Revista de Economía Aplicada EA *Número 58 (vol. XX), 2012, págs. 49 a 68*

A MULTIMARKET APPROACH FOR ESTIMATING A NEW KEYNESIAN PHILLIPS CURVE*

J.D. TENA
Universidad Carlos III y Università di Sassari

JORGE DRESDNER
IVÁN ARAYA
Universidad de Concepción

We propose a new approach for estimating a "hybrid" New Keynesian Phillips Curve (NKPC) that includes demand pressures coming from disequilibrium relations in three different markets: (1) monetary and financial, (2) international, and (3) labour. Econometric tests based on Chilean data indicate that this specification is superior to the traditional NKPC, which includes a single variable to account for demand pressures. Moreover, considering cointegration relationships in the monetary and labour markets is particularly useful for forecasting the dynamics of inflation compared to an alternative specification based on a standard measure of output fluctuation.

Key words: New Keynesian Phillips Curve, cointegration, monetary policy.

JEL Classification: E3, C3.

In this paper, we propose a new approach for estimating the New Keynesian Phillips Curve (NKPC). We modify the traditional aggregated NKPC to include excess demand pressures arising from three different markets. For the empirical application, we use data from the Chilean economy. Our results indicate that inflation is explained by pressures coming from all the markets, but that shocks in the monetary and labour markets and short-run deviations of output and inflation are relatively more important than other shocks when explaining inflation. We conclude that the proposed specification is superior to the traditional NKPC specification that includes excess demand pressures from a single market.

Based on the New Keynesian theoretical framework, we estimate a menu of models for the New Keynesian Phillips Curve (NKPC henceforth) that considers forward-looking, optimization behaviour by firms and households in the context

(*) We appreciate the comments by two anonymous referees, Klaus Schmidt-Hebbel and participants at the 2007 Encuentro de la Sociedad de Economía de Chile. The usual qualifier applies.

of price rigidities, as well as hybrid models that include backward-looking behaviour. These have been previously evaluated with data for both industrial countries like the U.S. and the Euro Zone and middle income countries. Gali and Gertler (1999), Gali *et al*. (2001), and Fanelli (2008) found evidence supporting the NKPC and the relevance of forward-looking behaviour. In contrast, Rudd and Whelan (2005) cast doubts on Gali and Gertler's empirical results. Indeed, they found evidence that the NKPC is consistent with a backward-looking Phillips curve based on the hybrid version approach. Cespedes *et al*. (2005) found a similar result in their NKPC estimation for Chile.

One of the limitations of the empirical model in the literature is the narrow approach used to treat the excess demand variable in the NKPC. Results in Hendry (2001) suggest that there is no "single-cause" explanation for inflation in the UK because this variable responds to excess demand from all sectors in the economy. In spite of this result, empirical estimations of the NKPC in the previous literature only considered demand pressures coming from, at most, two (internal and external) markets [see Gali and Gertler (1999), Gali *et al*. (2001), and Fanelli (2008), among others].

This research proposes a new approach for estimating a "hybrid" New Keynesian Phillips Curve that considers a large number of different variables separately in order to measure demand pressures in different markets. This methodology also considers a NKPC that includes long and short-run inflation dynamics. We follow a two-step procedure. In the first step, potential cointegration relationships are analyzed separately using three different vector equilibrium models in which inflation is explained by (1) the effect of monetary and financial variables, (2) information on international markets, and (3) production costs. In doing this, we follow the principle of 'specific to general' in the choice of variables, albeit 'general to specific' in the choice of the statistical model [see Juselius and MacDonald (2004), Juselius (1992), and Garrat *et al*. (2000)]. VAR models are very powerful for detailed analyses of small systems, but almost unmanageable in large systems. Moreover, the cointegration property is invariant to changes in the information set. Thus, any cointegration result found for a given set of variables can also be found in an extended analysis. Once the cointegration relationships are identified in the three markets, we use this information in the second step to obtain the structural parameters of the NKPC. In this model, inflation can be affected by shocks in the cointegration relations in the three markets as well as by the short-run dynamics induced by a group of fundamental variables. Econometric tests prove that this specification is superior to other NKPC curves in the literature such as the traditional NKPC, which uses a single variable to account for demand pressures, and the model proposed by Fanelli (2008). The empirical application of this procedure is based on data from the Chilean economy.

The next section discusses the main problems in extending the traditional NKPC to a model that includes demand pressures from different markets that affect long-run and short-run inflation dynamics. Section 2 analyzes separately the presence of cointegration relationships in the money, international, and labour markets. These cointegration relationships are used in Section 3 to specify and estimate an NKPC that reacts to disequilibria in the different markets and compares it to other more conventional NKPC in the literature. Some concluding remarks follow in Section 4.

1. Inflation and multiple market pressures

The "hybrid" New Keynesian Phillips Curve (NKPC) expresses the inflation rate as a function of the expected rate of inflation, lagged inflation, and a set of control variables. Galí *et al*. (2001) propose the following functional form:

$$\pi_t = \theta E_t \pi_{t+1} + \delta \pi_{t-1} + \lambda' x_t + a_t \qquad [1]$$

where π_t is the inflation rate at time t; $E_t\ \pi_{t+1}$ is the expected value at time t of the inflation rate at time $t+1$; x_t is a (nx1) vector or exogenous explanatory variables; a_t is an error term; and θ, δ, and λ are structural parameters with λ as a (nx1) vector.

The micro-foundation of this specification can be derived from a monopolistic competitive framework in which firms face some constraints on price adjustment [see, for example, Gali and Gertler (1999) and Eichenbaum and Fisher (2007), among others]. It is typically assumed that there is a fraction of "forward looking" firms that set prices optimally and a remaining fraction of "backward looking" firms that set prices based on a rule of thumb derived from past information. In turn, in order to aggregate the individual price decisions by "forward looking" firms, it is common to employ an assumption in Calvo (1983). The idea is that, at any given period, each firm adjusts its price in relation to a given probability and, hence, keeps the price unchanged with respect to probability $(1 - w)$. Under these assumptions, optimal price decisions by these firms can be defined as a mark-up over a weighted average of expected future nominal cost. On the other hand, "backward looking" firms set their price as a function of past information on aggregate price behavior. The NKPC that specifies inflation as a linear function of past inflation, expected future inflation and real marginal costs, can be obtained by expressing the individual price decisions taken by forward and backward looking firms relative to the aggregate price index and by aggregating these variables. In order to express the NKPC as a function of the output gap, instead of real marginal costs, it can be assumed that real marginal costs can be expressed as a linear function of the output gap [see Gali and Gertler (1999)].

Generally, x_t is a single variable that indicates demand pressure, such as the output gap or the unemployment rate. Only some small open economy versions of the NKPC include a second variable for x_t to account for external demand pressure [Peturson (1998), Batini *et al*. (2005)].

Here, as the main contribution of this paper, we separately consider a large number of different variables that indicate demand pressures in different markets. Let us split x_t into two groups of variables, $x_t' = (x_{l,t}', x_{c,t}')$, where $x_{l,t}$ is a n_1xl vector of non-stationary variables (or a combination of variables) that work in an equilibrium relationship with π_t; and $\pi_{c,t}$ is a n_2xl vector of stationary variables that can potentially affect short-run inflation dynamics, $n_1 + n_2 = n$.

Using the last definition, equation [1] can be expressed as:

$$\pi_t = \theta E_t \pi_{t+1} + \delta \pi_{t-1} + \lambda_1{}' x_{l,t} + \lambda_2{}' x_{c,t} + a_t \qquad [2]$$

where λ_1 and λ_2 are vectors of parameters with dimensions (n_1x1) and (n_2x1), respectively.

The one-step-ahead inflation forecast implied by this equation is:

$$E_t \Delta \pi_{t+1} = \left(\frac{1-\theta-\delta}{\theta}\right) \sum_{i=1}^{n_1} \left(\frac{\pi_t}{n_1} - \beta^i x_{l,t}^i\right) + \left(\frac{\delta}{\theta}\right) \Delta \pi_t - \left(\frac{\lambda_2{}'}{\theta}\right) x_{c,t} \qquad [3]$$

where $\beta^i = \dfrac{\lambda_l^i}{(1-\theta-\delta)}$, Δ is a difference operator, and $x_{l,t}^i$ and λ_l^i are respectively the ith elements in vectors $x_{l,t}$ and λ_l. Also, note that to obtain expression [3] we have used the fact that $\pi_t \equiv \dfrac{n_1 \pi_t}{n_1}$ [1].

An important issue to be explained about the last expression is that the term n_1 is only a scale factor in the cointegration relationship $\dfrac{\pi_t}{n_1} - \beta^i x^i$. It can be expressed as $\pi_t - \beta^{*i} x^i$ with $\beta^{*i} = n_1 \beta^i$. To avoid confusion, we will refer to the cointegration relationship using the more simple notation $\pi_t - \beta^{*i} x^i$.

Equation [3] is very similar to the one proposed by case 2 in Fanelli's specification (2006)[2]. However, an important difference in this approach is that, in the structural Phillips curve, we jointly consider the set of variables affecting the short and long-run inflation dynamics instead of studying them as two alternative cases. A second difference lies in the fact that, in our model, each of the $x_{l,t}^i$ variables corresponds to demand pressures coming from equilibrium relationships in different markets. This last issue will be more specifically outlined in the next section.

It is straightforward to obtain the parameters in [3] from a reduced form specification similar to this:

$$\Delta \pi_t = c + \sum_{i=1}^{n_1} \alpha_i \left(\pi_{t-1} - \gamma^i x_{l,t-1}^i\right) + \phi \Delta \pi_{t-1} - \rho'^{x_{c,t-1}} + u_t \qquad [4]$$

where ρ is a n_2x1 vector of all parameters and c, α_i, γ^i, and φ are scalars. Then, structural parameters can be obtained by matching expressions [3] and [4]. It turns out that:

(1) Expression [3] imposes the normality assumption by dividing parameters by θ, Gali and Gertler (1999). Although these authors show that this assumption may matter in some cases, we checked that our estimation is not affected by the normality assumption.

(2) This case can only be found in the working paper version but not in Fanelli (2008).

$$\alpha_i = \left(\frac{1-\theta-\delta}{\theta}\right) \qquad [5]$$

$$\phi = \frac{\delta}{\theta} \qquad [6]$$

$$\rho' = \frac{\lambda_2{}'}{\theta} \qquad [7]$$

In the following section, we specify and estimate a reduced form model similar to [4] that includes information about cointegration relationships in three markets: labour, money, and external markets. Then, restrictions [5], [6] and [7] can be imposed on the reduced form equation in order to obtain the parameters of a hybrid NKPC that contains information about demand pressures in several markets.

2. Cointegration analysis in three markets

We follow a two-step procedure in order to specify and estimate a Phillips curve for the Chilean economy. The first step identifies long-run relations in three different markets. Then, in the second step, an inflation equation is estimated including all the information about the equilibrium relationships found in the three markets considered. This section focuses on the first step, in which potential cointegration relationships are separately analyzed in three different vector equilibrium correction (VeqC) models in which Chilean inflation is explained by (1) the effect of monetary and financial variables, (2) information on international markets, and (3) production costs.

We consider the following quarterly series[3] (Table 1).

Table 1: Cointegration models

	Variable	Definition
		Asset market model
i	Short-run nominal interest rate	Banking system average buyer interest rate for 30-89-day deposits
π	Inflation rate	First difference of the log of the consumer price index
y	Real income	Gross Domestic Product in constant prices
Δm	Rate of change of nominal money	First difference of the log of M2 monetary aggregate

(3) A more detailed description of the different series and the sources from which these series were obtained can be found in the appendix.

Table 1: Cointegration models (continuation)

	Variable	Definition
		Foreign market model
i^*	Foreign interest rate	United States FEDs nominal interest rates
	Foreign inflation rate	First difference of the log of the consumer price index in the United States
i	Short-run nominal interest rate	Banking system average buyer interest rate for 30-89-day deposits
π	Inflation rate	First difference of the log of the consumer price index
e	Real effective exchange rate index - Chile index	Exchange rate - real effective exchange Rate index - Chile index base: 2000 = 1.00, based on relative consumer prices, period average
		Labour market
Pr	Labour productivity	Average ratio of real GDP over number of employed workers
w	Real wages	Ratio of nominal wage index over the consumer price index
π	Inflation rate	First difference of the log of the consumer price index
u	Unemployment rate	Ratio of unemployed workers over labour force

Source: Own estimation.

Our analysis covers the period 1987:Q1-2009:Q4. Different cointegration relationships are investigated for the different groups of variables. The first group refers to the money and financial market and includes i_t, π_t, y_t and Δm_t. The variables in the second group are i^*_t, π^*_t, i_t, π_t and e_t. These variables account for the influence of the foreign market on Chilean inflation. We consider a last set of variables (Pr_t, w_t, u_t and π_t) to measure inflationary pressures that come from the labour market.

The augmented Dickey-Fuller test applied to each of the individual series is reported in Table 2. From this table, we can conclude that, at 5%, all the variables considered in the cointegration analysis are either stationary or I(1). Variable w_t is the only exception to this general result. However, the reason for the no rejection of the null hypothesis of nonstationarity in this case is the presence of a strong outlier in 2006:Q4. Once we include a dummy variable to take this event into account, the value of its DF statistics is -4.91, which leads to the rejection of the nonstationarity hypothesis at all conventional levels.

Note that variable Δm can be considered as stationary at the 5% level following the unit root test. However, this contradiction is not a problem given that the standard methodology to specify VAR models allows for the inclusion of both stationary and non-stationary I(1), [see for example, in Juselius (2006) and most of the citations in that book]. If this variable is stationary we should be able to reduce the cointegration relationships in the monetary market to contain only Δm. One advantage of the Johansen test over standard unit root tests is that it takes into account the information provided by other variables.

Table 2: AUGMENTED DICKEY-FULLER (ADF) UNIT ROOT TESTS

Series	Series in levels		Series in first differences	
	D-F Statistics	Lags	D-F Statistics	Lags
i	-1.34	4	-5.58 (**)	6
π	-1.14	3	-11.51 (**)	2
y	-2.72	5	-3.42 (*)	4
Δm	-6.97 (**)	0	-11.36 (**)	2
i^*	-2.16	1	-4.26 (**)	0
π^*	-1.99	7	-6.81 (**)	6
e	-1.45	2	-9.48 (**)	1
Pr	-1.82	4	-3.54 (**)	3
w	-1.17	5	-1.79	4
u	-2.57	5	-4.19 (**)	4
lp	-1.20	1	-6.34(**)	0

Source: Own estimation.

Note: Number of lags are chosen to minimize the Schwarz criterion.

** (*) denotes rejection at the 0.01 (0.05) significance level. *lp* is the log of the copper price.

Another key point to discuss is the fact that, according to the ADF test in Table 2, Chilean inflation is generated by a non-stationary I(1) process. Estimating a NKPC with non-stationary inflation is not new in the literature and has been considered, for example, by Russell and Banerjee (2008) and Fanelli (2008). In our case, the non-stationary state of Chilean inflation is not imposed; it is a result of the empirical analysis. This result is also consistent with the graphical inspection as the first difference of the log of the Chilean consumer price index for the period 1987:Q1-2009:Q4 shows a downward trend.

A thorough discussion of the empirical implications of describing inflation with a stationary or a unit root process can be found in Culver and Papell (1997). There is no general consensus on this issue. In fact, whether inflation is stationary or non-stationary is not an "intrinsic" property but an empirical matter closely related to the behavior of a certain variable over a specific period.

If we accept the hypothesis that Chilean inflation may contain a unit root, a sound econometric methodology requires inference about the presence of cointegration relationships across the different variables of the model. Otherwise, inferences on the structural parameters based on a NKPC would be seriously flawed[4].

Given the previous discussion, for each of the three markets, we start with the estimation of the following vector equilibrium correction model:

$$\Delta Y_t = \mu + \Pi \begin{pmatrix} Y_{t-1} \\ 1 \\ t \end{pmatrix} + \Phi \Delta Y_{t-1} + \varepsilon_t \qquad [8]$$

where Y_t is the vector of variables included in a given group; $\prod$ is the matrix of parameters whose rank is restricted by the number of cointegration relationships; t is a deterministic trend scalar; μ is a vector of intercept parameters; Φ is a matrix of parameters; and ε_t is a vector of serially uncorrelated errors.

Note that we allow for two deterministic intercepts in expression [8]; one is included in the cointegration equation and the other only affects the short-run dynamics. However, the trend component is only included in the cointegration equation since imposing a quadratic deterministic trend in Y_t is generally an implausible assumption. The presence of deterministic components in the initial specification [8] can be tested in subsequent steps following the principle 'from general to specific' in the specification of the model [see Juselius (2007)].

The following results are obtained from Johansen's cointegration tests:

a) *Monetary and financial market:* the trace test indicates that the null hypothesis of no cointegration can be rejected at 1%. The first column in Table 3 shows the estimated relationship after imposing the normality restriction.

A particularly interesting test, given our previous discussion about unit roots, is to check whether this cointegration relationship could be reduced to contain only Δm_t, i.e. Δm_t is a stationary variable. However, this hypothesis is clearly rejected at the conventional levels; χ^2 (4) = 23.74, p-value = 0.005.

According to the sign of the estimated parameters, the equilibrium relationship can be interpreted as a monetary rule, as the interest rate increases with inflation and output but is negatively affected by money. This rule shows that, given inflation and detrended output, the interest rate will be negatively correlated with changes in nominal balances[5].

Note that our period of analysis includes different monetary regimes. The Central Bank of Chile becomes an autonomous institution from 1989 and the

(4) Cogley and Sbordone (2008) show that it is important to take into account changes in the trend inflation for the NKPC estimation. Although we do not explicitly control for different objectives of the central bank, by estimating cointegration relationships we analyze the group of variables that could endogenously explain changes in inflation trend.

(5) An alternative way to present the monetary rule is as $0.02\left(y_t - \frac{0.0003}{0.02}t\right) - 0.09\Delta m_t$. This formulation allows for the interpretation of the "detrended" income as a short-run determinant of the interest rate. In our case, detrended output is non-stationary and, therefore, provides different information from other standard measures of output gap, such as first differences of GDP (in logs) or the detrended GDP obtained by using the HP filter.

year of adoption of inflation target is 1991 with a range target between 1991 and 1996 and a point target thereafter. Hence, our estimation does not contradict the fact that the Central Bank of Chile has done a very good job in stabilizing inflation in the most recent period. Moreover, the estimated rule indicates that a one percent change in the inflation rate has five times a greater impact on the interest rate than a one percent change in money, showing the preponderance of inflation over emission in the monetary rule.

b) *External Market:* Economic theory typically expects that, in order to fulfill the purchasing power parity, the national rate of inflation should be in equilibrium with the U.S. inflation and the real effective exchange rate index. However, our empirical analysis does not allow us to accept the purchasing power parity cointegration relationship. Consistent with Juselius (1995), we find that a cointegration relationship holds once we introduce interest rate variables to equilibrate the goods and capital markets.

The trace test indicates the presence of a single cointegration relationship. Results of this estimation are reported in the second column of Table 3.

We impose the overidentifying restriction that the parameter associated with U.S. inflation takes a value of -1. This restriction cannot be rejected at the conventional levels using a standard likelihood ratio test (χ^2 (4) = 0.52, p-value = 0.77).

This cointegration equation can be interpreted as an equilibrium relationship between Chilean and U.S. interest rates in real terms.

c) *Labour Market:* the trace test indicates that it is possible to reject the null hypothesis of no cointegration and at least one cointegration relationship at the 1% level. However, it is not possible to reject the null hypothesis of at least two cointegration relationships at the conventional levels.

After imposing overidentifying restrictions implying that real wages are governed by labour productivity and that inflation reacts negatively to productivity growth, the estimated cointegration relationships can be found in the last two columns of Table 3.

The first cointegration relationship (column 3 in Table 3) can be interpreted as a (real) wage function in which wages is related to productivity and unemployment. The second cointegration relation indicates a long-run relationship that involves inflation, productivity and unemployment. Since productivity is inversely related to unit costs, this could be interpreted as a marginal cost pricing equation.

Figure 1 shows the four cointegration equations found in this analysis. A visual inspection of the figure reveals that these are stationary. Although these cointegration relationships are not unique, they satisfy our purpose to obtain equilibrium relationships in the three markets that are simple and can be interpreted economically.

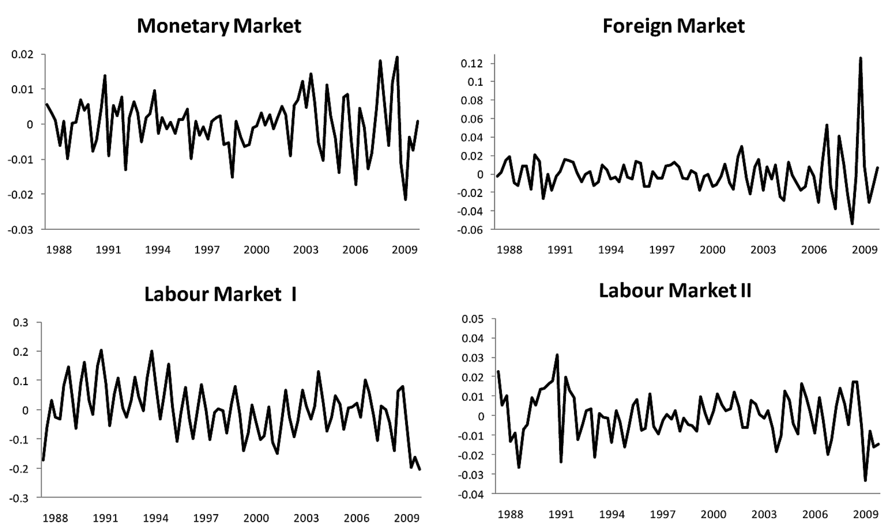
Table 3: ESTIMATED COINTEGRATION RELATIONSHIPS IN THE THREE RELEVANT MARKETS

	Monetary		External		Labour	
i_t	1.00	i_t^*	1.00	w_t	1.00	0.00
π_t	-0.39	π_t^*	-1.00	π_t	0.00	1.00
	(0.02)					
y_t	-0.02	i_t	-0.55	Pr_t	-1.95	0.13
	(0.004)		(0.14)		(0.16)	(0.03)
Δm_t	0.09	π_t	0.26	u_t	-0.03	0.003
	(0.02)		(0.06)		(0.006)	(0.002)
t	0.0003	e_t	-0.003	t	0.01	0.0007
	(5.2x10^{-5})		(0.004)		(0.001)	(0.003)

Source: Own estimation.

Note: Values in brackets are standard errors.

Figure 1: COINTEGRATION RELATIONSHIPS IN THE MONETARY, FOREIGN AND LABOUR MARKETS (*)



Source: Own estimation.

Note: (*) Cointegration relationships area centered by substracting their average values.

3. Estimation Of An Extended Phillips Curve

The information about the equilibrium relationships in different markets can be easily incorporated into the reduced form Phillips curve presented in [4]. To observe this, notice that the estimated equations in Table 3 indicate a long-run equilibrium relation between inflation and a combination of variables. Therefore, they can be written respectively as $\pi_t - \beta^m x^m_{l,t}$, $\pi_t - \beta^x x^x_{l,t}$ and $\pi_t - \beta^l x^l_{l,t}$, where the elements $x^m_{l,t}$, $x^x_{l,t}$ and $x^l_{l,t}$ are combinations of variables that determine the long-run Chilean inflation in the three markets of interest. Additionally, we include the vector of variables affecting short-run inflation dynamics, i.e., $x_{c,t}$ in [4], that contains lagged values of Δy_t; the rest are taken into account in the analysis after being transformed into stationary variables.

We also looked for series related to the non-tradable sector, housing sector or commodity prices, with the only exception being that of the copper price. However, it was not possible to obtain them for the time period considered in the analysis. Therefore, we include the series of the commodity price index of copper in first differences as the ADF applied to this series indicates that it has been generated by a I(1) process. More specifically, the null hypothesis that the log of the price of copper has a unit root can be accepted (p-value 0.67), whereas the null hypothesis that the first differences of the price of copper has a unit root (p-value 0.00001) can be rejected at the conventionally significant levels. The trace test for cointegration between the log of the price of copper and Chilean inflation indicates that the null of no cointegration cannot be rejected at the 5% level. However, although there is no cointegration relationship between these two variables, given Chilean economists, traditional concern about the evolution of the price of copper, we study the impact of the inclusion of this variable in the results.

Table 4 reports results of the two estimations. The first column of the table includes an estimation of the reduced form model obtained following a step-wise procedure but always keeping the lagged dependent variable, the price of copper and the second cointegration relationships in the labour market. In this estimation, the first cointegration relationship in the labour market is not considered as its effect is highly insignificant. In this specification, we considered the inclusion of the stationary transformation of all the variables in the analysis as explanatory variables, but only the lag two of output turned out to be significant.

The sign of the estimated parameters in this estimation is consistent with economic theory. In the asset, exchange rate, and labour markets, when inflation is greater than what is compatible with equilibrium, the disequilibrium position will generate a stabilizing downward movement in inflation rates. The most important factors to determine Chilean inflation in the long-run are the money and labour markets. Excess demand in the foreign markets is not so important, but shows the expected sign. Additionally, inflation is significantly affected in the short run by increased production, as expected, with a two-quarter lag.

Note that, in this specification, the price of copper and the cointegration relationship in the foreign market are not significant at the conventional levels. For this reason, in the second column of Table 4, we present an estimation that excludes all the nonsignificant variables with the only exception of lagged inflation that is required in order to identify the NKPC. This last estimation is more parsimonious and with a lower Akaike criterion and, therefore, it will be used in the subsequent analysis. However, there are no material changes in our analysis if the more complex specification is chosen.

Table 4: ESTIMATION OF A REDUCED FORM EQUATION FOR INFLATION DYNAMICS, $\Delta\pi_t$

	Equation 1	Equation 2
Constant	-0.002	-0.002 (*)
	(0.001)	(0.001)
$\pi_{t\text{-}1} - \beta^m x_{l,t\text{-}1}{}^m$	-0.55 (*)	-0.40 (**)
	(0.21)	(0.15)
$\pi_{t\text{-}1} - \beta^l x_{l,t\text{-}1}{}^t$	-0.41 (**)	-0.57 (**)
	(0.15)	(0.21)
$\pi_{t\text{-}1} - \beta^x{}_{xi,t\text{-}1}{}^x$	-0.11	—
	(0.07)	
$\Delta lp_{t\text{-}1}$	-0.02	—
	(0.01)	
$\Delta\pi_{t\text{-}1}$	-0.03	-0.001
	(0.10)	(0.09)
$\Delta y_{t\text{-}2}$	0.13 (**)	0.15 (**)
	(0.03)	(0.03)
Adj-R^2	0.49	0.49
Akaike	-6.14	-6.16

Source: Own estimation.

Note: ** (*) denotes rejection at the 0.01 (0.05) significance level.

It is also of interest to test two alternative Phillips curves found in the literature. First, we test a traditional NKPC that only allows for short-run dynamics, which amounts to assuming that the two parameters of adjustment to the equilibrium in the three markets are equal to zero. Using an F-test, we reject this restriction at the 1% level. A second alternative model is case 2 in Fanelli's nomenclature (2008). He considers the possibility of common unit roots in the variables included in the Phillips curve. However, his structural model, given cointegrated variables, only allows for the possibility of long-run adjustments. We test this by restricting $\Delta y_{t\text{-}2}$ and $\Delta lp_{t\text{-}1}$ to have no effect on $\Delta\pi_t$. However, this restriction is rejected using an F-test at the 1% level.

Restrictions from [5], [6] and [7] should be imposed on the estimation to obtain the structural equation [3]. We test these restrictions with a standard F-statistic. The test has a p-value of 0.60 and, therefore, we cannot reject these restrictions at the conventional levels. Structural parameters could be obtained by an indirect procedure such as the delta method, which matches the parameters in the reduced form and the structural model using [5], [6] and [7]. However, we do this directly, imposing these restrictions on the estimation of the model. Table 5 shows the results of the estimation of model (3). This table also includes the estimation results of both a model that does not include short-run dynamics and one without long-run dynamics.

Table 5: Structural Estimation

	Model (3)	Model without short-run dynamics	Model without long-run dynamics
Constant	-0.002 (*)	-0.0006	-0.003
	(0.001)	(0.001)	(0.001)
$\hat{\theta}$	1.46 (**)	1.57 (**)	
	(0.09)	(0.09)	
$\hat{\delta}$	0.004	-0.08	-0.26 (**)
	(0.09)	(0.10)	(0.09)
$\hat{\lambda}$	0.15 (**)	0.17 (**)	
	(0.03)	(0.04)	
Adj-R^2	0.49	0.35	0.17
Akaike	-6.18	-5.95	-5.67

Source: Own estimation.

Note: Values in brackets are standard errors.

** (*) denotes rejection at the 0.01 (0.05) significance level.

The fit of the model that includes both short and long-run adjustments clearly outperforms a model with only long-run or only short-run adjustments. Moreover, imposing these restrictions affects the value of the estimated structural parameters. We also tested whether the backward and forward looking coefficients add to unity in the global specification [see Galí and Gertler (1999)]. Note that imposing this restriction amounts to assuming that the cointegration relationships do not affect the dynamics of inflation in expression [3]. However, this restriction is clearly rejected at all the conventional levels, (F-statistics = 34.73, p-value = 0.003).

The first thing to notice is that the inflation dynamics resulting from the three estimations are stable because the roots of the three characteristic equations lie outside the unit circle. More interestingly, expected (and not lagged) inflation has a significant effect on current price decisions. This is consistent with the Phillips curve (equation 2) proposed by Clarida *et al*. (1998) in the sense that there is no arbitrary inertia or lagged dependence on inflation. Rather, inflation depends entirely on current and expected economic conditions, as firms set nominal prices based on the expectations of future marginal costs. Empirical estimations by Gali *et al.* (2001) using data from the U.S. and the Euro Zone lead them to a similar conclusion.

We test the robustness of our results by comparing the estimation reported in Table 3 with those obtained from the GMM estimation of a multimarket approach version of the Chilean NKPC. In general, GMM and ML can be regarded as two competing procedures used in the economic literature to estimate NKPC. Under ideal circumstances, GMM is preferred as it requires minimum assumptions about the explanatory variables. However, it is also well documented that GMM estimates can be markedly biased in small samples because of "weak instrument" issues [see Stock *et al*. (2002)].

Given this discussion, we consider it a highly relevant experiment to further explore the consideration of information coming from demand pressures in different markets in order to estimate a multimarket approach of the NKPC by using the GMM estimation procedure. An additional reason for this experiment is the fact that estimating a NKPC in which the dependent variable is in first differences (instead of level) of inflation is not consistent with standard practice in the literature. Therefore, the vector of variables that indicates demand pressures is redefined in order to include information from cointegration relationships in the three relevant markets and also to ensure that all variables included in vector x_t are stationary. Hence, x_t is now split as $x_t' = (x_{l,t}', x_{c,t}')$, where $c_{l,t}'$ is the vector of stationary cointegration relationships found in the previous section and $x_{c,t}$ is a vector of individual stationary variables.

Thus, the following model is estimated by GMM

$$\pi_t = \theta E_t \pi_{t+1} + \delta \pi_{t-1} + \lambda_1' c_{l,t} + \lambda_2' x_{c,t} + \varepsilon_t \qquad [9]$$

where ε_t is a stochastic error term.

Note that model [9] also encompasses the case of a traditional NKPC in which x_t is a scalar that does not contain information about cointegration relationships in different markets.

Similar to the previous estimation, in vector $c_{l,t}$ we include the cointegration relationships found in the 1) monetary and financial markets, 2) international markets, and 3) labour markets. Vector $x_{c,t}$ includes the stationary transformation of the Chilean GDP and the price of copper. Our set of instruments considers three lagged values of all the variables included in the previous estimation. In addition, as in Stock *et al*. (2002) and Kleibergen and Sophocles (2009), we also take into account the case of a NKPC with non-stationary inflation by using first differences of the Chilean inflation rate, instead of levels of inflation, as instrument variables in the GMM estimation. This case imposes the restriction $\theta + \delta = 1$ in the NKPC. Results of these estimations are reported in Table 6. Two important results to highlight, that are consistent with our previous structural estimation, are the significant role of both expectations on future inflation (given by $\hat{\theta}$) and, more importantly, the cointegration relationships found in the three key markets (given by $\hat{\lambda}_1$). We also test the restriction $\theta + \delta = 1$ with this specification and find that this restriction could be accepted at all conventional levels (F-statistics = 1.13, p-value = 0.29).

Table 6: Gmm structural estimation

	Unrestricted	Restricted ($\theta + \delta = 1$)
Constant	0.0002 (0.001)	0.0004 (0.0007)
$\hat{\theta}$	0.96 (**) (0.07)	0.72 (**) (0.06)
$\hat{\delta}$	-0.01 (0.07)	
$\hat{\lambda}_1$	0.44 (**) (0.06)	0.46 (**) (0.04)
$\hat{\lambda}_2$	0.08 (**) (0.02)	0.08 (**) (0.01)
Adj-R^2 J-statistic	0.54 0.13	0.74 0.16

Source: Own estimation.

Note: The model is $E(Z_t\ (\theta\ E_t\ \pi_{t+1} + \delta\ \pi_{t-1} + \lambda_1'\ x_{m,t} + \lambda_2'\ x_{c,t})) = 0$. Instruments include a constant and three lags of π_t, $x_{l,t}$ and $x_{c,t}$ (lags of n_t are replaced by lags of Δn_t in the restricted model). Values in brackets are standard errors.

From this analysis it is clear that considering cointegration relationships in the labour and monetary markets improves the description of inflation dynamics over the period of analysis: 1987:Q1-2009:Q4. However, it is still unclear whether this consideration is useful to forecast future inflation movements. To check for this, we forecast the Chilean inflation for the period 2006:Q1-2009:Q4[6]. The initial estimation sample covers the period 1987:Q1-2005:Q4 and forecasts are then obtained by following a recursive scheme [see, for example, Faust *et al.* (2005), and West (2006)]. Under this approach, the sample size used to estimate the parameters of the different models grows as one makes predictions for successive observations, which allows the parameters of the econometric models to change according to the new information in the samples.

Four alternative models are considered in this forecasting experiment. The first one (denoted by total) considers a reduced form equation similar to the one expressed in Table 4 that contains information on cointegration relationships in

(6) The forecasting experiment was stopped at 2009:Q4 as there is an important discontinuity in the employment series at this date. The new series shows different seasonal behavior and patterns of growth that are explained by changes made by the Chilean National Institute of Statistics in the methodology used to collect the series' data and not by a structural change in the labour market in Chile.

the labour and monetary markets as well as output fluctuation and lagged inflation. We compare the inflation forecast under this strategy with those obtained using information on cointegration relationships in each of the three markets separately (monetary, labour and external markets) and from a standard equation that considers only stationary output fluctuations and lagged inflation.

Root mean inflation forecast errors obtained under the different strategies are reported in Table 7. Results indicate that the inflation forecasts estimated by models which consider cointegration relationships in some of the relevant markets always outperform those which use a standard approach. Moreover, according to the Diebold and Mariano test, forecasting the Chilean inflation by means of considering cointegration relationships (especially in the monetary and labour markets) leads to a significant improvement compared to the standard approach at the conventional levels in many cases.

4. Concluding Remarks

In this paper, we propose a new empirical approach for estimating a New Keynesian Phillips Curve that jointly incorporates demand pressures from different markets. This procedure also allows us to analyse inflation dynamics in the short- and long-run. Methodologically, we apply a two-step procedure. In the first step, potential cointegration relationships are analysed separately in three different vector equilibrium models. Inflation is explained by [1] the effect of monetary and financial variables, [2] information on international markets, and [3] production costs. In Chile, a cointegration relationship has been found in the monetary market that could be interpreted as a Taylor rule. The cointegration relation in the foreign market equalizes the real interest rate in the U.S. with the Chilean one, minus the real exchange rate. In the labour market, two cointegration equations were found: one relating productivity to salary and inflation and the other relating price inflation to productivity growth. In the second step, we incorporated the information obtained in the first step about the equilibrium relationships in different markets to estimate the structural parameters of the NKPC. Inflation reacts to shocks in different markets, but equilibrium in the monetary and labour markets and deviations in output are the most important explanatory variables. Our approach can be considered a general case that nests other specifications in the literature, such as a traditional Phillips curve that only includes demand pressures coming from the output gap and the model proposed by Fanelli. Based on econometric tests, we can reject these specifications in favour of our model.

A future line of research could be to apply our procedure to estimate a Complete New Keynesian model instead of just a NKPC. The estimation of New Keynesian models has been previously proposed by Rotemberg and Woodford (1999), Christiano *et al*. (2001), and Boivin and Giannoni (2003), among others. An interesting contribution would be to extend the methodology used herein to include inertial forces coming from cointegration relationships from different markets in a system of equations.

Table 7: Forecast of chilean inflation using alternative econometric specification (Rmsfe)

	Total	Monetary	Labour	External	Standard
			1-step-ahead		
	0.30%	0.30%	0.34%	0.31%	0.36%
			Diebold-Mariano test		
Total	—	-0.17	-1.07	-0.40	-1.86
Monetary		—	-1.14	-0.25	-1.49
Labour			—	0.66	-0.43
External				—	-1.05
Standard					—
			2-step-ahead		
	0.37%	0.33%	0.54%	0.39%	0.38%
			Diebold-Mariano test		
Total	—	0.70	-2.81 (*)	-0.26	-0.37
Monetary		—	-3.54 (**)	-1.10	-2.30 (*)
Labour			—	2.40 (*)	3.50 (**)
External				—	0.08
Standard					—
			3-step-ahead		
	0.39%	0.23%	0.66%	0.39%	0.40%
			Diebold-Mariano test		
Total	—	2.09 (*)	-2.81 (*)	0.05	-0.08
Monetary		—	-4.48 (**)	-1.64	-2.55 (**)
Labour			—	4.00 (**)	5.04 (**)
External				—	-0.20
Standard					—
			4-step-ahead		
	0.36%	0.25%	0.78%	0.44%	0.32%
			Diebold-Mariano test		
Total	—	1.75	-3.68 (**)	-0.65	0.69
Monetary		—	-3.97 (**)	-1.47	-1.92
Labour			—	3.86 (**)	4.15 (**)
External				—	1.28
Standard					—

Source: Own estimation.

Note: ** and * indicate rejection at the 1% and 5%.

Appendix. Source Of The Series

The paper uses series from the first quarter of 1987 to the fourth quarter of 2009. The sources of the basic series are:

Real GDP: National Accounts, Chilean Central Bank.

Employment: National Employment Survey, National Statistics Institute, Chile.

Nominal Wages: Wage Index, National Statistics Institute, Chile.

Consumer Price Index: Consumer Price Survey, National Statistics Institute, Chile.

Unemployment Rate: National Employment Survey, National Statistics Institute, Chile.

Short-Run Nominal Interest Rate: Average bank borrowing rate from 30 to 89 days, Chilean Central Bank.

M2 Money Definition: Chilean Central Bank.

Exchange rate - real effective exchange rate index - Chile index base: 2000 = 1.00, based on relative consumer prices, period average: International Monetary Fund.

U.S. Nominal Fed Rates: Federal Funds Monthly Effective Rate, Federal Reserve of the United States.

U.S. Consumer Price Index: United States Department of Labor, Bureau of Labor Statistics.

PPI Commodity data special indexes - copper and copper products Index Base: 1982 = 100, Bureau of Labour Statistics.

Monthly series were averaged to a quarterly frequency. Nominal wages and M2 were deflated using the consumer price index to obtain series in real terms. The average labour productivity was obtained as the ratio between the real GDP and employment.

REFERENCES

Batini, N., Jackson, B. and Nickell, S. (2005): "An open-economy New Keynesian Phillips curve for the U.K.", *Journal of Monetary Economics,* n. 52, pp. 1061-1071.

Boivin, J. and Giannoni, M. (2003): "Has monetary policy Become more Effective?", *NBER Working Paper*, n. 9459.

Calvo, G. (1983): "Staggered Prices in a Utility-Maximising Framework", *Journal of Monetary Economics*, n. 12, pp. 983-998.

Céspedes, L.F., Ochoa, M. and C. Soto (2005): "The New Keynesian Phillips Curve in an Emerging Market Economy: The Case of Chile", *Central Bank of Chile, Working Paper*, n. 355, December.

Christiano, L., Eichenbaum, M. and Evans, C. (2001): "Nominal rigidities and the dynamic effects of a shock to monetary policy", *NBER Working Paper*, n. 8403.

Clarida, R., Galí, J. and Gertler, M. (1998): "The Science of Monetary Policy: A New Keynesian Perspective", *Journal of Economic Perspectives,* n. 37, 1999, pp. 1661-1707.

Cogley, T. and Sbordone, A.M. (2008): "Trend Inflation, Indexation, and Inflation Persistence in the Newkeynesian Phillips Curve", *American Economic Review, American Economic Association*, n. 98, pp. 2101-26.

Culver, S.E. and Papell, D.H. (1997): "Is there a Unit Root in the Inflation Rate? Evidence from Sequential Break and Panel Data Models", *Journal of Applied Econometrics*, n. 12, pp. 435-444.

Eichenbaum, M. and Fisher, J., (2007): "Estimating the frequency of price re-optimization in Calvo-style models", *Journal of Monetary Economics,* n. 54, pp. 2032-2047.

Faust, J., Rogers, J. and Wright, J. (2005): "News and Noise in G-7 GDP Announcements", *Journal of Money, Credit and Banking,* n. 37, pp. 403-19.

Fanelli, L. (2008): "Testing the New Keynesian Phillips curve through Vector Autoregressive models: results from the Euro area", *Oxford Bulletin of Economics and Statistics*, n. 70, pp. 53-66. Also published as *MPRA Paper*, n. 1617, University Library of Munich, Germany in 2006.

Fanelli, L. (2006): "Testing the New Keynesian Phillips curve through Vector Autoregressive models: Results from the Euro area", *MPRA Paper*, n. 1617, University Library of Munich, Germany.

Galí, J. and Gertler (1999): "Inflation Dynamics: A Structural Analysis"*, Journal of Monetary Economics*, n. 44, pp. 195-222.

Galí, J. and Gertler, M. and López-Salido, J.D. (2001): "European Inflation Dynamics", *European Economic Review*, n. 45, pp. 1237-1270.

Garrat, A., Lee, K., Pesaran, M.H. and Y. Shin (2000): "A structural cointegrating VAR approach to macroeconomic modelling", in S. Holly and M. Weale (eds): *Econometric Modelling: Techniques and Applications*, Cambridge University Press, Cambridge.

Hendry, D. (2001): "Modelling UK inflation, 1875-1991", *Journal of Applied Econometrics*, n. 16, pp. 255-275.

Johansen, S. (1995): "Likelihood-based Inference in Cointegrated Vector Autoregressive models", *Oxford University Press.*

Juselius, K. (1992): "Domestic and foreign effects on prices in an open economy, the case of Denmark", *Journal of Economic Policy Modeling*, n. 14, pp. 401-428.

Juselius, K. (1995): "Do the purchasing power parity and the uncovered interest rate parity hold in the long run? – An application of likelihood inference in Multivariate time series model", *Journal of Econometrics*, n. 69, pp. 211-240.

Juselius, K. (2007): "The Cointegrated VAR Model: Methodology and Applications", *Oxford University Press*.

Juselius, K. and MacDonald, R. (2004): "The International Parities Between USA and Japan", *Japan and the World Economy,* n. 16, 1, pp. 17-34.

Kleibergen, F. and Sophocles, M. (2009): "Weak Instruments Robust Tests in GMM and the New Keynesian Phillips Curve", *Journal of Business & Economic Statistics*, n. 27, pp. 293-311.

Peturson, T.G. (1998): "Price Determination and Rational Expectations", *International Journal of Finance and Economics*, n. 3, pp. 157-167.

Rotemberg, J. and Woodford, M. (1999): "Interest rate rules in an estimate sticky price model", in John B. Taylor (ed.): *Monetary Policy Rules Chicago: University of Chicago Press*, pp. 57-119.

Rudd, J. and Whelan, K. (2005): "New tests of the New Keynesian Phillips Curve", *Journal of Monetary Economics*, n. 52, pp. 1167-1181.

Russell, B. and Banerjee, A. (2008): "The long-run Phillips curve and non-stationary inflation", *Journal of Macroeconomics*, Elsevier, n. 30, pp. 1792-1815.

Stock, J.H., Wright, J.H. and Yogo, M. (2002): "A survey of weak instruments and weak identification in Generalized Method of Moments", *Journal of Business and Economic Statistics*, n. 20, pp. 518-529.

West, K.D. (2006): "Forecast Evaluation" in Elliott G., Granger C.W.J. and Timmermann A. (eds.): *Handbook of Economic Forecasting,* North Holland.

Fecha de recepción del original: agosto, 2009
Versión final: septiembre, 2011

RESUMEN

Se propone una nueva metodología para estimar una curva Neo-Keynesiana "híbrida" (NKPC) que incluye presiones de demanda provenientes de desequilibrios en tres diferentes mercados: (1) monetarios y financieros, (2) internacional y (3) trabajo. Contrastes econométricos basados en datos chilenos indican que esta especificación es superior a una curva NKPC tradicional, que incluye una única variable para capturar presiones de demanda. Es más, la consideración de relaciones de cointegración en los mercados monetario y laboral resulta particularmente útil para predecir la dinámica de la inflación en comparación con especificaciones alternativas basadas en medidas estándar de fluctuaciones de producción.

Palabras clave: Curva Neo-Keynesiana, cointegración, política monetaria.

Clasificación JEL: E3, C3.