01/68, International Monetary Fund.

Chong, Y. and D. Hendry. 1986. "Econometric Evaluation of Linear Macroeconomic Models." *Review of Economic Studies* 53: 671-90.

Coletti, D. 1992. "The long-run behaviour of key Canadian non-energy commodity prices (1900 to 1991)." *Bank of Canada Review (*Winter 1992-1993): 47-56.

Cuddington, J. and L. Hong. 2000. "Will the Emergence of the Euro Affect World Commodity Prices?" Staff Papers No. WP/00/208: International Monetary Fund.

Davidson, R. and J. MacKinnon. 1993. *Estimation and Inference on Econometrics*, New York: Oxford University Press.

Harvey, A. C. 1985. "Trends and Cycles in Macroeconomic Time Series." *Journal of Business & Economic Statistics* 3: 216-27.

Harvey, A. C. 1989. *Forecasting, Structural Time Series Models and the Kalman filter*, New York: Cambridge University Press.

Hunt, B. 1995. "The effect of foreign demand shocks on the Canadian economy: An analysis using QPM." *Bank of Canada Review*: Autumn 1992.

Lalonde, R. 1998. "Le PIB potentiel des États-Unis et ses déterminants: la productivité de la main-d'oeuvre et le taux d'activité." Working Paper No. 13. Ottawa: Bank of Canada.

Lalonde, R. 2000. "Le modèle U.S.M d'analyse et de projection de l'économie américaine." Working Paper No. 19. Ottawa: Bank of Canada.

Newey, Whitney L. 1985. "Generalized Method of Moments Specification Testing." *Journal of Econometrics* 29: 229-56.

Quah, D. 1992. "The Relative Importance of Permanent and Transitory Components: Identification and Some Theoretical Bounds." *Econometrica* 60: 107-18.

Reinhart, C. M. 1991. "Fiscal Policy, the Real Exchange Rate, and Commodity Prices." Staff Paper 38: 506-24: International Monetary Fund.

Reinhart, C. M. and P. Wickham. 1994. "Commodity Prices: Cyclical Weakness or Secular Decline?" Staff Papers 41: 175 - 213: International Monetary Fund.

St-Amant, P. and S. van Norden. 1997. "Measurement of the Output Gap: A Discussion of Recent Research at the Bank of Canada." Technical Report No. 79. Ottawa: Bank of Canada.

Zivot, E. and D. W. K. Andrews. 1992. "Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis." *Journal of Business And Statistics* 10: 251-70.

Prudence Serju

Monetary conditions and core inflation: an application of neural networks

1. INTRODUCTION

Previous approaches to forecasting inflation by the Bank of Jamaica (BOJ) employed a Vector Error Correction Model (VECM) (Robinson 1997) and a model that decomposes inflation into its local and foreign cost components. These models have been complemented by conjectural analysis and rely on a diverse set of indicators. The BOJ models are used for the short-term forecasting of headline inflation.

However, it is generally accepted that monetary policy should focus on underlying or core inflation. Roger (1995) suggests that shocks to the general price level, which are perceived as one-time events, should not have a lasting effect on the inflation rate, and as such, it would be inappropriate for the policy makers, who are

Paper prepared by P. Serju, Research Services Department, Research and Economic Programming Division, Bank of Jamaica. The paper was presented at the VII Meeting of the Network of America Central Banks Researchers, organized by the Banco de Guatemala, in Guatemala City, November 20-22, 2001. The views expressed are those of the author and do not necessarily reflect those of the Bank of Jamaica. (The figures in the Appendix have been omitted for editorial reasons, readers interested in the figures may contact the author: Prudences@boj.org.jm).

targeting the inflation rate, to respond. Similarly, Motley (1997) argues that temporary price shocks are often due to supply shocks, such as unusual weather, which affect harvests. These supply shocks while affecting the level of prices do not necessarily affect its long-run growth rate. Thus, if the goal of policy makers is to control inflation, core inflation should be selected as the policy target, in preference to headline, as the former inherently avoids shocks or disturbances that add noise to measured inflation.

Accordingly, this paper develops and assesses the forecast performance of various models for core inflation. The remaining sections of the paper are organized as follows. Section 2 gives a brief discussion on the history of core inflation in Jamaica. Section 3 describes methodological issues relating to the neural network model. Section 4 presents the results and evaluates the competing model forecasts. The conclusion is presented in the final section.

2. INFLATION IN JAMAICA

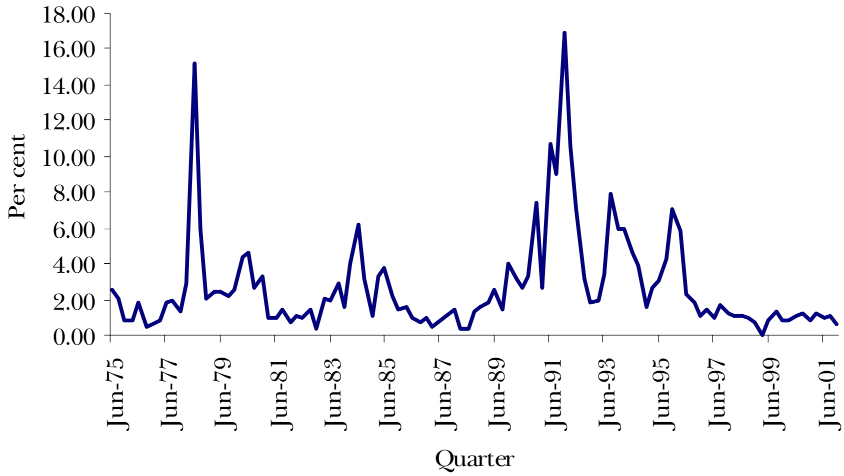
The quarterly trend in core inflation[1] in Jamaica between 1975 and 2001 is depicted in figure 1. The average quarterly core inflation over the period was 2.7 per cent. With the exception of 1978 and 1984, core inflation was fairly stable over the period 1975 and 1990. The 1978 hyper inflationary episode resulted from the devaluation of the exchange rate due to the abolition of the dual exchange rate regime and the implementation of a crawling peg regime. Similarly, the rise in inflation in 1984 coincided with a change in the exchange rate regime and the introduction of the foreign exchange auction mechanism in March of that year, which saw the devaluation of the exchange rate. Core inflation peaked in 1991, following the liberalization of the foreign exchange rate regime. It is important to note that the incidence of inflation over the review period has been associated with strong money growth, which has served to undermine the stability of the exchange rate regime. The average quarterly expansion of the monetary base between 1975 and 2001 was 5.8 per cent, with particularly strong growth being recorded between 1982 and 1984.

Since 1997, core inflation has been fairly stable, reflecting the Central Bank's focus on containing underlying inflation. This

[1] The measure of core used in this paper is the trimmed mean. See Allen (1997).

trend in core inflation has occurred alongside major changes in the Jamaican economy, including substantial reductions in tariffs, partial elimination of price controls, subsidies and quantitative restrictions on commodity trade.

FIGURE 1. CORE INFLATION IN JAMAICA, 1975-2001



There are many different theories as to the causes of inflation, but there is no universally accepted theory that explains inflation in all countries. For our purposes, our variable selection process will be guided by previous empirical work done on inflation in Jamaica, which was essentially based on the monetarist views.

Bourne and Persaud, in 1977, conducted one of the first econometric investigations of inflationary sources in the Jamaican economy. The study showed that a devaluation of the local currency and increased foreign prices were the main causes of inflation in Jamaica in the 1960 and 1970. Downes (1992) examined a structural/monetarist model of inflation in Jamaica, among other countries using an error correction model. He found that monetary policy, the exchange rate vis-à-vis the United States of America (USA) dollar and USA price inflation were significant in explaining inflation in Jamaica, with the monetary policy variable being the most important. Using monthly data for the period 1978:1 to 1990:6, Thomas (1994) employed a monetarist model to evaluate the impact of changes in the exchange rate, net domestic credit, foreign assets and international prices on Jamaica's inflationary process. He found the Treasury Bill rate, exchange rate, foreign prices and net foreign assets to be the important variables in elucidating long-run price behaviour in Jamaica. However, he found net domestic credit to be insignificant in explaining inflation in Jamaica. Robinson (1995) tested for cost-push against demand-pull inflation using a model developed by Harberger in 1963. Over the sample period, 1986 to 1994, he found a bi-directional causation between prices and wages. He concluded that, " Excessive cash holdings are translated first into consumer demand, then into higher prices and then into higher wages and then into higher prices". This suggests that wages do not necessarily initiate inflation in the Jamaican economy. The dominant result he found was that current money supply changes were most significant.

The literature reviewed above suggests that monetary policy changes, exchange rate movements and foreign inflation have been the more dominant factors in explaining inflation in Jamaica. Underlying inflation is generally determined by demand pressures primarily associated with output gaps and monetary impulses. However given the absence of consistent intra-year data on aggregate supply and demand and consistent with the previous studies on inflation in Jamaica, the explanatory variables used in this study are base money to capture monetary policy changes, the exchange rate and foreign prices. The impact of imported inflation is captured mainly by oil prices as it was found that domestic inflation reacted more significantly to oil prices relative to general foreign consumer prices. In addition the 30- day treasury-bill rate is included to capture shifts in monetary policy.[2]

3. NEURAL NETWORK MODEL

Over the last two decades there has been increased research on the forecasting performance of various time series models. (Makridakis et al. (1982, 1986)). It has been found that no single method dominate the forecasting landscape. However, it has been established that simple and parsimonious models are robust under a wide range of conditions.

Recently, Artificial Neural Network models (ANN) have emerged as an alternate forecasting tool. Remus & O'Connor (1998) indicate that ANN model excel in pattern recognition and forecasting from pattern clusters. The ANN models have two ad-

[2] See Robinson (1998).

vantages when compared with other traditional methods of forecasting. Firstly, they are universal approximators of functions in that they can approximate whatever functional form best characterizes the time series. In this context, they are inherently nonlinear, but can overcome the limitations of linear forecasting models. Secondly, ANN models have been proven to be better than traditional forecasting methods for *long term* forecast horizons, but are often as good as traditional forecasting methods over shorter forecast horizons.

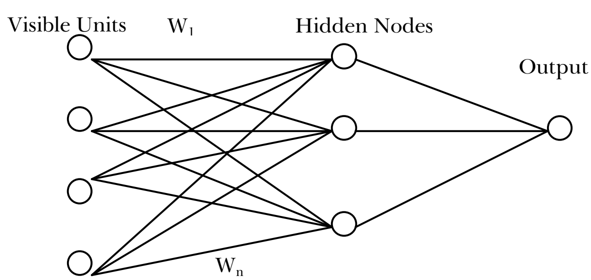
McCulloch's (1943) paper laid the theoretical basis for the development of ANN models. Minsky (1954) developed the computing platform on which current ANN models are processed. Analysis using ANN was further enhanced with the development of the back-propagation technique (Rumelhart & McClelland, 1986).

ANN models are inspired from biological neural networks. They are developed using software that attempt to mimic the human brain's ability to classify patterns or to make predictions or decisions based on past experience (Gately, (1996)). While the human brain relies on inputs from the five senses, ANN uses inputs from data sets.

3.1 Architecture of ANN models

ANN models typically contain three or more layers consisting of input (visible), hidden and output units.

FIGURE 2. TYPICAL NEURAL NETWORK ARCHITECTURE



Input layers receive input patterns directly, while hidden layers neither receive inputs directly nor are given direct feedback. Hidden units are the stock of units from which new features and new internal representations can be created. In a neural network model, the user may specify a pattern of inputs to the visible units but not the external inputs to the hidden unit. The hidden unit net input is based only on the outputs from other units to which they are connected. In this paper, the exchange rate, base money, three seasonal factors, a pulse dummy and an autoregressive term were specified in the input layer. The exchange rate and base money were differenced before they entered the model.

3.2 Interactive activation

The units in a neural network take on continuous activation values between a maximum and minimum value, though their output – the signal they transmit to other units – is not necessarily identical to their activation. In the ANN model, output = $[a_j]^+$. Here, a_j refers to the activation of unit j, and the expression $[a_j]^+$ has value $a_j > 0$; and zero otherwise. The index j ranges over all units with connections to unit 'i'.

Units change their activation based on the current activation of the unit and the net input to the unit from other units, or from outside the network. The net input to a particular unit (say unit i) is the weighted sum of all the output from other units plus any external input:

$$net_i = \sum_j w_{ij} output_j + extinput_i \tag{1}$$

In general the weights (w_{ij}) can be positive or negative.

Once the net input into a unit has been computed, the resulting change in the activation of the unit is as follows:

$$\Delta a_i = \begin{cases} (\max = a_i) net_i - decay(a_i - rest) & \text{if } (net_i > 0) \\ (a_i - \min) net_i - decay(a_i - rest) & \text{otherwise} \end{cases} \tag{2}$$

where max, min, rest and decay are all parameters. In particular, we choose max = 1, min ≤ rest ≤ 0, and decay between 0 and 1. Note also that a_i is assumed to start, and to stay within the interval [max, min].

The optimal activation of the unit occurs when the incremental activation of the unit is zero. Setting $\Delta a_i = 0$ and rearranging expression 2 results in the following equilibrium condition for the activation of the unit:

$$a_i = \frac{(\max)(net_i) + (rest)(decay)}{net_i + decay} \tag{3}$$

Using max = 1 and rest = 0, this simplifies to:

$$a_i = \frac{(net_i)}{net_i + decay} \tag{4}$$

Equation 4 indicates that the equilibrium activation of a unit will always increase as the net input increases, however, it can never exceed 1 (or, in the general case, max). The decay term acts as a kind of restoring force [as an equilibrating force] that tends to bring the activation of the unit back to zero (or to rest in the general case). Decreasing the value of this decay parameter increases the equilibrium activation of the unit.

3.3 Learning

Neural network models are of interest because they learn, naturally and incrementally, in the course of processing. One classical procedure for learning (i.e. understanding the data generation process) is the error correcting or delta learning rule as studied by Widrow and Hoff (1960) and by Rosenblatt (1959). The delta rule in its simplest form, can be written as:

$$\Delta w_{ij} = \varepsilon \delta_i a_j \tag{5}$$

where ε is the value of the learning parameter and δ_i, the error for unit *i,* is the difference between its teaching input (t_i) and its obtained activation (a_i):

$$\delta_i = t_i - a_i \tag{6}$$

Note that if $\delta_{pi} < 0$, the adjustment to weight w_{ij} will be negative so that the influence of input *i* is reduced. The system chooses a set of weights to minimize the error. The procedure used is the gradient descent, in which the weights themselves are minimized. The system is then said to decline in weight-space and attains equilibrium when all the weights have been minimized.

For a simple network with say two input units and a single output unit, learning occurs by activating each unit, preparing an output, and then comparing this output with the teaching input. The error between the teaching input and the output of the network is then used to adjust the weights through the fixed learning parameter. The correct set of weights is approached asymptotically if the training procedure is continued through several sweeps, each of these sweeps being referred to as a *training epoch*.

Each epoch results in, theoretically, a set of weights that is closer to the perfect solution. To get a measure of the closeness of the approximation to a perfect solution, we can calculate the total sum of squared errors that result on each epoch. This measure of the state of learning of the network gets smaller over each epoch, as do the changes in the strength of the connections. Minsky and Papert (1969) have shown that the error correcting rule will find a set of weights that drives the error as close to zero as desired, provided that such a set of weights exist.

It should be noted that such weights as described above exists only if for each input-pattern – target-pair, the target can be predicted from a linear combination of the activation units. That is, the weights must satisfy the linear predictability constraint:

$$t_{ip} = \sum_j w_{ij} a_{jp} \tag{7}$$

for output unit i in all patterns p. This constraint can be overcome by the use of hidden units, which in turn introduces problems relating to the training of the network.

3.4 Training hidden units: back propagation

The application of the back propagation rule involves two phases. In the first phase, the input is propagated forward through the network to compute the output value a_{pj} for each unit (we will assume a single output unit for simplicity). This output is then compared with the target, resulting in a δ term for the output unit.

$$\delta_{pi} = (t_{pi} - a_{pi}) f'_i (net_{pi}) \tag{8}$$

where $net_{pi} = \sum_j w_{ij} a_{pj} + bias_i$ is the activation function, and $f'_i(net_{pi})$ is the first derivative of the activation function with respect to a change in the net input to the unit. The second phase involves a backward pass through the network (analogous to the initial forward pass) during which the δ term is computed for each unit in the network. In the case of the hidden units, there is no specified target so that δ is determined recursively in relation to the δ terms of the units to which they directly connect and the weights of those connections. That is:

$$\delta_{pi} = f'_i (net_{pi}) \sum_k \delta_{pk} w_{ki} \tag{9}$$

Once these two phases are complete, we can compute the weighted error derivative for each weight. These weighted error derivatives can then be used to compute actual weight changes on a pattern by pattern basis, or they may be accumulated over the ensemble of patterns.

4. RESULTS

This section presents the predictions from the estimated neural network model. These are compared with the forecasts from a univariate ARIMA model and a VECM. The ARIMA was estimated using the Box-Jenkins (1976) methodology and the best model was chosen using the Schwartz Bayesian Criteria. With respect to the VECM, the likelihood ratio test was employed to determine the most appropriate lag length.

As noted previously, the variables considered are core inflation (CORE), base money (BAS), exchange rate (EXR), oil prices (OIL) and the Treasury bill rate (TBILL). All variables are measured in logs covering the period March 1975 to December 2001. Except for oil prices, which were taken from West Texas Intermediate Crude Oil Price listings, all the variables were taken from the Bank of Jamaica's database. For the core series, initial work had been done for the period 1992 to 1999. For the purpose of this paper, the CPI was collected from the Statistical Institute of Jamaica (STATIN) for the period 1975 to 1998 and the index extended using the same methodology currently employed by BOJ in reporting core inflation. Table A, Appendix, gives the results of the unit root tests, which indicate that all the variables are I(1).

4.1 Neural network

The network was trained in RATS. One hidden layer with three nodes was specified along with an R-square of 0.95 as the convergence criteria. The network converged after 15,587 epochs. The most parsimonious model included the exchange rate and base money, at four lags, three seasonal factors, the pulse dummies, and an autoregressive term. The Treasury Bill rate was found to be insignificant.

The Llung-Box Q-statistic indicates the presence of mild serial correlation in the error term of the neural forecast. Perhaps, one way to remedy this would be the specification of a more complex network or to use alternative learning algorithms such as the feedforward training techniques. RATS unfortunately does not cover multi-layer network design, nor does it contain alternative algorithms to the back-propogation technique. For more in depth work on neural nets, software such as "Matrix Backpropagation", "WinNN" or "The Brain" would be required.[3] The presence of serial correlation indicates that the estimates from the ANN model will not have minimum variance.

4.2 ARMA

Table B in Appendix gives the SBC results of the various Box-Jenkins models, which indicate that an ARIMA (1,1,0) model was

TABLE 1. RESULT OF ARIMA (1,1,0) MODEL, SAMPLE 1975:1-1998:4

Variable	*Coefficient*	*Std. Error*	*t-Statistic*	*Prob.*
C	0.0002	0.006	0.030	0.976
Dummy 1	0.103	0.012	8.708	0.000
Dummy 2	0.110	0.013	80194	0.000
Q1	0.001	0.003	0.210	0.834
Q2	0.001	0.003	0.435	0.665
Q3	0.002	0.003	0.795	0.429
Filter Core	0.834	0.193	4.064	0.000
Core (1)	0.430	0.095	4.533	0.000
Llung-Box				
Period	Q-Statistic	Probability		
4	0.34	0.95		
8	5.19	0.64		
12	13.87	0.24		
16	15.58	0.41		
	F Statistics	R Squared	Probability	
White's Test	1.59	12.23	0.14	

R-squared: 0.80; SBC: -5.61.

[3] WinNN, for example, is a Neural Networks package for windows 3.1 and above. WinNN incorporates a very user-friendly interface with a powerful computational engine. It provides an alternative to using more expensive and hard to use packages. WinNN can implement feed forward, multi-layered NN and uses a modified fast back-propagation for training. It also has various neuron functions. It allows testing of the network performance and generalization. All training parameters can be easily modified while WinNN is training. Results can be saved on disk or copied to the clipboard. It also supports plotting of the outputs and weight distribution.

favoured. Two pulse dummies were created to take account of the two one - off shocks to core inflation. The first pulse dummy was used to capture the effect of the 1978 shock, equalling one in June 1978 and zero elsewhere. The second is a combination of a gradually changing and prolonged pulse dummy, equalling one in December 1991, and a half (0.5) in June and September 1991 and March and June 1992. Three seasonal dummies were also created to capture the seasonal patterns in the data. In addition, a time trend was included in the model.

The results of the ARIMA model are contained in table 1. The two pulse dummies were found to be statistically significant, as well as the filtered series and the AR component. The Llung-Box Q-Statistics (Table 1) indicated the absence of serial correlation along with the Breusch-Godfrey serial correlation LM Test (F statistics of 0.65 and probability of 0.69), while the White's test (Table 1) suggested that the error term was homoskedastic.

4.3 VAR

Based on the time series properties of the variables a VECM was estimated to capture both the long and short run dynamics of core inflation. The likelihood ratio tests indicated an optimal lag length of two (Table D, Appendix). The Johansen Cointegration test (Table C, Appendix) on the variables revealed one cointegrating equation at the 5% significance level.

The long-run equilibrium results (Table 2) of the VECM indicate the nature of the relationships between core inflation and its specified determinants. The exchange rate, base money, and foreign prices all have positive effects on core inflation, with the exchange rate having the greatest impact. This is consistent with *a priori* expectations. A one-unit shock to the exchange rate causes core inflation to increase by 49 per cent over the long run, while shocks to the Treasury Bill rate causes core inflation to decline marginally by 1 per cent. The effects of oil prices in the long run appear to be relatively strong.

The impulse responses from the VECM are shown in figures A, B and C (Appendix). The effect runs from oil prices to base money

TABLE 2. NORMALISED LONG RUN COEFFICIENTS

Coe Index	*Ex-rate*	*Tbill*	*Base Money*	*Oil Prices*
1.000	0.490	-0.090	0.095	0.206
Standard Error	0.263	0.034	0.205	0.139

to Treasury bill rate to exchange rate and to core inflation. A unit shock to the exchange rate has a positive impact on core inflation. This impact is highly significant for the first six quarters, after which it dies out at approximately the 34th quarter. A unit shock to base money, has a positive impact on core inflation in the second quarter after the shock, which lasts for as much as seven quarters. This is consistent with Allen (1997), where it was found that core inflation responded within three to six months after a shock to the monetary base. Likewise, a unit shock to foreign prices has a positive impact on core inflation, peaking in the second quarter before settling to its equilibrium level. A unit shock to the Treasury bill rate has a cyclical effect on core inflation up to the 38th quarter, after which it peters out. Shocks to core inflation from itself have the most significant influence, suggesting that the inflationary process in Jamaica has significant inertia. A unit shock to core inflation has an immediate, positive and significant effect on itself over the first five quarters.

Based on the impulse responses in figures B and C (Appendix), base money has a significant impact on the exchange rate, while the exchange rate has a marginal impact on base money. By deduction, the causation appears to run from base money to the exchange rate, and finally to prices. These results are consistent with those found by Robinson (1997), which showed that expansionary monetary policy has an unambiguous expansionary effect on prices, the lag effect of monetary policy was found to be at least two months and exchange rate stabilization was found to be the most effective means of short-term stabilization.

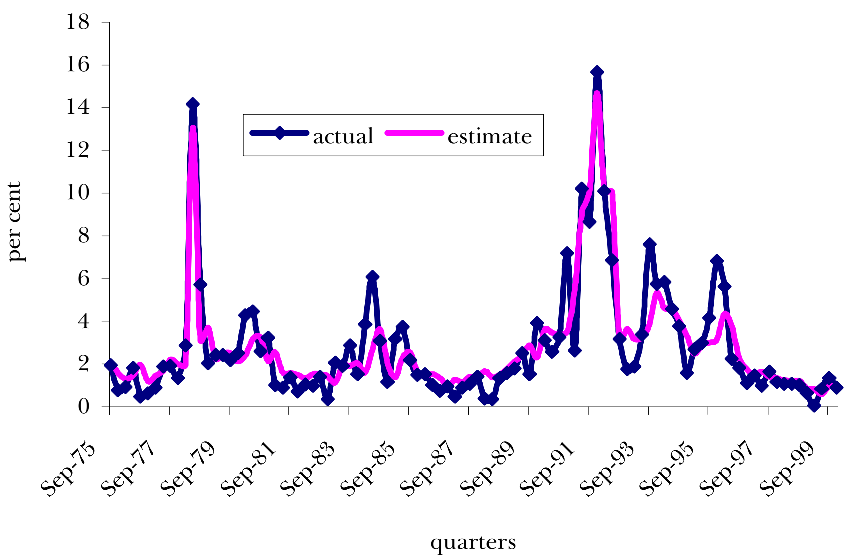
To determine the contribution of each variable to the core inflation process, the paper assesses the variance decomposition for the sixty-step-ahead forecasts. The results (Table E, Appendix) show that most of the variability in core inflation was caused mainly by shocks to itself, base money and the exchange rate, throughout the period. Shocks associated with foreign prices have a marginal effect on core inflation over the forecast horizon.

4.4 Forecast evaluation

Tables F and G (Appendix) provide a comparison of the forecasting accuracy of the three models under consideration. Based on these statistics, the ANN model has the greatest predictive power for in-sample forecasts. It has the lowest mean squared error (MSE), root mean squared error, and mean absolute error.

Figures 3, 4, and 5 show the (in sample) forecasts compared with the actual series for the three models. All the models captured the shock to core inflation in 1978. The ANN model captured the major turning points in core inflation well, although there are indications that between 1986 and 1989 and towards the end of the period it over-estimated the series. The other two models appear to have done better over these periods. Of note, the ARIMA model captured the dynamics in core inflation best, towards the end of the period. Also, the liberalization effect was best captured by the ANN model, followed by the ARIMA model. The VEC model (and to some extent the ARIMA model) did not capture the full effect of this policy, and apparently did a poor job at estimating the second to last and last spikes during the period.

FIGURE 3. ARIMA FORECAST OF CORE INFLATION (IN-SAMPLE), 1975-99



A forecast encompassing test was also used to assess the ARIMA and ANN forecasting abilities. If one model forecast encompasses another, that model's forecast is said to be unbiased, contains all the information present in the other, but contains more useful information. Failure of one model's forecasts to encompass another indicates that it is possible to gain by combining the forecasts. The results in table H (Appendix) indicated that the ANN model forecast encompasses the ARIMA model.

Despite the fact that the ANN model dominated the in-sample

FIGURE 4. VEC FORECAST OF CORE INFLATION (IN-SAMPLE), 1975-99

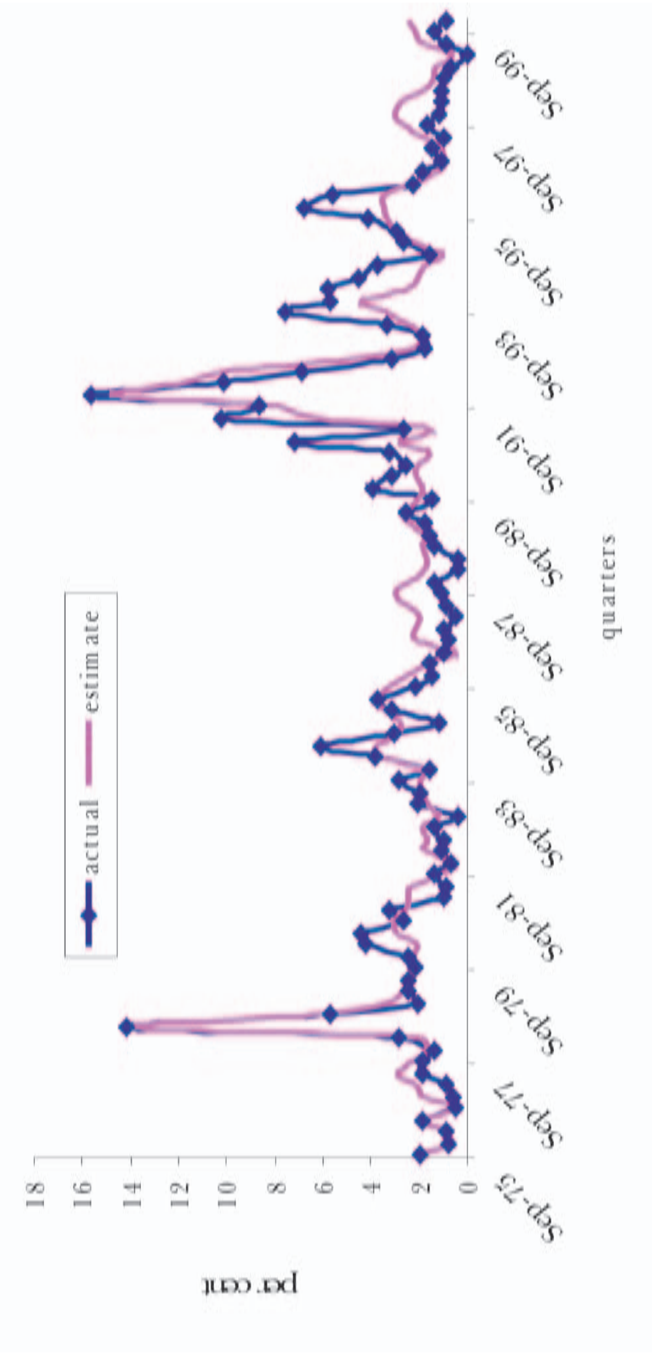
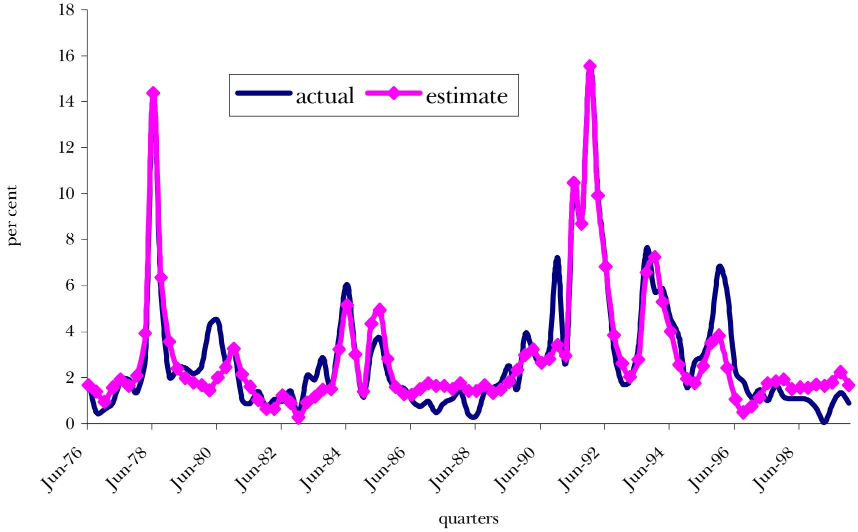


FIGURE 5. ANN FORECAST OF CORE INFLATION (IN-SAMPLE), 1976-98



forecast performance, the ARIMA and VEC models did a better job with the out-of-sample forecast. Significantly, the Janus quotient (J-Quotient) for the ARIMA model was approximately seven times smaller than the quotient for the ANN model. It should be noted that the Janus quotient is a more robust estimator than the Theil-U coefficient when dealing with out-of-sample forecast evaluations.

FIGURE 6. OUT-OF-SAMPLE FORECASTS OF CORE INFLATION, 2000-2001

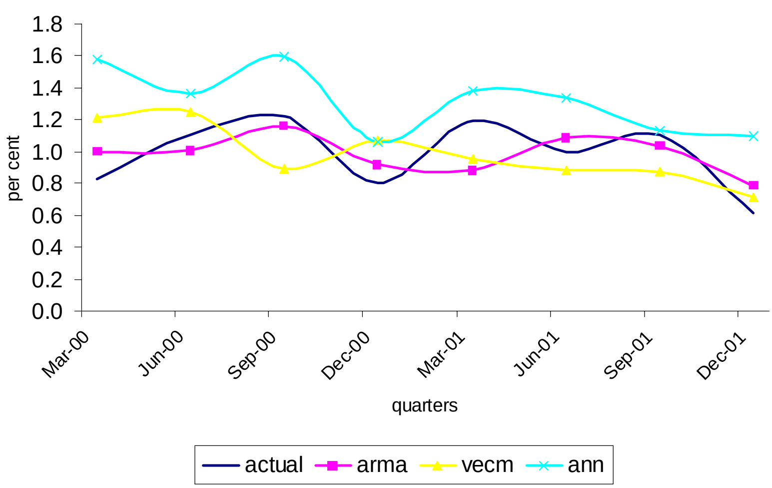


Figure 6 shows the out-of-sample forecast. The graph indicates that the ARIMA and VECM models remain closest to the actual core inflation out-turn between the first quarter of 2000 and the last quarter of 2001. The ANN model, although suggesting a higher inflation out-turn, captured the major turning points in the core series. The ARIMA model captured only the first turning point in core inflation while the VECM model did not capture any of the major turning points. Thus although the statistical measures suggest that the forecasts from the univariate model is superior, the ANN model seems to have been able to capture the data generating process more adequately. The omission of an income variable and possible shifts in the relations among the macroeconomic variables over time may have limited the forecasting power of the multivariate models.

5. CONCLUSION

Based on the results from the models, volatility in core inflation was due mainly to innovations to itself, to base money, the exchange rate, and to a lesser extent foreign prices. The results of the impulse responses indicated that core inflation responded immediately to a shock to base money.

Of the three models estimated the ANN model was the most appropriate in making in-sample forecast of core inflation. The ARIMA model performed marginally better than the VECM for in-sample forecasts. In addition, the ARIMA model performed better than the VECM when making out-of-sample forecast. The ANN model was the only model that captured the major turning points in core inflation when making out of sample forecast. Whilst further work is required in terms of the type of propagation mechanism to be used, the analysis suggests that ANN model can be a useful addition to the set of tools used in analyzing and forecasting inflation in Jamaica.

Appendix

TABLE A. AUGMENTED DICKEY-FULLER TEST

Variables	*T Statistics*		*Lag*
	Levels	*First Difference*	
Core	-1.85	-3.8	1
Base Money	-0.46	-3.65	4
Exchange Rate	-2.58	-5.23	1
Tbill	-1.44	-4.6	9
Oil Prices	-2.25	-8.09	1
5% critical value	-3-45	-3.45	n/a
1% critical value	-4.04	-4.04	n/a

TABLE B. MODEL SELECTION USING AIC AND SBC CRITERIA: BOX-JENKINS MODELS

Models	*AIC*	*SBC*
ARIMA (1,1,1)	-5.802	-5.565
ARIMA (1,1,2)	-5.802	-5.566
ARIMA (1,1,0)	-5.823	-5.612
ARIMA (1,1,4)	-5.803	-5.565
ARIMA (0,1,4)	-5.639	-5.43

TABLE C. JOHANSEN COINTEGRATION TEST, SAMPLE 1976:4-1999:4

Null Hypothesis	*Eigenvalue*	*Likelihood Ratio*	*5% Critical Value*	*1% Critical Value*
r = 0	0.605	117.146	68.52	76.07
r < 1	0.174	34.425	47.21	54.46
r <2	0.15	17.401	29.68	35.65
r < 3	0.033	2.973	15.41	20.04

NOTE: Likelihood ratio test indicates one cointegrating equation at the 5% significance level.

TABLE D. RESULTS OF LIKELIHOOD RATIO TEST

Lag Lengths	*Chi-Square Ratio*	*Significance Level*
8 vs 4	90.39	0.74
4 vs 2	57.64	0.21
2 vs 1	76.03	0.00
	SBC	
2	-5.12	
4	-4.73	
8	-396	

TABLE E. VARIANCE DECOMPOSITION OF CORE INFLATION (%)

Periods	*SE*	*Core Inflation*	*Exchange Rate*	*Treasury Bill*	*Base Money*	*Oil Prices*
5	0.04	84.50	3.68	1.50	9.15	1.21
10	0.06	81.14	4.21	1.48	12.27	0.90
15	0.08	79.86	4.60	1.57	13.10	0.86
20	0.09	79.65	4.77	1.38	13.37	8.83
25	0.11	79.41	4.82	1.35	13.61	0.81
30	0.12	79.23	4.89	1.33	13.75	0.80
35	0.13	79.16	4.92	1.29	13.89	0.79
40	0.13	79.08	4.95	1.28	13.91	0.79
45	0.14	79.02	4.97	1.26	13.96	0.78
50	0.15	78.98	4.99	1.25	14.00	0.78
55	0.16	78.94	5.00	1.24	14.04	0.78
60	0.17	78.91	5.01	1.23	14.07	0.77

TABLE F. MODEL FORECAST EVALUATIONS, IN-SAMPLE FORECAST, SAMPLE 1975:1-1999:04

Model	*MSE*	*RMSE*	*MAE*	*Theil U*	*Janus*
ARMA	0.00150	0.013	0.0085	0.16	0.02
VEC	0.00020	0.014	0.0106	0.18	0.03
Neural	0.00010	0.012	0.0072	0.13	0.15

TABLE G. MODEL FORECAST EVALUATIONS, IN-SAMPLE FORECAST: SAMPLE 2000:1-2001:04

Model	*MSE*	*RMSE*	*MAE*	*Theil U*	*Janus*
ARMA	0.00000	0.0016	0.0014	0.10	0.02
VEC	0.00001	0.0025	0.0023	0.12	0.03
Neural	0.00002	0.0039	0.0034	0.17	0.15

TABLE H. FORECAST ENCOMPASSING TEST RESULTS, SAMPLE 1975:1-1999:4

Null Hypothesis	*F-Statistics*	*P-Value*
ARMA	3.9294	0.0254
ANN	0.3828	0.6832

BIBLIOGRAPHY

Allen, Courtney, 1997, "Measuring Core Inflation in Jamaica", *Bank of Jamaica Working Papers*, September 2000, pp. 79 – 104.

Bond, Marian, 1980, "Exchange Rates, Inflation, and the Vicious Circle", *Finance and Development*, (March), pp. 27 – 31.

Bourne, C., and W. Persaud, 1977, "Financial Variables in the Inflationary Process: Two Caribbean Cases", *in Inflation in the Caribbean*, ed. By C. Bourne (Kingston: Institute of Social and Economic Research), pp. 54 – 71.

Claus, Iris, 1997, "A Measure of Underlying Inflation in the United States", *Bank of Canada Working Paper,* 97-20.

Diebold, Francis X., 1998, "Elements of Forecasting", University of Pennsylvania, South-Western College Publishing.

Downes, Andrew S., 1992, "Macroeconomic Adjustment and the Inflation Experience of CARICOM Countries," paper presented at the XXIX Meeting of Technicians of Central Banks of the American Continents, Bridgetown, Barbados.

Enders, Walter, 1995, "Applied Econometric Time Series", Canada, John Wiley & Sons.

Enders, Walter, 1996, "Rats Handbook For Econometric Time Series", United States of America, John Wiley & Sons.

Engert, Walter & Scott Hendry, 1997, "Forecasting Inflation with the M1- VECM: Part 2", Department of Monetary and Financial Analysis, Bank of Canada.

Gately, Edward, 1996, "Neural Networks for Financial Forecasting", United States of America, John Wiley & Sons, Inc.

Holden, Ken, 1995, "Vector Autoregression Modeling and Forecasting", *Journal of Forecasting*, Vol. 14, pp. 159 – 166.

McCulloch, W. S., & Pitts, W., 1943, "A Logical Calculus of the Ideas Immanent in Nervous Activity", *Bulletin of Mathematical Biophysics,* 5, pp. 115 – 133.

Minsky, M., 1954, "Neural Nets and the Brain-Model Problem", Unpublished doctoral dissertation, Princeton University.

Minsky, M., & Papert, S., 1969, "Perceptrons", Cambridge, MA: MIT Press.

Norden, Simon van, 1992, "A Note on Forecast Encompassing Tests", Canada.

Quah, Danny T, & Shaun P Valley, 1995, "Measuring Core Inflation", *Bank of England Working Paper Series no. 31*.

Remus, William & Marcus O'Conner, 1997, "Neural Networks for Time Series Forecasting", Principles of Forecasting.

Robinson, Wayne, 1996, "Forecasting Inflation using VAR Analysis", *Bank of Jamaica Research Papers*, September 2000, 55 - 78.

Roger, Scott, 1995, "Measures of Underlying Inflation in New Zealand, 198195," *Reserve Bank of New Zealand Discussion Paper Series No. G95/5*.

Rosenblatt, F., 1959, "Two Theorems of Statistical Separability in the Perceptron", in *Mechanisation of Thought Processes: Proceed-*

ings of a Symposium Held at the National Physical Laboratory, November 1958, Vol. 1, pp. 421 – 456. London: HM Stationery Office.

Rumelhart, David & James McClelland, 1988, "Parallel Distributed Processing", United States of America, Massachusetts Institute of Technology.

Stanley, Jeannette, 1990, "Introduction to Neural Networks, Computer Simulations of Biological Intelligence", Printed in the United States of America.

Stennett, Robert, 1998, "Projecting Velocity in Jamaica", Research Department, Bank of Jamaica.

Thomas, Desmond, 1994, "Estimating a Monetary Model of Inflation Determinants in Jamaica," paper presented at the XXVI Annual Conference of the Regional Programme of Monetary Studies.

Warren McCullock's and Walter Pitts', "A logical calculus of ideas immanent in Nervous Activity" (Bulletin of Mathematical Biophysics Volume 5, pp 115-133),

Webb, Roy H., 1995, "Forecasts of Inflation from VAR Models", *Journal of Forecasting*, Vol. 14, 267 – 285.

Widrow, G., & Hoff, M. E., 1960, " Adaptive Switching Circuits", *Institute of Radio Engineers, Western Electronic Show and Convention, C*onvention Record, Part 4, pp. 96 – 104.

Víctor Olivo

Taylor rules and inflation targeting do not work with systematic foreign exchange market intervention

1. INTRODUCTION

Inflation targeting became widely popular among central banks during the 1990s. The first wave of central banks adopted inflation targeting in the first half of this decade in industrialized countries. In the late 1990s, however, several Latin American central banks embarked on inflation targeting. Currently, the central banks of Chile, Mexico, Brazil and Colombia practice inflation targeting.

Jointly with the spreading of inflation targeting, an explosion occurred in the literature that covers different practical and theoretical aspects of this monetary policy regime. Bofinger (2001), however, points out that in contrast to monetary targeting that was applied by central banks after intensive academic discussion,

Paper prepared by V. Olivo, Economist, Vice-presidency of Research, Central Bank of Venezuela. The views expressed in this paper are those of the author and should not be interpreted as those of the Central Bank of Venezuela. *The earlier version of this paper was also presented at the VII Meeting of the Central Banks Research Network of the Americas hosted by the Central Bank of Guatemala, in Guatemala City, November 22-24, 2002. S. M. Miller made important contributions to previous versions of the paper. Any remaining errors are the sole responsibility of the author.* *E-mail address: volivo@bcv.org.ve*

inflation targeting was first put in practice by central banks and then researchers began to provide the theoretical support. The literature available on inflation targeting has covered an ample spectrum of topics, such as: i) country experiences with the scheme (Leiderman and Svensson, 1995, Bernanke et al. 1999); ii) implementation of inflation targeting as inflation forecast targeting (Svensson, 1996); iii) analysis of the responses to different shocks, consequences of model uncertainty, effects of interest rate smoothing and stabilization, comparison with nominal GDP targeting, and implications of forward-looking behavior (Svensson, 1997); iv) studying inflation targeting in the context of monetary policy rules (Svensson, 1998); v) examining the efficiency of inflation targeting in relation to Taylor rules in closed economies (Ball, 1997) and open economies (Ball, 1998).

Although inflation targeting currently reflects the monetary policy strategy choice of those countries that want to provide the economy with a nominal anchor while avoiding the excessive rigidity of "hard" or "soft" pegs, Calvo and Reinhart (2000) argue that many countries that say they allow their exchange rate to float actually do not. Those authors content that an epidemic case exists of "fear of floating."

This paper studies how the systematic attempt to influence directly the path of the nominal exchange rate due to "fear of floating" affects the conduct of monetary policy under a Taylor rule and inflation targeting. The paper demonstrates that implementing a Taylor rule or an inflation-targeting scheme simultaneously with policies that try to moderate the rate of depreciation of the nominal exchange rate may result in an inconsistent monetary policy. In addition, the paper shows in a relatively simple framework the close theoretical connection between a Taylor rule and inflation targeting.

This issue is important, because as mentioned previously, inflation targeting has spread rapidly in countries that may experience simultaneously "fear of floating".

The paper is organized as follows. Section 2 examines the relationship between a Taylor rule and inflation targeting in a closed economy. Section 3 extends the analysis to an open economy. In particular, it focuses on the effect in a Taylor rule and inflation targeting of adopting simultaneously an exchange rate system based on a crawling peg. This analysis implicitly assumes imperfect capital mobility, so the monetary authority can exert some degree of control over the domestic interest rate. Section 4 examines the effect in a Taylor rule and inflation targeting regime of systematic intervention in the foreign exchange market to curb the rate of depreciation of the nominal exchange rate without pursuing any precise target for that variable. Section 5 concludes.

2. TAYLOR RULES AND INFLATION TARGETING: CLOSED ECONOMY

The closed economy model developed here closely follows Ball (1997). It consists of a dynamic IS function, and a Phillips curve. In contrast to Ball, I substitute the real interest rate (r) using the Fisher ex post relationship, and solve the model for the nominal interest rate (i). I measure all variables as deviations with respect to their respective means in logarithms.

In this model, it takes two periods for monetary policy to affect inflation (π) through changes in the interest rate -- one period for policy to affect output (y), and one period for output to affect inflation. The model includes the following equations:

IS curve:

$$y_t = -\beta r_{t-1} + \lambda y_{t-1} + \varepsilon_t . \tag{1}$$

Phillips curve:

$$\pi_t = \pi_{t-1} + \alpha y_{t-1} + \eta_t . \tag{2}$$

Fisher ex post relation:

$$i_{t-1} = r_{t-1} + \pi_{t-1} . \tag{3}$$

Substituting this relation into equation (1) produces:

$$y_t = -\beta i_{t-1} + \beta \pi_{t-1} + \lambda y_{t-1} + \varepsilon_t , \tag{1a}$$

and

$$\pi_t = \pi_{t-1} + \alpha y_{t-1} + \eta_t . \tag{2a}$$

For period t+1, the model is written as follows:

$$y_{t+1} = -\beta i_t + \beta \pi_t + \lambda y_t + \varepsilon_{t+1} , \tag{1b}$$

and

$$\pi_{t+1} = \pi_t + \alpha y_t + \eta_{t+1} . \tag{2b}$$

Taylor rule

The transmission mechanism of monetary policy assumed in the model implies that the policymaker minimizes the following expected quadratic loss function in terms of output and inflation variance:

$$E_t(L) = cE_t(y_{t+1})^2 + E_t(\pi_{t+2})^2 . \quad (4)$$

Substituting the assumed structure of the economy in the loss function yields:

$$E_t(L) = c(-\beta i_t + \beta\pi_t + \lambda y_t)^2 + [E_t(\pi_{t+1}) + \alpha E_t(y_{t+1})]^2 , \quad (4a)$$

and

$$E_t(L) = c(-\beta i_t + \beta\pi_t + \lambda y_t)^2 + [\pi_t + \alpha y_t + \alpha(-\beta i_t + \beta\pi_t + \lambda y_t)]^2 \quad (4b)$$

Minimizing E(L) with respect to i_t and solving for i_t produces:

$$i_t = [\frac{\lambda}{\beta} + \frac{\alpha^2}{\beta(c+\alpha^2)}]y_t + [1 + \frac{\alpha^2}{\alpha\beta(c+\alpha^2)}]\pi_t . \quad (5)$$

Hence, this model generates an optimal policy rule, Taylor rule, where its parameters depend on the structure of the economy and the weight that the policymaker assigns to output variability in the loss function (c).

When the policymaker assigns no weight to output fluctuations (c=0), the interest rate rules reduces to:

$$i_t = \frac{(1+\lambda)}{\beta} y_t + (1 + \frac{1}{\alpha\beta})\pi_t . \quad (6)$$

Strict inflation targeting

This section demonstrates that inflation targeting implicitly implies an interest rate rule.

Ball (1997) defines strict inflation targeting as a policy that minimizes the variance of inflation around its average level. With monetary policy affecting inflation with a two-period lag, we define strict inflation targeting as follows:

$$E_t(\pi_{t+2}) = 0 . \quad (7)$$

Using equation (2) updated two periods causes equation (8) to equal:

$$E_t(\pi_{t+1}) + \alpha E_t(y_{t+1}) = 0\,. \tag{8}$$

Substituting equations (1b) and (2b) into this expression results in:

$$\pi_t + \alpha y_t + \alpha(-\beta i_t + \beta\pi_t + \lambda y_t) = 0\,. \tag{9}$$

Solving this expression for i_t, produces:

$$i_t = \frac{(1+\lambda)}{\beta} y_t + (1 + \frac{1}{\alpha\beta})\pi_t\,. \tag{10}$$

The previous equation equals the Taylor rule with c=0 (i.e., Equation 5). Hence, strict inflation targeting matches a Taylor rule where the policymaker gives zero weight to output variance. Because the Taylor rule is the optimal rule in this model, strict inflation targeting is also an optimal rule for the particular case that c=0.

Although the policymaker assigns a zero weight to output fluctuations, the implicit policy rule still responds to this variable due to the structure of the model.

Flexible inflation targeting

Inflation targeting more generally equals a partial adjustment rule where expected inflation in period t+2 equals a fraction of expected inflation in period t+1. That is:

$$E_t(\pi_{t+2}) = \theta E_t(\pi_{t+1})\,, \quad \theta<1. \tag{11}$$

Once again, using equation (2) updated two periods, equation (11) equals:

$$E_t(\pi_{t+1}) + \alpha E_t(y_{t+1}) = \theta E_t(\pi_{t+1})\,. \tag{12}$$

Substituting equations (1b) and (2b) into this expression and re- arranging terms leads to:

$$(1-\theta)(\pi_t + \alpha y_t) + \alpha(-\beta i_t + \beta\pi_t + \lambda y_t) = 0\,. \tag{13}$$

Solving for i_t produces:

$$i_t = [\frac{(1-\theta)+\lambda}{\beta}]y_t + [1+\frac{(1-\theta)}{\alpha\beta}]\pi_t \,. \quad (14)$$

If we define:

$$(1-\theta) = \frac{\alpha^2}{c+\alpha^2} \text{ or } \theta = \frac{c}{c+\alpha^2}\,, \quad (15)$$

and substitute it into equation (14), then equation (5) emerges. This implies that flexible inflation targeting is, in general, the optimal policy rule in this model. The value of θ implicitly depends on the weight that the policymaker assigns to output volatility in the loss function (c). A larger value of θ, that is a more gradual path toward the inflation target, associates with a larger c, given the value of α.

3. TAYLOR RULES AND INFLATION TARGETING: CRAWLING PEG

This section develops a model similar to that of Ball (1998), but assumes that the exchange rate follows a crawling peg. Combining a Taylor rule or inflation targeting with a crawling peg may seem strange at first, but that policy strategy currently reflects the Hungarian and Israeli situations (IMF-IFS, May 2002).

In the open economy model, I introduce the logarithm of the real exchange rate (q) into the IS function and the change of the nominal exchange rate (e_t- e_{t-1}) into the Phillips function. The rate of the crawl (Ω) is specified as a fraction of the differential between domestic and foreign inflation, where the latter is normalized to zero $(0 \leq \Omega \leq 1)$). Thus, the model setup is as follows:

IS curve:

$$y_t = -\beta r_{t-1} + \lambda y_{t-1} + \delta q_{t-1} + \varepsilon_t \,. \quad (16)$$

Phillips curve:

$$\pi_t = \pi_{t-1} + \alpha y_{t-1} + \psi(e_{t-1} - e_{t-2}) + \eta_t \,. \quad (17)$$

Fisher ex post relation:

$$i_{t-1} = r_{t-1} + \pi_{t-1} \,. \quad (18)$$

Crawling peg rule:

$$e_{t-1} - e_{t-2} = \Omega\pi_{t-1}; \pi^*_{t-1} = 0 . \tag{19}$$

Definition of the real exchange rate:

$$q_{t-1} - q_{t-2} = (e_{t-1} - e_{t-2}) - \pi_{t-1}; \tag{20}$$

$$q_{t-1} = q_{t-2} + (\Omega - 1)\pi_{t-1}. \tag{21}$$

Substituting the previous definitions into equations (16) and (17) yields:

$$y_t = -\beta i_{t-1} + (\beta + \delta(\Omega - 1))\pi_{t-1} + \lambda y_{t-1} + \delta q_{t-2} + \varepsilon_t , \tag{22}$$

and

$$\pi_t = (1 + \psi\Omega)\pi_{t-1} + \alpha y_{t-1} + \eta_t . \tag{23}$$

Moving one period ahead leads to:

$$y_{t+1} = -\beta i_t + (\beta + \delta(\Omega - 1))\pi_t + \lambda y_t + \delta q_{t-1} + \varepsilon_{t+1} , \tag{22a}$$

and

$$\pi_{t+1} = (1 + \psi\Omega)\pi_t + \alpha y_t + \eta_{t+1} . \tag{23a}$$

In this model, the introduction of the crawling peg rule maintains the two-period lag with which monetary policy affects inflation. The crawling rule breaks the explicit connection between the exchange rate and the interest rate. The interest rate, however, still affects the behavior of the exchange rate indirectly through its effect on the amount of intervention necessary to maintain the crawling rule.

Taylor rule

As before, the loss function equals:

$$E_t(L) = cE_t(y_{t+1})^2 + E_t(\pi_{t+2})^2 . \tag{24}$$

Substituting the equations of the model into the loss function produces:

$$\begin{aligned} E_t(L) = c[-\beta i_t + (\beta + \delta(\Omega - 1))\pi_t + \lambda y_t + \delta q_{t-1}]^2 \\ + [(1 + \psi\Omega)E_t(\pi_{t+1}) + \alpha E_t(y_{t+1})]^2 \end{aligned} , \tag{25}$$

and then

$$E_t(L) = c\{-\beta i_t + (\beta + \delta(\Omega - 1))\pi_t + \lambda y_t + \delta q_{t-1}]\}^2$$
$$+\{(1+\psi\Omega)[(1+\psi\Omega)\pi_t + \alpha y_t] + \alpha[-\beta i_t + (\beta + \delta(\Omega-1))\pi_t + \lambda y_t + \delta q_{t-1}]\}^2 \quad (26)$$

Minimizing E(L) with respect to i_t and solving for i_t leads to:

$$i_t = [\frac{\lambda}{\beta} + \frac{\alpha^2(1+\psi\Omega)}{\beta(c+\alpha^2)}]y_t + [1 + \frac{\delta(\Omega-1)}{\beta} + \frac{\alpha\ (1+\psi\Omega)^2}{\beta(c+\alpha^2)}]\pi_t + \frac{\delta}{\beta}q_{t-1} \quad (27)$$

The optimal policy in this model equals a Taylor rule plus a term that depends on the real exchange rate lagged one period. Now, the parameters of the policy rule depend positively on the rate of the crawl (Ω).

The case where c=0 yields the following interest rate rule:

$$i_t = [\frac{\lambda}{\beta} + \frac{(1+\psi\Omega)}{\beta}]y_t + [1 + \frac{\delta(\Omega-1)}{\beta} + \frac{(1+\psi\Omega)^2}{\alpha\beta}]\pi_t + \frac{\delta}{\beta}q_{t-1}. \quad (28)$$

Strict inflation targeting

As in the closed economy model, strict inflation targeting implies that:

$$E_t(\pi_{t+2}) = 0\ . \quad (29)$$

This leads to:

$$(1+\psi\Omega)E_t(\pi_{t+1}) + \alpha E_t(y_{t+1}) = 0\ , \quad (30)$$

or

$$(1+\psi\Omega)[(1+\psi\Omega)\pi_t + \alpha y_t] + \alpha[-\beta i_t + (\beta + \delta(\Omega-1))\pi_t + \lambda y_t + \delta q_{t-1}] = 0\ . \quad (30a)$$

Solving for i_t results in:

$$i_t = [\frac{\lambda}{\beta} + \frac{(1+\psi\Omega)}{\beta}]y_t + [1 + \frac{\delta(\Omega-1)}{\beta} + \frac{(1+\psi\Omega)^2}{\alpha\beta}]\pi_t + \frac{\delta}{\beta}q_{t-1}. \quad (31)$$

Strict inflation targeting equals the optimal Taylor rule with c=0.

Flexible inflation targeting

As in the closed economy case, a gradual inflation targeting is defined as follows:

$$E_t(\pi_{t+2}) = \theta E_t(\pi_{t+1}) . \tag{32}$$

Equation (32) can be re-written as follows:

$$(1+\psi\Omega)E_t(\pi_{t+1}) + \alpha E_t(y_{t+1}) = \theta E_t(\pi_{t+1}) , \tag{33}$$

or

$$\begin{aligned} &[(1+\psi\Omega)-\theta][(1+\psi\Omega)\pi_t + \alpha y_t] + \\ &\alpha[-\beta i_t + (\beta + \delta(\Omega-1))\pi_t + \lambda y_t + \delta q_{t-1}] = 0 \end{aligned} . \tag{33a}$$

Solving for i_t yields:

$$\begin{aligned} i_t = &[\frac{\lambda}{\beta} + \frac{(1-\theta)+\psi\Omega}{\beta}]y_t + \{1 + \frac{\delta(\Omega-1)}{\beta} + \frac{(1+\psi\Omega)[(1-\theta)+\psi\Omega]}{\alpha\beta}\}\pi_t \\ &+ \frac{\delta}{\beta} q_{t-1} \end{aligned} . \tag{34}$$

If we define:

$$(1-\theta)+\psi\Omega = \frac{\alpha^2}{c+\alpha^2}(1+\psi\Omega) \text{ or } \theta = \frac{c(1+\psi\Omega)}{c+\alpha^2} ,$$

and substitute this expression into equation (34), we obtain equation (27). Therefore, flexible inflation targeting is, in general, an optimal policy. Note that when c=0, then equations (28), (31) and (34) are all identical.

Taylor rules and inflation targeting under a crawling peg regime

The previous results show that the parameters of the nominal interest rate rule depend positively on the rate of the crawl (Ω). Setting a relatively low value of Ω to reduce inflation towards some target value, diminishes the response of the interest rate for a given deviation of output and inflation with respect to their respective means. In addition, as long as the crawling peg regime produces an appreciation of the real exchange rate, the last term of the policy rule also indicates a lower interest rate for given deviations of output and inflation. The reduced response of the interest rate, in turn, generates pressures on inflation by rising aggregate demand, and on international reserves as it increases the amount of intervention in the foreign exchange market necessary to support the crawling rule.[1] Hence, in a crawling peg regime, the interest rate rule or its inflation targeting counterpart may generate an interest rate too low to maintain the peg. This is a result that one can intuitively expect. If a policymaker that observes a positive gap between actual and target inflation chooses to reduce the rate of depreciation of the nominal exchange rate through direct intervention in the foreign exchange market, then she wants to avoid a direct tightening of monetary policy to close this gap.

4. TAYLOR RULES AND INFLATION TARGETING: SYSTEMATIC FOREIGN EXCHANGE MARKET INTERVENTION

This section develops an analysis of the effect on Taylor rules and inflation targeting of systematic intervention of the monetary authority in the foreign exchange market to moderate the rate of depreciation of the nominal exchange rate. In contrast to the crawling peg model, this section focuses on the case where policymakers do not have an explicit target for the nominal exchange rate, but by trying to control what they perceive as high rates of depreciation not related to fundamentals, direct intervention in the foreign exchange market become systematic. In this model the amount of intervention is an exogenous variable.

We specify a model similar to that of Ball (1998), but assume that the rate of depreciation of the nominal exchange rate depends negatively on the interest rate and the systematic intervention of the monetary authority in the foreign exchange market (fx). The model includes the following equations:

IS curve:

$$y_t = -\beta r_{t-1} + \lambda y_{t-1} + \delta q_{t-1} + \varepsilon_t . \tag{35}$$

Phillips curve:

$$\pi_t = \pi_{t-1} + \alpha y_{t-1} + \psi(e_{t-1} - e_{t-2}) + \eta_t . \tag{36}$$

Fisher ex post relation:

$$i_{t-1} = r_{t-1} + \pi_{t-1} . \tag{37}$$

[1] Thus in this model, the amount of intervention in the foreign exchange market is an endogenous variable that depends on the interest rate.

Nominal exchange rate adjustment:

$$e_{t-1} - e_{t-2} = -\tau i_{t-1} - \gamma f x_{t-1} + \upsilon_{t-1}\,. \tag{38}$$

Definition of the real exchange rate:

$$q_{t-1} = e_{t-1} - \pi_{t-1}; \pi^{*}_{t-1} = 0\,. \tag{39}$$

Substituting the previous definitions in the equations of the basic model and moving one period ahead, yields the following:

$$y_{t+1} = -(\beta + \delta\tau) i_t + (\beta - \delta)\pi_t + \lambda y_t + \delta(e_{t-1} - \gamma f x_t) + \varepsilon_{t+1} + \delta v_{t+1}, \tag{40}$$

and

$$\pi_{t+1} = \pi_t + \alpha y_t - \psi\tau i_t - \psi\gamma f x_t + \eta_{t+1} + \psi v_{t+1}\,. \tag{41}$$

Taylor rule

In contrast to the previous model, in the present monetary policy affects both output and inflation with one period lag. Hence, in this case, the loss function to minimize is the following:

$$E_t(L) = cE_t(y_{t+1})^2 + E_t(\pi_{t+1})^2\,. \tag{42}$$

Minimizing E(L) with respect to i_t and solving for i_t subject to the assumed structure of the economy produces:

$$i_t = \frac{c\lambda(\beta+\delta\tau) + \alpha\psi\tau}{c(\beta+\delta\tau)^2 + (\psi\tau)^2} y_t + \frac{c(\beta-\delta)(\beta+\delta\tau) + \psi\tau}{c(\beta+\delta\tau)^2 + (\psi\tau)^2} \pi_t + \frac{c\delta(\beta+\delta\tau)}{c(\beta+\delta\tau)^2 + (\psi\tau)^2} e_{t-1} - \frac{c\delta\gamma(\beta+\delta\tau) + \psi^2\gamma\tau}{c(\beta+\delta\tau)^2 + (\psi\tau)^2} f x_t\,. \tag{43}$$

Notice that in (43) the interest rate responds to the lagged value of the nominal exchange rate and to the level of intervention in the foreign exchange market.

Ball (1998) holds that "In open economies, inflation targets and Taylor rules are suboptimal unless they are modified in important ways. Different rules are required because monetary policy affects the economy through exchange-rate as well as interest-rate channels" (p. 1). The interest-rate rule in equation (43) -- assuming zero intervention in the foreign exchange market -- differs from a standard Taylor rule, because it includes the exchange rate lagged one period. This term captures the effect of the exchange rate on inflation that Ball (1998) considers the source of suboptimality of the standard Taylor rule in open economies.[2]

The case where c=0 generates the following Taylor rule:

$$i_t = \frac{\alpha}{\psi\tau} y_t + \frac{1}{\psi\tau}\pi_t - \frac{\gamma}{\tau} fx_t \tag{44}$$

Strict inflation targeting

In this case, we define strict inflation targeting as follows:

$$E_t(\pi_{t+1}) = 0\,. \tag{45}$$

Substituting equation (36) into the inflation targeting definition results in the following expression:

$$\pi_t + \alpha y_t - \psi\tau i_t - \psi\gamma fx_t = 0\,. \tag{46}$$

Solving for i_t yields an expression equivalent to equation (44):

$$i_t = \frac{\alpha}{\psi\tau} y_t + \frac{1}{\psi\tau}\pi_t - \frac{\gamma}{\tau} fx_t\,. \tag{47}$$

In equations (44) and (47), the lagged exchange rate term does not appear, so as Ball (1998) argues, strict inflation targeting may not be optimal in an open economy. Whether the reaction of the interest rate rule to foreign exchange intervention is consistent with inflation targeting is analyzed below.

Flexible inflation targeting

As discussed by Ball (1998), monetary policy in this model can control inflation period by period. In this context, the definition of flexible inflation targeting (i.e., $E_t\pi_{t+2}=\theta\ E_t\pi_{t+1}$) does not determine a unique policy rule. Hence, it is not possible to identify the Taylor rule equivalent to flexible inflation targeting in this model. Ball (1998) proposes, however, that targeting long-run inflation defined as $\pi_t^L = \pi_t - \psi e_{t-1}$ makes possible to obtain an op-

[2] How equation (43) compares to Ball´s rule based on a monetary condition index in terms of optimality is beyond the scope of this paper.

erative definition of flexible inflation targeting. Substituting this definition of inflation in equations (40) and (41) yields:

$$y_{t+1} = -(\beta + \delta\tau)i_t + (\beta - \delta)\pi_t^L + \lambda y_t + [\delta + (\beta - \delta)\psi]e_{t-1} - \delta\gamma f x_t + \varepsilon_{t+1} + \delta v_{t+1}, \quad (48)$$

and

$$\pi_{t+1}^L = \pi_t^L + \alpha y_t + \eta_{t+1}. \quad (49)$$

The definition of flexible inflation targeting is now cast in terms of long-run inflation:

$$E_t(\pi_{t+2}^L) = \theta E_t(\pi_{t+1}^L) \ . \quad (50)$$

Making the appropriate substitutions into equation (50) and solving for i_t produces:

$$i_t = [\frac{(1-\theta)}{(\beta + \delta\tau)} + \frac{\lambda}{(\beta + \delta\tau)}]y_t + [\frac{(1-\theta)}{\alpha(\beta + \delta\tau)} + \frac{(\beta - \delta)}{(\beta + \delta\tau)}]\pi_t^L + [\frac{\delta + (\beta - \delta)\psi}{(\beta + \delta\tau)}]e_{t-1} - [\frac{\delta\gamma}{(\beta + \delta\tau)} f x_t] \ . \quad (51)$$

If we define:

$$(1-\theta) = \frac{\alpha^2}{c + \alpha^2} \text{ or } \theta = \frac{c}{c + \alpha^2},$$

and substitute this expression into equation (51), we obtain an equation equal to the one derived by minimizing the expected loss function $E_t(L) = cE_t(y_{t+1})^2 + E_t(\pi_{t+2}^L)^2$ with respect to i_t subject to the assumed structure of the economy. The value of θ is equal to that derived in the closed economy model.

What is important to highlight, however, is that this result does not change the fact that intervention in the foreign exchange market appears as a variable in the interest rate rule.

Taylor rules and inflation targeting with systematic intervention in the foreign exchange market

Equations (43), (44), and (47) show that systematic intervention in the foreign exchange market to decrease the rate of depreciation of the nominal exchange rate reduces the response of the interest rate to deviations in output and inflation from their mean values. In contrast to the crawling peg case, intervention in the foreign exchange market appears as an explicit variable in the Taylor rule instead of affecting its parameters.

Again, the reduced response of the interest rate is a result that one can intuitively expect. If a policymaker that observes a positive gap between actual and target inflation chooses to moderate the rate of depreciation of the nominal exchange rate through direct intervention in the foreign exchange market, then she wants to avoid a direct tightening of monetary policy to close this gap. This diminished response of the interest rate, however, generates pressures on the inflation rate by increasing aggregate demand and on the nominal exchange rate to depreciate, thus reducing the credibility of the anti-inflation stance of monetary policy. Therefore, policymakers concerned with the effects of the behavior of the exchange rate on the inflation target should adopt a monetary policy strategy that adjusts the interest rate to changes in the nominal exchange rate, and avoids direct intervention in the foreign exchange market.

Intervention in the foreign exchange market under inflation targeting has received some attention lately. In an inflation targeting seminar organized by the Bank of Mexico (March 4-5, 2002), a panel discussion was held on this topic. The participant from the Central Bank of Chile argued that central banks that choose to intervene in the foreign exchange market in the context of inflation targeting should not compromise the credibility of the policy regime. In general, the panel agreed that non-systematic and pre-defined (in time and amount) intervention to attenuate fluctuations of the exchange rate due to factors not directly related to fundamentals can occur along with inflation targeting. Very occasional interventions in a context where fundamentals are robust, may signal to economic agents the policymakers´perception that strong movements of the nominal exchange rate are not justified. For example, policymakers in country A may decide to intervene in the foreign exchange market for a very short time and in a limited amount, if they perceive that a financial crisis in country B is causing a strong depreciation of its currency that they do not think accords with the fundamentals of country A´s economy. Also, rules to control volatility of the nominal exchange rate should be designed and monitor carefully to avoid systematic intervention and its consequences on the consistency of monetary policy. The problem for inflation targeting emerges when what should be occasional interventions turn systematic, because it is not possible to distinguish when fundamentals and non-fundamentals factors are driving the foreign exchange market.

5. CONCLUSIONS

This paper uses a model similar to that developed by Ball (1998) to analyze how the attempt to influence directly the behavior of the nominal exchange rate affects the conduct monetary policy under a Taylor rule and inflation targeting.

In section 3, I extend Ball´s (1998) model to an open economy with a crawling exchange rate regime. In this case, the policymakers have explicit targets for the nominal exchange rate that require a certain – endogenous – level of intervention. In this model, the parameters of the optimal nominal interest rate rule depend positively on the rate of the crawling peg (Ω). This implies that setting a relatively low value of Ω to reduce inflation towards some target value, decreases the level of the interest rate for a given deviation of output and inflation with respect to their means. This, in turn, generates pressures on aggregate demand and international reserves.

In section 4, I extend Ball´s (1998) model to include systematic intervention of the monetary authority in the foreign exchange market to moderate the rate of depreciation of the nominal exchange rate. In this case, policymakers do not have nominal exchange rate targets, but in trying to control what they consider high rates of depreciation, their intervention actions become systematic. We show that intervention reduces the response of the interest rate to deviations in output and inflation from their mean values. This decreased response of the interest rate generates pressures on aggregate demand and the nominal exchange rate to depreciate, thus reducing the credibility of the anti-inflation stance of monetary policy.

The main conclusion of the paper is that a conflict exists between the implementation of a Taylor rule or inflation targeting, and systematic intervention in the foreign exchange market. Policymakers should know that such potential conflicts may arise in attempts to implement an independent monetary policy while simultaneously trying to control the behavior of the nominal exchange rate. Thus, our results formalize the notion that a monetary policy based on a Taylor rule or inflation targeting requires a fairly "clean" flexible exchange rate regime. This issue is relevant because as Calvo and Reinhart (2000) point out "fear of floating" is a generalized problem. They argue, however, that in many countries, interest rate policy replaces foreign exchange intervention as the preferred means of smoothing exchange rates.[3] Nevertheless, the temptation to intervene directly to reduce volatility may still be present. Calvo and Reinhart (2000) hold: "In the context of less-than-freely-floating exchange rates, purchases and sales of international reserves are routinely a means for smoothing exchange rate fluctuations (often, alongside interest rate policy, as discussed)" (p. 10).

Recently, some economists argue, however, that non-systematic intervention in the foreign exchange market may be compatible with inflation targeting. On this issue, central banks that have adopted inflation targeting should evaluate carefully, given the specific characteristics of their respective economies, intervention strategies that do not erode the credibility of the monetary policy regime. In particular, central banks adopting inflation targeting without a strong reputation may stifle the effectiveness of the scheme, if they embark very early in exchange rate smoothing through direct intervention in the foreign exchange market.

REFERENCES

Ball, Laurence (1997). *Efficient Rules for Monetary Policy*. National Bureau of Economic Research, Working Paper 5952.

Ball, Laurence (1998). *Policy Rules for Open Economies*. National Bureau of Economic of Economic Research, Working Paper 6760.

Bernanke, Ben; Thomas Laubach; Frederic Mishkin; Adam Posen (1999). *Inflation Targeting: Lessons from the International Experience*. Princeton University Press.

Bofinger, Peter (2001). *Monetary Policy: Goals, Institutions, Strategies and Instruments*. Oxford University Press.

Calvo, Guillermo, Carmen Reinhart (2000). *Fear of Floating*. Unpublished manuscript University of Maryland.

International Monetary Fund (2002). *International Financial Statistics*. May.

[3] In an inflation targeting regime excessive response to exchange rate fluctuations may also generate some problems. This subject, however, lies beyond the scope of this paper.

Leiderman, Leonardo; Lars E. O. Svensson (Editors) (1995). *Inflation Targets*. Center for Economic Policy Research.

Rudebusch, Glenn; Lars E.O. Svensson (1999). *Policy Rules for Inflation Targeting*. Published in Monetary Policy Rules, John B. Taylor, editor. The University of Chicago Press.

Svensson, Lars E. O. (1996). *Inflation Forecast Targeting: Implementing and Monitoring Inflation Targets*. National Bureau of Economic Research, Working Paper 5797.

Svensson, Lars E. O. (1997). *Inflation Targeting: Some Extensions*. National Bureau of Economic Research, Working Paper 5962.

Svensson, Lars E. O. (1998). *Inflation Targeting as a Monetary Policy Rule*. National Bureau of Economic Research, Working Paper 6790.

Taylor, John B. (1994). *The Inflation/Output Variability Trade-off Revisited*. Published in Goals, Guidelines, and Constraints Facing Monetary Policymakers: Proceedings of a Conference held in June 1994, Jeffrey Fuhrer, editor. Federal Reserve Bank of Boston.

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Price inflation and exchange rate pass-through in Chile

I. INTRODUCTION

This article estimates a price equation using quarterly Chilean data from 1986:1 to 2001:1. Several issues that are important for understanding and anticipating the behaviour of inflation could motivate the estimation of such an equation. For instance: i) elasticity of inflation to the output gap ii) the permanent and cyclical movements of mark-ups, iii) effects of labour productivity growth on inflation, iv) credibility and v) the size of the exchange rate pass-through.

Even though the estimation is related to all these subjects, we take a closer look at the pass-through effect, defined as the effect

Paper prepared by C. J. García and J. E. Restrepo, Senior Economists, Economic Analysis Department, Central Bank of Chile. The paper was presented at the VI Meeting of the Network of America Central Banks Researchers, organized by the Central Bank of Uruguay, in Montevideo, Uruguay, October 17-18, 2001. The views expressed in this paper are those of the authors and should not be attributed to the Central Bank of Chile. The authors would like to thank Francisco Figueiredo, Pablo García, Esteban Jadresic, Leonardo Luna, Carlos Esteban Posada, Rodrigo Valdés, for their comments and suggestions. *Corresponding author: Jorge Enrique Restrepo. Postal Address: Central Bank of Chile, Agustinas 1180, 4º piso, Santiago, Chile. E-mail: jrestrep@bcentral.cl*

of exchange-rate changes on domestic inflation, because apparently, this factor has substantially changed in recent years.[1] The exchange rate pass-through has been low recently despite the fact that Chile is a small open economy. In fact, there has been significant peso depreciation since 1997 without having a strong impact on inflation. Why is this? Is a low pass-through a new permanent characteristic of the Chilean economy? Will depreciation's impact on inflation take place as soon as demand takes off again? The answer to these questions is crucial for defining monetary policy. The international evidence shows that a very low exchange-rate pass-through has also been observed in New Zealand, Brazil and Australia, where substantial depreciation has taken place after 1997 without having a proportional effect on inflation.[2]

In general, the exchange rate affects the price of any tradable good.[3] However, the most important channel for passing depreciation to inflation is the direct short-term effect of the exchange rate on the imported part of the basket of goods that make up the CPI and the imported inputs. The larger the share of imported goods within the CPI basket, the greater the exchange rate effect on prices. In Chile, about 48% of CPI goods are considered importable. The exchange rate also directly affects the cost structure of companies using imported inputs. Thus, the greater the proportion of imported inputs making up the costs, the more depreciation will affect these companies' prices. It is important to note that we have considered in our estimation a narrower definition of prices than CPI, which we call core inflation throughout the paper.[4]

[1] We compute it here as the ratio between accumulated inflation and accumulated depreciation after a shock has hit the last variable.

[2] There has been a great amount of articles written on pass-through over the years. For a survey see Golberg and Knetter (1997). Most of them try to estimate how much exchange rate fluctuations are responsible for the behaviour of inflation. Some use CPI inflation others use producer price inflation. There is also a wide range of estimation techniques used to obtain a quantitative result. It goes from ordinary least squares (Woo, 1984), to panel data (De Gregorio and Borensztein, 1999; Goldfajn and Werlang, 2000), vector auto regression (McCarthy, 1999), cointegration analysis and error correction models (Beaumont et al. 1994; Kim, 1998; Kim, 1990), and state-space models (Kim, 1990).

[3] The exchange rate effect on prices of exportable goods is empirically less important given that some of Chile's main exports are not included in the core price index, like: fruits, or even in the CPI like: copper, fishmeal, wood pulp, salmon, methanol, etc.

[4] The price index is around 70% of CPI because we took out perishable goods and regulated services. Therefore, our estimation of the pass-through co-

Monetary policy and agents' expectations also influence the effect on prices of exchange–rate depreciation. Although in the short-term, inflation may rise due to depreciation, in the medium– and long–term, inflation should fall back to the target or range level defined by the Central Bank. Of course, an active monetary policy implies affecting aggregate demand and the exchange rate itself.[5]

Phillips curves have been estimated for Chile before (García, Herrera, and Valdés, 2001). Nonetheless, we will follow a different approach here. Instead of estimating a reduced form relation between the change in inflation and the output gap, we will estimate an equation for price inflation that considers explicitly a model of nominal price setting by imperfect competitive firms. In addition, we use the linear quadratic price adjustment cost model (LQAC) in Rotemberg (1982), where the representative firm chooses a sequence of prices for solving its intertemporal problem. As a result, inflation can be represented as an error correction equation (Euler equation), relating this variable to expected inflation as well as to the gap between the "equilibrium" and actual price level. The error correction in the price equation ensures that in the steady state, the price level is a mark-up on unit labour costs. Different versions of the core–price equation are estimated by using a single–equation error correction procedure.

Moreover, an I(2) analysis of inflation and the mark-up is undertaken. We find that the price level is best described as an I(2) process. It is worth stressing that Chile is not an exception in this regard. In fact, Roberts (2001) models inflation as a unit–root process for the USA. Johansen (1992) and Engstead and Haldrup (1999) do the same for UK. Barnerjee, Cockerell and Russell (2001) find that prices in Australia are better described as an I(2) process as well. In order to deal with I(2) processes, we incorporate inflation as an additional component of the "equilibrium" price in the Euler equation (Engsted and Haldrup, 1999 also Barnerjee, et al. 2001). On the other hand, we deal with the inflation expectation term using the limited-information approach due to McCallum (1976).

Taking the estimated price equation, the exchange-rate pass-

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efficient should be taken carefully, since some regulated services are very sensitive to changes in the exchange rate.

[5] Another factor affecting the degree of pass-through is how permanent agents perceive the shock. We will not deal with exchange-rate volatility here, though.

through is analysed by simulating an exchange rate shock, with and without an output gap since pass-through is related to economic activity, but also with and without full wage indexation. Had a negative output gap not existed after 1997, the exchange–rate effect on inflation would have been significantly higher. In addition, as a consequence of prices being positively related to wages, the simulation shows that when wage indexation is not complete –and wages grow less than past inflation– pass-through is lower in the long run. The results also show, as expected, that labour productivity reduces unit labour costs and inflation.

This article is organised as follows. The second section introduces some preliminary evidence on the recent pass-through reduction and presents a price model. The third section presents the estimation results and the pass-through simulation. Finally there are some conclusions.

II. PASS-THROUGH FROM DEPRECIATION TO INFLATION

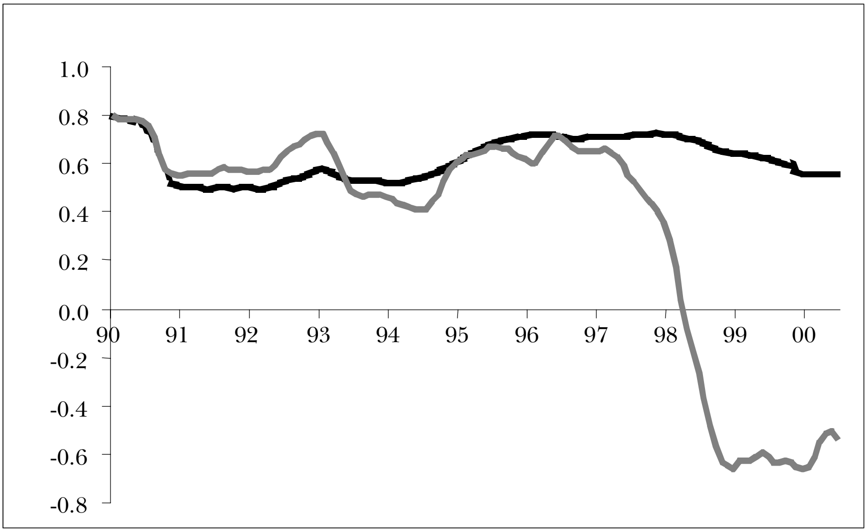
a. Stylised facts

The simplest exercise one can realise regarding the pass-through is to compute the correlation between inflation and exchange–rate depreciation. In this case, two rolling correlation statistics were computed (Figure 1). The first one (dark line) has its starting date fixed (1986:1) and the correlation coefficient is calculated adding a new observation at a time, starting in 1990. Therefore, each computation has a larger sample than the last one. Even though this coefficient is rather stable, it has some movement. It decreases at the beginning of the nineties, grows again from 1994 to 1996 and falls steadily after 1998.

The second statistic in Figure 1 (grey line) has a fully moving sample. Thus, each time the correlation coefficient is computed, both the starting and ending dates move. In this case the statistics fluctuates much more. The latter coefficient moves in a similar way to the former until to 1996. After that year it falls dramatically, to even become negative, showing an important change in the relationship between these two variables.

In addition, a rolling regression was estimated for annual inflation with exchange rate depreciation and a trend as right hand side variables using monthly data between 1986 and 2000 (Figure 2). Again the two types of rolling samples were used. The left side in Figure 2 shows the regression coefficient obtained when the

FIGURE 1. ANNUAL INFLATION AND DEPRECIATION: ROLLING CORRELATION COEFFICIENT, 1990-2000



initial date of the sample does not change. The sample used to estimate the right side in the figure has both (initial and last) dates moving. The left panel of Figure 2 shows that the coefficient started falling in 1996, earlier than in the right panel and Figure 1. As one would expect, the coefficient is less stable with both dates moving.[6]

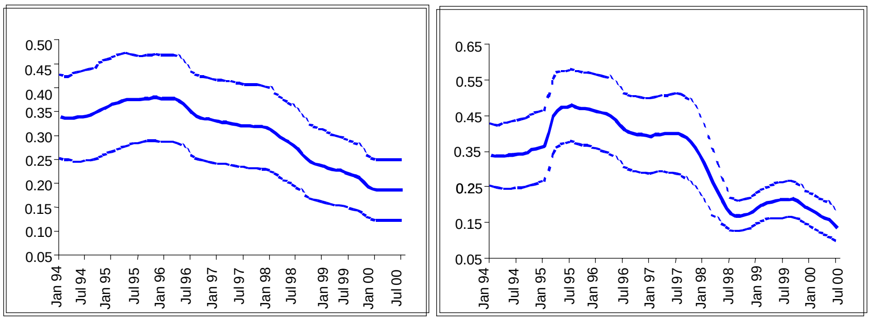
b. A price setting model

1. Price setting

In this section we derive a price model following Beaumont et al. (1994) and Layard et al. (1991). Given that prices adjust slowly, we also use the linear quadratic price adjustment model developed by Rotemberg (1982). Thus, firms should weigh the cost of changing prices against the cost of being away from the price that the firm would choose in case there were no adjustment cost (Roberts, 1995). One could think of the latter price as being the "optimal price" or the long-run equilibrium price, given that, in general, prices are sticky in the short run.

[6] As a matter of fact, it matches some stylised facts of the economy during the last decade. It is well known there was a consumption boom between 1995 and 1997 which coincides with a rebound of this coefficient.

FIGURE 2. ROLLING REGRESSION BETWEEN INFLATION AND EXCHANGE-RATE DEPRECIATION, 1994-2000



Firms are identical and get an output (*y*) by using labour (*l*) and an imported input (*z*):

$$y_i = a_1 + a_2\, l_i + (1 - a_2)\, z_i \qquad (1)$$

Each firm's demand is y_d-f, where f is the log of the number of identical firms. The demand curve faced by each firm would be:

$$y_{di} = -\eta\,(\tilde{p}_i - p) + y_d - f \qquad (2)$$

Where p_i is the firm's price, p is the price level and η is the elasticity of demand. Therefore, the price that maximises benefits in the long run is given by:

$$\tilde{p}_{di} = -\log\left[\frac{n}{n-1}\right] + MC = m + MC = m + a_1 + a_2 w_t + (1 - a_2)\, p_t^{*} \qquad (3)$$

Where the price (p_{di}) is fixed by charging a margin (*m*) over the marginal cost (*MC*). A pricing model based on a mark-up over costs would be inappropriate if it were applied to markets close to perfect competition like the ones for agricultural products (Woo, 1984).

It is assumed that firms desire a constant mark-up, *m* in the long term. However, in the short run firms could postpone price adjustments and accept deviations of their mark-up from the desired level. In doing so, firms could be motivated by both market share and the actual cost of changing prices or menu cost (Ghosh and Wolf, 2001). Demand fluctuations and anything affecting market power could have an impact on the mark-up (Barnerjee, Cockerell and Rusell, 2001; Small, 1997; Taylor, 2000). If the economy is in the midst of a recession, companies will find it difficult to transfer to their prices higher costs due to depreciation.

In addition, margins and inflation may also be either positively or negatively related because there are two opposite effects. In Taylor's words "firms in low inflation economies will appear to have less pricing power than firms in high inflation economies." (Taylor, 2000). Second, one would also expect that inflation imposes costs on firms and therefore the mark-up net of inflation is reduced with higher inflation rates (Benabou, 1992; Banerjee, Cockerell, and Russell, 2001).

Based on what was said in the last two paragraphs we assumed that the mark-up equation depends on average labour productivity, inflation and the output gap i.e. it has a cyclical component:

$$m_t = c_1 + c_2 q_t + c_3 (y_t - y_t) + c_4 \Delta p_t \tag{4}$$

Following Beaumont et al. (1994), and Banerjee et al. (2001) one can approximate equation 4 by this expression:[7]

$$\tilde{p}_{di} = (a_1 + c_1) + a_2 (w_t - q_t) + (1 - a_2) p_t^{*} + c_3 (y_t - y_t) + c_4 \Delta p_t \tag{5}$$

Where p* is equal to foreign input prices adjusted by nominal exchange rate and taxes and w_t-q_t is wages minus average labour productivity (unit labour cost). Here we are imposing $a_2 = -c_2$, which implies that income shares are independent of the level of productivity in the long run. We drop output gap from the long-run price equation (5) on the basis that it is equal to zero at the steady-state level although in the short run mark-up (P – MC) depends on economic activity.[8] Following Layard et al. (1991) p_{di} and y_{di} are aggregated so $p_{di} = p$ and $y_{di}=y$.[9]

2. Price dynamics

The structural equation for inflation is in the spirit of the new Phillips curve literature. It evolves explicitly from a setting of im-

[7] Note also that w_t can be separated in private (wpr_t) and public wages (wpu_t).

[8] The theory about the relationship between margins and the cycle is ambiguous. Some models predict pro-cyclical margins (Kreps and Scheinkman, 1983; Haskel and Small, 1995; Small, 1997). Others, in contrast, predict that they are countercyclical (Rotemberg and Saloner, 1986; Rotemberg and Woodford, 1991).

[9] The optimal price has been aggregated as in Layard et al. (1991) chapter 7, page 436.

perfect competitive firms where nominal prices are rigid or adjust slowly. In doing this, we use the (Rotemberg, 1982) LQAC model of the representative firm, which minimises the loss of charging a different price for its product from the long-run level, weighed against the cost of changing its price. This intertemporal problem is solved by choosing a sequence of p_t, the decision variable, in order to:

$$\min_{\{p_{t+i}\}} E_t \sum_{i=0}^{\infty} \beta^i \left[\theta\,(p_{t+i} - \tilde{p}_{t+i})^2 + (p_{t+i} - p_{t+i-1})^2\right] \tag{6}$$

where E_t is the expectations operator conditional on the full public information set, β is the subjective discount rate, θ is the relative cost parameter, $\tilde{p}_t$ denotes the optimal price and p_t the current or actual price. After rearrangement, the Euler equation from the minimisation problem can be written as:

$$\Delta p_{t+i} = \beta\,\Delta p^{e}_{t+i+1} - \theta\left[p_{t+i} - \tilde{p}_{t+i}\right] \tag{7}$$

Where Δp^{e}_{t+i+1} denotes expected inflation. One could think of it as being an error-correction equation relating the rate of inflation to the gap between the equilibrium and actual price levels. In order for this to be a useful theory of inflation, the optimal price level needs to be defined as in (5).

The second step is to reparameterise equation (5) for carrying out the I(2) analysis. Following Haldrup (1995) the optimal price can be parameterised as:

$$\tilde{p}_t = \gamma_1 x_{1,t-1} + \gamma_1 \Delta x_{1,t} + \gamma_2 x_{2,t-1} + \gamma_2 \Delta x_{t-1} + \gamma_2 \Delta\, x_{2,t-1} + \gamma_2 \Delta^2 x_{2,t}$$

where x_1 are the I(1) variables $\{q_t, \Delta p_t\}$ while x_2 are the I(2) ones $\{w_t\}$. Therefore we transform the optimal price:

$$\tilde{p}_t = (1-a_2)\,p^{*}_{t-1} + a_2\,(w_{t-1} - q_{t-1}) + c_4 \Delta p_{t-1} + a_2 \Delta w_{t-1} + (1-a_2)\Delta p^{*}_t + a_2(\Delta^2 w_t - \Delta q_t) + c_4 \Delta^2 p_t$$

Now we transform $\theta\left[p_t - \tilde{p}_t\right]$ to get the cointegration error correction term.

In order to do that we add and subtract Δp_{t-1}, and we also use two identities $p_t \equiv p_{t-1} + \Delta p_t$ and $\Delta p_{t-1} \equiv \phi \Delta p_{t-1} + (1-\phi)\Delta p_{t-1}$ where $\phi = \dfrac{\beta}{1+\theta}$. Thus, equation (7) can be written in acceleration form:

$$\Delta^2 p_t = k_1(\Delta p^e{}_{t+1} - \Delta p_{t-1}) + k_2(1-a_2)\Delta p_t^* + k_2 a_2(\Delta^2 w_t - \Delta q_t) + \psi\left(y_{t-1} - \bar{y}_{t-1}\right)$$
$$-k_2\left(p_{t-1} - \left[(1-a_2)p_{t-1}^* + a_2(w_{t-1} - q_{t-1}) + a_2\Delta w_{t-1} + \left(c_4 + \frac{(1-\phi)(1+\theta)}{\theta}\right)\Delta p_{t-1}\right]\right) + \varepsilon_t \quad (8)$$

Where $k_1 = \dfrac{\beta}{1+\theta(1-c_4)}$ and $k_2 = \dfrac{\theta}{1+\theta(1-c_4)}$

Equation (8) is what we refer to as the Phillips curve. This equation relates inflation to expected inflation, wage growth, output gap, and average cost. In addition, there is an error correction term, which ensures that in steady state the price level is set adding a mark-up on unit labour cost and imported-input prices (Layard et al. 1991).

Finally, it is important to notice that expected inflation matters because prices are sticky. What happens with prices next period affects current prices. Note that expectations can be rational or adaptive. When expectations are rational, we will have a price curve similar to the New Phillips curve proposed by Galí (2000) and Roberts (1995) where the inflation rate can jump. However, it is usually found empirically that inflation shows a great amount of inertia.[10] A successful stabilisation program should take this into account, in order to reduce the risk of causing a sharp fall in the rate of output growth.

III. RESULTS

We present here the estimation results. Instead of applying the two step method proposed by Engle and Granger (1987) and Haldrup (1995), we estimated the long-run relationship together with the dynamics, as in equation (8), following Harris (1995).[11] As this author puts it, when estimating a long-run equation, superconsistency ensures that it is asymptotically valid to omit the stationary I(0) terms, however the long-run relationship estimates will be biased in finite samples (see also Phillips, 1986). Therefore, Harris concludes (citing Inder, 1993) that in the case of fi-

[10] In Chile inflation is highly persistent to the extent that it is best described as being an I(1) process.

[11] Harris, R. *Cointegration Analysis in Econometric Modelling,* Chapter 4, 60-61. See also Phillips and Loretan (1991) for a comparison among several one-step (uniequational) cointegration methods used to estimate long-run economic equilibria.

nite samples, "the unrestricted dynamic model gives... precise estimates [of long-run parameters] and valid t-statistics, even in the presence of endogenous explanatory variables" (Harris op. cit., p. 60-61). At the same time it is also possible to test the null hypothesis of no cointegration by testing the significance of the error correction coefficient (see Harris op cit., for an explanation and for critical values see Banerjee, Dolado, Galbraith, and Hendry, 1993, p. 223-233).

In addition, an I(2) analysis of inflation and the mark-up is done as in Haldrup (1995). We find that the levels of prices and unit labour costs are best described as I(2) processes, this result can probably be accounted for by the existence of an I(1) inflation target during the1990s. As said above, Chile is not unique in this regard. In fact, Roberts (2001) models inflation as a unit–root process for the USA. Johansen (1992) and Engstead and Haldrup (1999) also do the same for UK. Barnerjee, Cockerell and Russell (2001) who find that prices in Australia are also better described as an I(2) process as well. An I(1) inflation process is not necessarily inconsistent with inflation targeting, given that the target was not "stationary." Furthermore, an I(1) inflation could be allowed to wander inside a target range. Since in this case there would be an inaction zone, the monetary policy reaction function would be non-linear and is called by Orphanides and Wieland (2000) "inflation zone targeting." Such a stochastic process is known in the (continuous time) literature as a Brownian motion with barriers. Although this kind of process is expected to settle down to a stationary one in the long run, it is not yet the case in the sample we are considering here.[12]

a. Data, unit roots

The quarterly series (in logs) used in the model from 1986:1 to 2001:1 are the following:

p: price level, it excludes perishable food as well as gas, fuels and regulated services

q: average labour productivity

wpr: Nominal private wage

[12] Another alternative would be treating inflation as a stationary process but using a calibrated autoregressive coefficient very close to 1. In this case, inflation would still be very persistent although being a stationary process. However, it would be a calibration experiment instead of being an econometric estimation.

wpu: Nominal public wage

w : Nominal wage

e : Nominal exchange rate

p^*: Foreign prices

$y-\bar{y}$: The difference between actual output and its Hodrick-Prescott trend[13]

t : ln(1+VAT)

ta : ln(1+Tariff)

IT : Inflation target

Δ and Δ^2 : first and second difference, respectively.

We begin the empirical section testing the variables used in the estimations for the existence of unit roots. Table 1 indicates that price level is I(2). This confirms that the price equation can be estimated in acceleration form.

TABLE 1. UNIT ROOT TEST

Variables	*Level*	*First Difference*	*Second Difference*
Price	–	-3.04 (-3.49) 5	-4.41 (-2.92) 6
Private Wage	–	-3.07 (-3.49) 0	-4.47 (-2.92) 6
Public Wage	–	-3.46 (-3.49) 5	-4.48 (-2.92) 6
Foreign Price	-2.42 (-2.92) 6	-4.01 (-3.49) 5	–
Average Labour Productivity	-2.36 (-3.49) 1	-5.20 (-3.49) 4	–
Nominal Exch. Rate	-2.09 (-3.49) 1	-6.41 (-2.92) 0	–
Private Unit Labour Cost	–	-5.01 (-3.49) 0	-4.5 (-2.92) 6
Output Gap	-3.49 (-1.95) 3	–	–
Critical value 5% no constant	-1.95		
Critical value 5% with constant	-2.92		
Critical value 5% const. & trend	-3.49		

NOTES*:* Critical values in parenthesis. The number of lags used in the test is on the right side of every column.

In general one can say that Chilean inflation deviates from any given mean in the period here considered. Moreover, Chilean in-

[13] Potential output was computed with a Hodrick-Prescott filter but, given the problems this filter has with final data points, it was adjusted by building an autorregresive forecast from 1998 to 2001:1.

TABLE 2. PRICE EQUATION (DEPENDENT VARIABLE: $\Delta^2 P_T$), SAMPLE 1986.3-2001.3

Variables[a]	*Model 1*	*Variables*	*Model 2*
Constant	0.59 (2.75)	Constant	0.59 (3.19)
$\Delta p^e_{t+1}\text{-}\Delta p_{t-1}$[b]	0.45 (2.83)	$IT_{t+1}\text{-}\Delta p_{t-1}$	0.25 (1.62)
$[(y_{t-1}-\bar{y}_{t-1})+(y_{t-4}-\bar{y}_{t-4})]/2$	0.16 (3.0)	$[(y_{t-1}-\bar{y}_{t-1})+(y_{t-4}-\bar{y}_{t-4})]/2$	0.20 (4.0)
p_{t-1}	-0.23 (-4.8)	p_{t-1}	-0.23 (-4.6)
$wpr_{t-1}\text{-}q_{t-1}$	0.16 (4.0)	$wpr_{t-1}\text{-}q_{t-1}$	0.16 (4.21)
$wpu_{t-1}+t_{t-1}$	0.044 (3.5)	$Wpu_{t-1}+t_{t-1}$	0.04 (3.5)
$p^*_{t-1}+e_{t-1}$ $+ta_{t-1}+t_{t-1}$	0.24-0.16- 0.044=0.03	$p^*_{t-1}+e_{t-1}$ $+ta_{t-1}+t_{t-1}$	0.22-0.16 -0.05=0.03
Δp_{t-1}	-0.44 (-2.4)	Δp_{t-1}	-0.54 (-2.6)
$\Delta^2 wpr_t\text{-}\Delta q_t$	0.13 (2.6)	$\Delta^2 wpr_t\text{-}\Delta q_t$	0.15 (2.7)
$\Delta e_t+\Delta p^*_t$ $+\Delta t_t+\Delta ta_t$	0.06 (2.9)	$\Delta e_t+\Delta p^*_t$ $+\Delta t_t+\Delta ta_t$	0.05 (2.3)
Seas(2)	0.01 (3.8)	Seas(2)	0.006 (3.15)
Seas(3)	-0.003 (-1.1)	Seas(3)	-0.006 (-2.46)
Adjusted R^2	0.77		0.76
DW	2.1		2.2
ARCH(4)[c]	1.1 (37%)		0.48 (75%)
LM(4)[3]	0.72 (58%)		0.50 (73%)
Jarque–Bera[c]	0.65 (72%)		0.40 (82%)
Mean Absolute Error[d]	0.05 & 0.0015		0.05 & 0.008
Errors Unit Root test (1% critical value in brakets)	-6.29 (-2.6)		-6.06 (-2.6)

NOTES: [a] p_t is core inflation and each variable is in logs. [b] Δp^e_{t+1} is estimated by instrumental variables. We use as IV contemporaneous values Δp^*_t, Δp_{t-1}, Δq_t, Δwpu_t, Δwpr_t, Δe_t, $(y_t-\bar{y}_t)$, IT_t, $\Delta(oil\ price)_t$, contemporaneous unemployment rate, and Seasonal variables. [c] Probabilities are reported in brackets. [d] Out-of- Sample Forecast for 1998.1-2001 and also for 1999.1-2001.

flation has traditionally been very persistent due to generalised indexation. On the other hand, variables such as output gap and the nominal exchange rate are I(0) and I(1) respectively. Regarding cointegration we follow Banerjee, Dolado, Galbraith, and Hendry's (1993) approach which states that a test of the null hypothesis Ho: $c_2 = 0$ (the error–correction coefficient in equation 8) based on the t-statistic $t_k = 0$ is a valid test for cointegration. If the variables are not cointegrated, this coefficient should be zero. They also include computed critical values for this test in Table 7.6 in their book. Both equations pass the test since the critical t value is 4.06 at 1% of significance (Table 2).

b. Price equation

As stated in equation (8), price acceleration was run on wage, average labour productivity, output gap, lagged prices, foreign prices and several difference terms.[14] We have estimated two versions of equation 8.

- Model 1

In this estimation we imposed β6 = -β7, which implies that we can introduce unit labour costs instead of having private wages and labour productivity separated. Cost homogeneity (the various costs add up to prices) was also imposed: -β4= β5 + β6 + β7 + β8 + β9. In consequence an increase in all nominal inputs generates a proportional increase in prices.

$$\Delta^2 p_t = \beta_1 + \beta_2 (\Delta p^e_{t+1} \text{-} \Delta p_{t-1}) + \beta_3([(y_{t-1} - \bar{y}_{t-1}) + (y_{t-4} - \bar{y}_{t-4})]/2) + \beta_4 p_{t-1}$$
$$+ \beta_5(wpr_{t-1} - q_{t-1}) + \beta_6 (wpu_{t-1} + taxes) + (-\beta_4 - \beta_5 - \beta_6)(e_{t-1} + p^*_{t-1} + ta_{t-1} + t_{t-1})$$
$$\beta_7 \Delta p_{t-1} + \beta_8 (\Delta^2 wpr_t - \Delta q_t) + \beta_9(\Delta e_{t+} \Delta p^*_t + \Delta t_t + \Delta ta_t) + \text{Seasonal variables}$$

- Model 2

In this model we used the inflation target as a proxy for expected inflation:

[14] Based on De Gregorio and Borensztein (1999) and Goldfajn and Werlang (2000), we also tried including the real exchange rate misalignment multiplied by the depreciation rate, but the coefficient associated with this variable was not statistically significant. We also tried including the oil price in these regressions, to take into account short-run shocks to the system, but it was not significant.

$$\Delta^2 p_t = \beta_1 + \beta_2 (IT_{t+1} \text{-} \Delta p_{t-1}) + \beta_3([(y_{t-1} - \bar{y}_{t-1}) + (y_{t-4} - \bar{y}_{t-4})]/2) + \beta + \beta_4 p_{t-1}$$
$$+ \beta_5(wpr_{t-1} - q_{t-1}) + \beta_6 (wpu_{t-1} + taxes) + (-\beta_4 - \beta_5 - \beta_6)(e_{t-1} + p^*_{t-1} + ta_{t-1} + t_{t-1})$$
$$\beta_7 \Delta p_{t-1} + \beta_8 (\Delta^2 wpr_t - \Delta q_t) + \beta_9(\Delta e_{t+} \Delta p^*_t + \Delta t_t + \Delta ta_t) + \textit{Seasonal variables}$$

Regarding the possible endogeneity of the exchange rate in these single equation models, we can argue that it does not affect the long-run coefficient of the exchange rate in the cointegrating vector. This kind of bias could only affect the coefficient of the contemporaneous exchange rate dynamics.

The results of the estimation of equation (8) are presented in Table 2. The parameters of these econometric estimations have the expected signs and the restrictions of the model hold. The coefficient on output gap $[(y_{t-1} - \bar{y}_{t-1}) + (y_{t-4} - \bar{y}_{t-4})]/2$ is positive, indicating that the direct impact of a 10% output gap will be a 1.35% acceleration of the inflation rate.

The results also show, as expected, that labour productivity reduces unit labour costs and inflation. On the other hand, expected inflation acceleration $\Delta p^e_{t+1} - \Delta p_{t-1}$ is significant, confirming that expectations matter in determining inflation. Wages and foreign prices also accelerate inflation. Finally, lagged inflation (Δp_{t-1}) is also significant.[15]

Table 2 also shows the various diagnostic residual tests indicating that the models have the desired properties for OLS estimation.[16] Multivariate tests are satisfactory as seen in the lower part of the table. In general, the econometric fit is satisfactory with high R-squares and highly significant variables. Also, the results presented in Table 2 provide evidence of the existence of I(2) data trends and cointegration because the parameter of Δp_{t-1} is significant, according con Banerjee, Dolado, Galbraith and Hendry (1993) critical values, and the error terms are stationary.

We tested the two restrictions of model 1 using an unrestricted version of it. First, we tested the hypothesis of the coefficient on private wages being equal to the one on labour productivity, though with opposite signs. If this is the case we can include unit labour cost (w-q) as a variable in the model.

As shown in Table 3, the Wald test indicates that we fail to re-

[15] We are not able to distinguish its long-run effect on mark-ups from its contribution to the short-run dynamics.

[16] Standard errors were obtained with the Newey-West heteroscedasticity and autocorrelation consistent procedure.

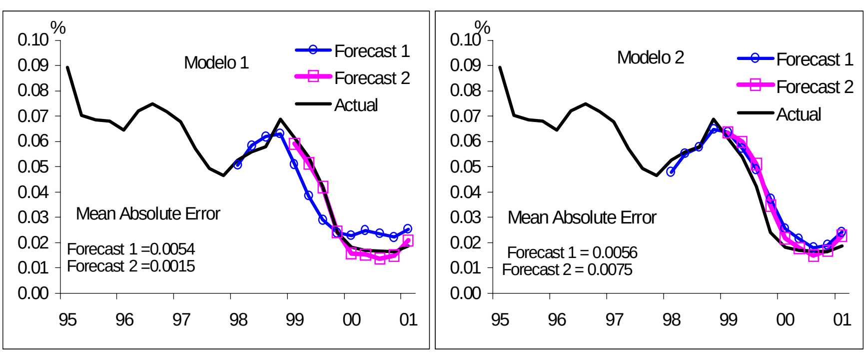
ject the null hypothesis of different coefficients at 90% of significance. Second, we tested in Model 1 the hypothesis that the various costs add up to prices (cost homogeneity). We also fail to reject this null hypothesis at a 61% of significance (Table 3). As a result we imposed both restrictions in Model 1.

TABLE 3. RESTRICTIONS TESTS ON MODEL 1

Wald Test Hypothesis	*Model 1*
Unit Labour Cost	90%
Linear Homogeneity and Unit Labour Cost	61%

In order to evaluate the models' forecasting ability, we estimated them until 1997:4 and generated out-of-sample inflation forecast from 1998:1 up to 2001:1 (Figure 3). Then, we estimated them again until 1998:4 and generated out-of-sample inflation forecast from 1999:1 up to 2001:1 (Figure 3). We find that both models follow actual inflation rather well. However, both overestimate inflation at the end of 2000. The mean absolute error of both forecasts is found in Table 2 and Figure 3.

FIGURE 3. OUT-OF-SAMPLE FORECAST OF ANNUAL INFLATION, 1999.1-2001.1



c. Pass-through[17]

Finally, we analyse the implications of our estimations on exchange rate pass-through by simulating an unexpected perma-

[17] Again, the reader should take into account that these results do not consider the exchange rate effect on regulated services and gas. Therefore exchange-rate pass-through is actually somewhat higher.

nent 10% shock to the nominal exchange rate.[18] In order to do so, besides using the estimated equation (first column in Table 2), we assume fully indexed wages. Thus, we solve the model (not estimate it) simultaneously, by also including the following equation: $w_t = w_{t-1} + \Delta p_{t-1}$, for both private and public sectors, respectively. After we introduce a shock to the nominal exchange rate, we compute the pass-through effect using the simulated paths followed by both the nominal exchange rate and domestic prices. It is worth noting that the exercise does not consider any monetary policy action.

FIGURE 4. EXCHANGE RATE PASS-THROUGH SIMULATION

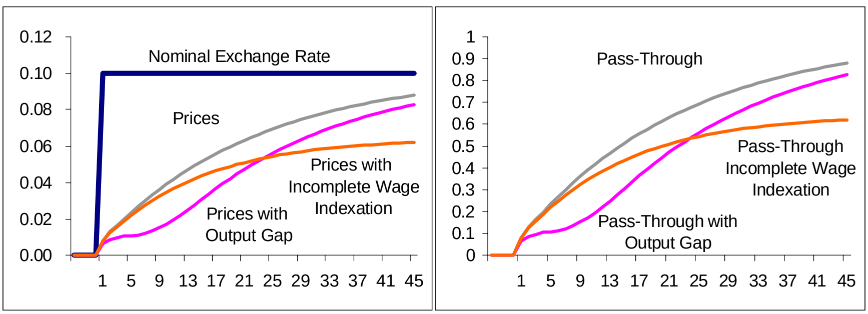


Figure 4 shows the exchange–rate pass-through effect when the nominal exchange rate unexpectedly increases. It indicates that a 100% rise in the exchange rate produces an accumulated impact on prices of around 33% in the first two years (8 quarters) which is considered the relevant policy horizon.[19] This means that a nominal devaluation will not be proportionally translated into

[18] Pass-through will be smaller if the exchange rate shock is not a permanent one. It could also be smaller, at least in the short run, if agents believe the shock is only temporary.

[19] International evidence indicates that the pass-through of the exchange rate to prices is lower in developed countries than in Latin America and Asia. In one panel estimate with 71 countries, Goldfjan and Werlang (2000) found a depreciation-to-inflation pass-through coefficient of 0.73 at the end of 12 months. When the sample was sorted between OECD members (Organization for Economic Cooperation and Development) and emerging economies, at the end of 12 months, pass-through coefficients of 0.6 and 0.91 respectively appeared. When this sample was sorted by regions, the 18-month coefficient for Europe was 0.46, while in America it was 1.24. Finally, as a result of an exercise based on their estimates, the authors found a bias toward predicting higher inflation than actually observed in several well-known cases of large depreciation.

prices in the short run, affecting the real exchange rate. After that, pass-through increases approximating to 100% in the very long run.

We also realised an exercise imposing the limitation that private wages do not have full indexation: $w_t = w_{t-1} + 0.9*\Delta p_{t-1}$. Figure 4 shows that in this case, pass-through is much smaller in the long run. Nonetheless, this effect is not large in the first two years. Of course, this scenario assumes that private wages will permanently bear the cost of a higher nominal exchange rate, which is not realistic.

Evidence suggests that there is a pass-through decrease when the output is below potential (recession) because a negative output gap tends to compensate the inflationary effect of depreciation by reducing margins. This is what usually happens when a currency depreciation is the result of a negative terms of trade shock with negative effects on output (Mishkin, 2001). Figure 4 shows that the exchange-rate effect is much smaller when there is an exogenous 2% negative output gap, which fades linearly in 3 years. In this case, a 100% increase in the exchange rate will translate into a 13,5% price increase i.e. less than half of what it was before. Hence, a fraction of the depreciation is not passed on to consumers. In the long run, as output gap disappears, pass-through approaches 100%.[20]

IV. CONCLUSIONS

A price equation based on a model of imperfect competition was estimated and used to generate out-of-sample forecasts for core inflation. The parameters of these econometric estimations have the expected signs and the restrictions of the model hold. It was empirically found, as expected, that labour productivity reduces unit labour costs and inflation. On the other hand, expected inflation acceleration $\Delta p^e_{t+1}-\Delta p_{t-1}$ is significant, confirming that expectations matter for determining inflation. Wages and foreign prices are also positively related to inflation. The coefficient on output gap $(y_{t-1}-\bar{y}_{t-1})$ is positive, indicating that the direct impact of a 10% output gap will be a 1.6% acceleration of the inflation rate.

We analyse the implications of our estimations on exchange

[20] Since 1998 the output gap in Chile has probably been higher than what we assumed for this simulation.

rate pass-through by simulating an unexpected permanent 10% shock to the nominal exchange rate. A nominal devaluation has real effects that disappear in the long run. In the case of incomplete wage indexation, pass-through is much smaller in the long run. Nonetheless, this effect is not large in the first two years.

The simulation also shows that a negative output gap tends to compensate the inflationary effect of depreciation since exchange–rate pass-through depends positively on economic activity. In this case, a fraction of the depreciation is not passed on to prices in the short run, explaining why pass-through has been so low in recent years. However, as time goes on and the output gap disappears, pass-through approaches 100%.

If the recent peso depreciation in Chile is permanent, one can conclude from the results, that pass-through will increase as soon as aggregate demand starts recuperating. Nevertheless, this will only be the case if monetary authorities take no action.

REFERENCES

Banerjee, A., Cockerell, L. and Russell, B. (2001), "An I(2) Analysis of Inflation and the Markup", *Journal of Applied Econometrics,* 16(3), 221-240.

Banerjee, A., Dolado, J. Galbraith, J. and Hendry, D. (1993), *Co-Integration, Error–Correction, and the Econometric Analysis of Non-Stationary Data*. Oxford University Press, London.

Beaumont, C. Cassino, V. and Mayes, D. (1994), *Approaches to Modelling Prices at the Reserve Bank of New Zealand,* Reserve Bank of New Zealand, Discussion Paper, No. 3.

Benabou, R. (1992), "Inflation and Markups: Theories and Evidence from the Retail Trade Sector", *European Economic Review,* 36, 566-574.

De Gregrorio, J. and Borensztein, E. (1999), *Devaluation and Inflation after Currency Crises*, Mimeo, International Monetary Fund Research Department.

Engsted, T. and Haldrup, N. (1999), "Estimating the LQAC Model with I(2) variables", *Journal of Applied Econometrics,* 14, 155-170.

García, P., Herrera, L. and Valdés, R. (2001), *New Frontiers for Monetary Policy in Chile*, Working Papers Series No. 125, Central Bank of Chile.

Goldberg, P., and Knetter, M. (1997), "Goods Prices and Exchange Rates: What Have We Learned?", *Journal of Economic Literature,* 35(3), 1243-1272.

Goldfajn, I., and Werlang, S. (2000), *The Pass-through from Depreciation to Inflation: A Panel Study*, Working Paper of the Central Bank of Brazil, July.

Haldrup, N. (1994), "The Asymptotics of Single-Equation Cointegration Regressions with I(1) and I(2) Variables", *Journal of Econometrics,* 63, 153-81.

Haldrup, N. (1995), "An Econometric Analysis of I(2) Variables", *Journal of Economic Surveys,* 12(5), 595-650.

Harris, R. (1995), *Using Cointegration Analysis in Econometric Modelling.* Prentice Hall, London.

Haskel, J., Martin, C. and Small, I. (1995), "Price, Marginal cost and the Business Cycle", *Oxford Bulletin of Economics and Statistics,* 57(1), 25-39.

Inder, B. (1993), "Estimating Long-run Relationships in Economics: a Comparison of Different Approaches", *Journal of Econometrics,* 57, 53-68.

Johansen, S. (1992), "Testing Weak Exogeneity and the Order of Cointegration in UK Money Demand Data", *Journal of Policy Modelling,* 14, 313-34.

Kim, K. (1998), "US Inflation and the Dollar Exchange Rate a Vector Error Correction Model", *Applied Economics,* 30, 613-619.

Kim, Y. (1990), "Exchange Rates and Import Prices in the United States: A Varying-Parameter Estimation of Exchange-Rate Pass-Through", *Journal of Business & Economic Statistics,* 8(3), 305-15.

Kreps, D. and Scheinkman, J. (1983), "Quantity Pre-Commitment and Bertrand Competition Yield Cournot Outcomes", *Rand Journal of Economics,* 14, 326-37.

Layard, R, Nickell S. and Jackman R. (1991), *Unemployment. Macroeconomic Performance and the labour market.* Oxford University Press; Spanish version translated and Edited by the Ministry of Labour and Social Security of Spain.

McCallum, B. (1976), "Rational Expectations and the Natural Rate Hypothesis: Some Consistent Estimates", *Econometrica,* 44, 43-52.

McCarthy, J. (1999), *Pass-Through of Exchange Rates and Import Prices to Domestic Inflation in Some Industrialised Economies*, Bank of International Settlements, Working Paper No. 79, November.

Mishkin, F. (2001), *The Transmission Mechanism and the Role of Asset*

Prices in Monetary Policy, National Bureau of Economic Research, Working Paper No. 8617.

Orphanides, A. and Wieland, V. (2000), "Inflation Zone Targeting", *European Economic Review,* 44(7), 1351-1388.

Phillips, P. (1986), "Understanding Spurious regressions in econometrics", *Journal of Econometrics,* 33, 335-46.

Phillips, P. and Loretan, M. (1991), "Estimating Long-run Economic Equilibria", *Review of Economic Studies,* 58, 407-36.

Roberts, J. (1995), "New Keynesian Economics and the Phillips Curve", *Journal of Money, Credit, and Banking,* 27(4), 974-84.

Roberts, J. (2001), *How Well Does the New Keynesian Sticky-Price Model Fit the Data?*, Finance and Economic Discussion Papers No. 13, Board of Governors of the Federal Reserve System, February.

Rotemberg, J. (1982), "Sticky Prices in the United States", *Journal of Political Economy*, 60, 1187-1211.

Rotemberg, J. and Saloner, G. (1986), "A Supergame Theoretic Model of Price Wars During Booms", *American Economic Review* 76, 390-407.

Rotemberg, J. and Woodford, M. (1991), "Mark-ups and the Business Cycle", *National Bureau of Economic Research Macroeconomics Annual*, 6, 63-128.

Russel, B., Evans, J. and Preston, B. (1997), *The impact of Inflation and Uncertainty on the Optimum Price Set by Firms*, Department of Economic Studies Discussion Papers, No. 84, University of Dundee.

Small, I. (1997), *The Cyclicality of Mark-ups and Profit Margins: Some Evidence for Manufacturing Services*, Working Paper No. 72, Bank of England.

Taylor, J. (2000), "Low Inflation, Pass-Through, and the Pricing Power of Firms", *European Economic Review,* 44(7), 1389-408.

Woo, W. (1984), "Exchange Rates and the Prices of Non-food, Non-fuel Products", *Brookings Papers on Economic Activity,* 2, 511-30.

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