The New Phillips Curve in Canada

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Introduction

The short-run dynamics of inflation and the cyclical interaction of inflation with real aggregates are important issues both in theory and in practice, especially for central banks in their conduct of monetary policy. Recent high levels of economic activity coupled with low inflation, observed in several countries, cast doubt on the traditional Phillips curve as a model of inflation dynamics.

A recent class of dynamic stochastic general-equilibrium model integrates Keynesian features, such as imperfect competition and nominal rigidities, resulting in a new view of the nature of inflation dynamics. These models are grounded in an optimizing framework where imperfectly competitive firms are constrained by costly price adjustments. Within this framework, the process of inflation is described by the so-called New Keynesian Phillips curve (NKPC), which has two distinguishing features. First, the inflation process has a forward-looking component and second, it is related to real marginal costs. These features are a consequence of the fact that in this framework firms set prices in anticipation of future demand and factor costs. Compared with traditional reduced-form Phillips curves, which are subject to the Lucas critique, the NKPC is a structural model with parameters that are unlikely to vary as the policy regime changes. This aspect is important in a country such as Canada, because parameter instability in reduced-form models is a likely possibility since the adoption of an explicit inflation-targeting regime. Furthermore, the NKPC specification has dramatic implications for the conduct of monetary policy in that a fully credible central bank can bring about disinflation at no recessionary cost if inflation is a purely forward-looking phenomenon. A crucial issue is therefore whether the NKPC is empirically relevant.

The recent work of Galí and Gertler (1999) and Galí, Gertler, and López-Salido (2001a) provide evidence supporting the NKPC for the United States and the euro area. These authors estimate hybrid versions of the NKPC, where lags of inflation are also incorporated, and conclude that the forward-looking component is more important and, furthermore, that real marginal costs are statistically significant. In these studies, parameter estimates are obtained by the Generalized Method of Moments (GMM) and statistical significance is assessed based on Newey-West estimates of the covariance matrix.

In this paper, we examine the empirical relevance of the NKPC for Canada. We address several important econometric issues with the standard approaches typically used for estimation and inference in NKPC models. The main issues are related to the potential bias of GMM estimates in the presence of many instruments and the low power of specification tests based on overidentifying restrictions. The approach adopted in this paper attempts to mitigate these econometric problems. Furthermore, we investigate the robustness of our estimation results based on this improved approach relative to the choice of instruments.

The rest of the paper is organized as follows. In section 1, we present the theoretical framework that yields the NKPC, outline alternative measures of marginal cost, and show how open-economy considerations can be incorporated. In section 2, we describe the econometric issues associated with standard GMM estimation, discuss particular issues with estimation of the closed-form version of the NKPC, and present our estimation strategy based on the bias-corrected continuous updating estimator (CUE). In particular, using the same data set as Galí and Gertler (1999), we demonstrate the sensitivity of standard GMM estimates to the choice of instruments. In section 3, we describe various measures of the labour share with Canadian data and then, in section 4, we present the estimation results. A discussion of the main findings follows in section 5, and the final section concludes.

1 New Phillips Curves

The NKPC, as advocated by Galí and Gertler (1999), is based on a model of price-setting by monopolistically competitive firms. Adopting a price-setting rule as in Calvo (1983) simplifies the aggregation problem. This price-adjustment rule is in the spirit of Taylor's (1980) model of staggered contracts. Following Calvo, each firm, in any given period, may adjust its price with a fixed probability $1 - \theta$ and, with probability θ, its price will be kept unchanged or proportional to trend inflation, Ω.[1] These adjustment probabilities are independent of the firm's price history, such that the proportion of firms that may adjust their price in each period is randomly selected. The average time over which a price is fixed is then given by $1/(1-\theta)$.

A firm that sets its price at the beginning of period *t* maximizes its stock market value by solving the following problem:

$$\max_{P_t^*(s)} E_t \sum_{i=0}^{\infty} (\theta\beta)^i \lambda_{t+i} [P_t^*(s)\Omega^i Y_{t+i}(s) - TC(Y_{t+i}(s))] \quad ,(1)$$

where $1-\theta$ is the probability that it may adjust its price at the beginning of a given period, β is the subjective discount rate of the representative owner of the firm, λ_{t+i} is the marginal utility of consumption of the representative owner in period $t+i$, and $Y_{t+i}(s)$ is the firm's output in period $t+i$. $TC(Y_{t+i}(s))$ is the firm's nominal total cost as a function of output. The firm faces a constant elasticity of demand for its output equal to μ. The solution of this maximization problem leads to optimal price-setting rules, which relate a firm's optimal price to its real marginal cost of production and to its expected future optimal price.

For a firm that adjusts its price at time *t*, the optimal reset price is given by:

$$p_t^* = (1-\beta\theta)E_t \sum_{j=0}^{\infty} (\beta\theta)^j mc_{t+j,t}, \quad (2)$$

where $mc_{t+j,t}$ is the firm's nominal marginal cost (as a percentage deviation of the steady state) for a optimal price fixed at time *t*. This expression relates the optimal price to the stream of the future path of discounted nominal marginal cost of the individual firm. It can also be shown that the aggregate price, p_t, depends on the optimal reset price, p_t^*, and the lagged price level p_{t-1} through:

$$p_t = (1-\theta)p_t^* + \theta p_{t-1}. \quad (3)$$

By combining equations (2) and (3), a Phillips curve relationship can be derived relating current inflation to expected future inflation and to firms' real marginal costs. To obtain a Phillips curve relationship of the average real marginal costs of a firm, the firm's real marginal cost has to be aggregated. Unfortunately, the aggregation problem has been solved only

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1. This adjustment is necessary if there is trend inflation in order to preserve monetary neutrality in the aggregate.

under very restrictive assumptions. Yun (1996) and Goodfriend and King (1997) assumed that individual firms can instantaneously adjust their own capital stocks, so that the marginal productivity of capital is the same across all firms, and all firms have the same marginal cost. Danthine and Donaldson (2002) have criticized this approach, since it amounts to assuming that the costs of adjusting physical capital stocks are an order of magnitude smaller than the costs of adjusting prices. Sbordone (2001) showed that under the assumption that firms' relative capital stocks do not vary with their relative prices, and (with a Cobb-Douglas production technology) firms' average marginal costs can be approximated by average unit labour costs. This assumption seems as unsatisfactory as Yun's, since if there is aggregate capital accumulation, then firms have identical rates of net investment. The approaches of both Yun and Sbordone are theoretically unappealing and may be at odds with the data.

Ambler, Kurmann, and Guay (2002) show how to relate the average real marginal cost of firms that adjust their price to the average real marginal cost of firms without specific assumptions.

This New Phillips curve has the same functional form as previous New Phillips curves in the literature, but its parameters depend differently on the underlying structural parameters. In particular, the effects of average real marginal costs and future expected inflation on current inflation depend on both the elasticity of real marginal cost with respect to output and on the demand elasticities of firms.

By a first-order expansion:

$$mc_{t+j} \approx mc_{t+j,t} + \frac{\partial mc_{t+j,t}}{\partial y_{t+j,t}} \frac{\partial y_{t+j,t}}{\partial p_t^*} (p_{t+j} - p_t^*),$$

where $mc_{t+j,t}$ is the real marginal cost of the firm at $t+j$ when its price is fixed at t, and $y_{t+j,t}$ is the firm's output at $t+j$ for a price fixed at t. We can rewrite this as:

$$mc_{t+j,t} \approx mc_{t+j} - \eta\mu(p_t^* - p_{t+j}), \qquad (4)$$

where η is the elasticity of marginal cost with respect to output, and μ represents the demand elasticities of firms.

The New Phillips curve is then given by:

$$\pi_t = \frac{(1-\theta)(1-\theta\beta)}{\theta - \theta\eta\mu} mc_t + \beta E_t \pi_{t+1}. \qquad (5)$$

The derivations in Yun (1996) and Goodfriend and King (1997) correspond to the case where the elasticity of marginal cost with respect to output (η) is equal to zero. Indeed, the hypothesis that individual firms can instantaneously adjust their own capital stocks implies that firms act as price-takers in the input market. Combined with a constant return to scale technology, real marginal cost is then independent of output.

Under the assumption that the relative capital stock does not vary with changes in the relative price or relative output, the individual firm's real marginal cost is related to the aggregate real marginal cost by the following expression:

$$mc_{t,t} = mc_t + \frac{\alpha}{(1-\alpha)}\mu(p_t^* - p_t),$$

which can be rewritten as

$$mc_{t,t} = mc_t + \frac{\alpha}{(1-\alpha)}\frac{\theta}{(1-\theta)}\mu\pi_t,$$

where α is the capital share in the constant return to scale Cobb-Douglas production. With the general formulation (4), this assumption implies the following relationship:

$$\frac{\alpha}{(1-\alpha)} = \eta,$$

which was also shown to hold by Ambler, Kurmann, and Guay (2002).

Finally, under certain conditions, average marginal costs are in turn related to output, thereby linking the New Phillips curve with the traditional Phillips curve (which has the output gap as a key explanatory variable).

Galí and Gertler (1999) extend the basic Calvo model to allow a subset of firms to use a backward-looking rule of thumb to capture the inertia in inflation. The hybrid version of the Phillips curve for the general formulation developed by Ambler, Kurmann, and Guay (2002) is given by:

$$\pi_t = \lambda\left(\frac{1}{(1-\eta\mu)}\right)mc_t + \gamma_f E_t\pi_{t+1} + \gamma_b\pi_{t-1},$$

where

$$\lambda = \left(\frac{(1-\omega)(1-\theta)(1-\theta\beta)}{\theta}\right)\phi^{-1},$$

$$\gamma_f = \beta\theta\phi^{-1},$$

$$\gamma_b = \omega\phi^{-1},$$

$$\phi = \theta + \omega[1 - \theta(1-\beta)],$$

and where ω is the proportion of firms that use a backward-looking rule of thumb. The corresponding hybrid New Phillips curve for the aggregate assumption considered by Yun (1996) and Goodfriend and King (1997) is derived in Galí and Gertler (1999) and the one based on the assumption of Sbordone (2001) in Galí, Gertler, and López-Salido (2001). One can easily retrieve these specific forms from the general form given above.

1.1 Measures of marginal cost

Alternative measures of the marginal cost have been considered in empirical investigations of the NKPC. The simplest measure of real marginal cost is based on the assumption of Cobb-Douglas technology (see Galí and Gertler 1999). Suppose the following Cobb-Douglas production function:

$$Y_t = K_t^{\alpha}(A_t H_t)^{(1-\alpha)},$$

where K_t is the capital stock, A_t is labour-augmenting technology, and H_t represents hours worked. Real marginal cost is then given by $S_t/(1-\alpha)$, where $S_t = W_t H_t / P_t Y_t$ is the labour-income share. In log-linear deviation from the steady state, we have:

$$mc_t = s_t = w_t + h_t - p_t - y_t.$$

Rotemberg and Woodford (1999), Galí, Gertler, and López-Salido (2001a), Gagnon and Khan (2001), and Sbordone (2001) consider a Cobb-Douglas technology with overhead labour cost. In this case, the production function has the following form:

$$Y_t = K_t^{\alpha}(A_t(H_t - \overline{H}))^{(1-\alpha)},$$

where the term $\overline{H}$ is the hours that need to be worked irrespective of the level of production. Expressed in log-linear deviations, we have from Sbordone (2001) that

$$mc_t = s_t + bh_t,$$

where

$$b = \frac{\bar{H}/H}{1-\bar{H}/H}$$

and H is the number of hours worked at the steady state. The measure of marginal cost is in this case augmented by a term that depends on hours worked.

Finally, we can consider adjustment costs of labour. To this end, we specify the following functional form for the adjustment cost of labour:[2]

$$\frac{\phi}{2}(H_t - H_{t-1})^2,$$

where ϕ is the coefficient controlling the adjustment labour cost. It can be shown that real marginal cost in log-linear deviations is then given by:

$$mc_t = s_t + \phi H\left(\frac{H}{(1-\alpha)Y}\right)\Delta H_t - \beta\phi H\left(\frac{H}{(1-\alpha)Y}\right)E_t\Delta H_{t+1}, \qquad (6)$$

where H/Y is the steady-state value of the hours worked share in output. This specification of the real marginal cost implies more dynamics by the intermediary of the expectation of hours worked than the previously considered specifications.

1.2 Open-economy considerations

Following Galí and Monacelli (2002), in the case of an open economy, the real marginal cost can be expressed in log-linear deviations as:

$$mc_t = s_t + (\tilde{p}_t - p_t),$$

where $\tilde{p}_t$ corresponds to the consumer price index.[3] The variable $\tilde{p}_t$ is defined by the following expression:

$$\tilde{p}_t = (1-\Theta)p_t + \Theta p_{mt},$$

where p_{mt} is the import price index (in log deviation), and Θ measures the degree to which the economy is open. By combining these two expressions, the real marginal cost can be rewritten as:

2. Ambler, Guay, and Phaneuf (1999) present estimates of the parameter associated with the adjustment cost of labour in a dynamic stochastic general-equilibrium model.

3. See also Ambler, Kurmann, and Guay (2002) for an alternative derivation of this expression for the real marginal cost.

$$mc_t = s_t + \Theta(p_{mt} - p_t), \tag{7}$$

where $(p_{mt} - p_t)$ corresponds to the terms of trade.

By the law of one price, the real exchange rate is proportional to the terms of trade (Galí and Monacelli 2002). Therefore,

$$q_t = (1 - \Theta)(p_{mt} - p_t),$$

where q_t is the real exchange rate. The real marginal cost can then be expressed as a function of the real exchange rate, as follows:

$$mc_t = s_t + \frac{\Theta}{1 - \Theta} q_t. \tag{8}$$

The estimation of the NKPC in the open-economy case will be based on specifications (7) and (8).

2 Estimation Issues

2.1 Standard GMM approach

The hybrid model in reduced form can be written as:

$$\pi_t = \gamma_f \pi_{t+1} + \gamma_b \pi_{t-1} + \lambda mc_t + \varepsilon_{t+1}, \tag{9}$$

where ε_{t+1} is an expectational error term orthogonal to the information set in period *t*, i.e.,

$$E[(\pi_t - \gamma_f \pi_{t+1} - \gamma_b \pi_{t-1} - \lambda mc_t) Z_t] = 0, \tag{10}$$

where Z_t is a vector of instruments dated *t* and earlier. The orthogonality condition in equation (10) then forms the basis for estimating the model by the GMM. Galí and Gertler (1999) use this technique with four lags each of inflation, the labour income share, the output gap,[4] the long-short interest rate spread, wage inflation, and commodity price inflation. Finally, they use a 12-lag Newey-West estimate of the covariance matrix to obtain standard errors for the model parameters. Based on these choices, they conclude that: (i) the model is statistically significant; and (ii) γ_f is statistically larger than γ_b. They interpret these results as support for the NKPC in the case of the United States. Given the relatively large number of moment conditions,[5] the

4. Typically, the output gap is obtained by application of the Hodrick-Prescott filter or by fitting a quadratic trend to the entire sample. Using such measures of the output gap as instruments is invalid since they violate the basic GMM orthogonality condition.

5. In fact, 24 moment conditions to estimate three reduced-form parameters.

estimates reported by Galí and Gertler are potentially biased, since it is well known that the estimation bias increases with the number of moment conditions in the standard GMM approach (Newey and Smith 2001). To illustrate this effect, consider the following Monte Carlo experiment. Suppose the data are generated by the ARMA process:

$$y_t = \rho y_{t-1} + \varepsilon_t + \theta\varepsilon_{t-1},$$

where $\rho = 0.1$, $\theta = 0.5$, and $\varepsilon_t \sim i.i.d.\ N(0, 1)$. Consistent estimates of ρ are obtained by GMM. The moment conditions are based on

$$E(\varepsilon_t Z_t) = 0,$$

where $Z_t = (y_{t-2}, y_{t-3}, \ldots, y_{t-k})'$ is a vector of valid instruments (since it excludes y_{t-1}). The sample size is fixed at 100, and we study the effect of an increase in the number of moment conditions. The Monte Carlo experiment is based on 10,000 replications, and the automatic lag selection procedure of Newey and West (1994) is used to obtain an estimate of the weighting matrix. Table 1 reports the bias of the GMM estimator as a function of the number of moment conditions, $k - 2$. The bias of the GMM estimator clearly increases with the number of moments (lags of y_t) included in the vector of instruments. With two instruments, the estimator is nearly unbiased. With ten instruments, the bias appears to be of the same order as the true parameter value. This simple Monte Carlo experiment concurs with the theoretical results of Newey and Smith (2001).

Table 1
Bias of GMM estimator

$k-2$	ρ_{GMM}	bias
2	0.0942	–0.0058
3	0.1157	0.0157
4	0.1209	0.0109
5	0.1410	0.0410
6	0.1446	0.0446
7	0.1607	0.0607
8	0.1716	0.0716
9	0.1799	0.0799
10	0.1932	0.0932

A number of studies have also estimated NKPCs in countries other than the United States, applying equally arbitrary choices for the instrument set and the number of lags used in the construction of Newey-West standard errors.[6] See, for example, Batini, Jackson, and Nickell (2002); Galí, Gertler, and López-Salido (2001a); Gagnon and Khan (2001); and Balakrishnan and López-Salido (2002).

To appreciate the relative importance of these choices within a standard GMM context, let $\gamma = \gamma_f$ and consider the reduced form under the constraint $\gamma_f + \gamma_b = 1$:

$$\pi_t(\gamma) = \lambda(\gamma) mc_t + \varepsilon_{t+1}, \tag{11}$$

where $\pi_t(\gamma) = \pi_t - \pi_{t-1} - \gamma(\pi_{t+1} - \pi_{t-1})$. For a fixed value of $\gamma \in [0, 1]$, the parameter $\lambda(\gamma)$ can be consistently estimated by instrumental variables, using lagged values of real marginal cost dated t and earlier.

Using the same data set[7] as Galí and Gertler (1999), Figure 1 shows the effects of different instruments and those of various lags in constructing Newey-West estimates of the standard deviation. For a given instrument, it appears that there is little effect whether 8, 12, or 16 lags are used for the Newey-West standard errors. On the other hand, it is clear that the choice of instrument is crucial, especially at the upper end of the interval [0,1], where the forward-looking component in the New Phillips curve is more important. When the sixth lag of marginal cost is used as the instrument, marginal costs tend to appear marginally significant for some values of the forward-looking component parameter near 0.7, while it is clearly insignificant when the fourth lag is used as the instrument. Note also the increased precision when the fourth lag is used as the instrument as reflected by the relatively tighter confidence bands. The difference in the width of the confidence bands is expected, since the more recent lags are more strongly correlated with contemporaneous marginal cost and hence are better instruments.

Overall, these results cast doubt on the robustness of the results reported by Galí and Gertler (1999) and on the significance of marginal costs in explaining U.S. inflation.

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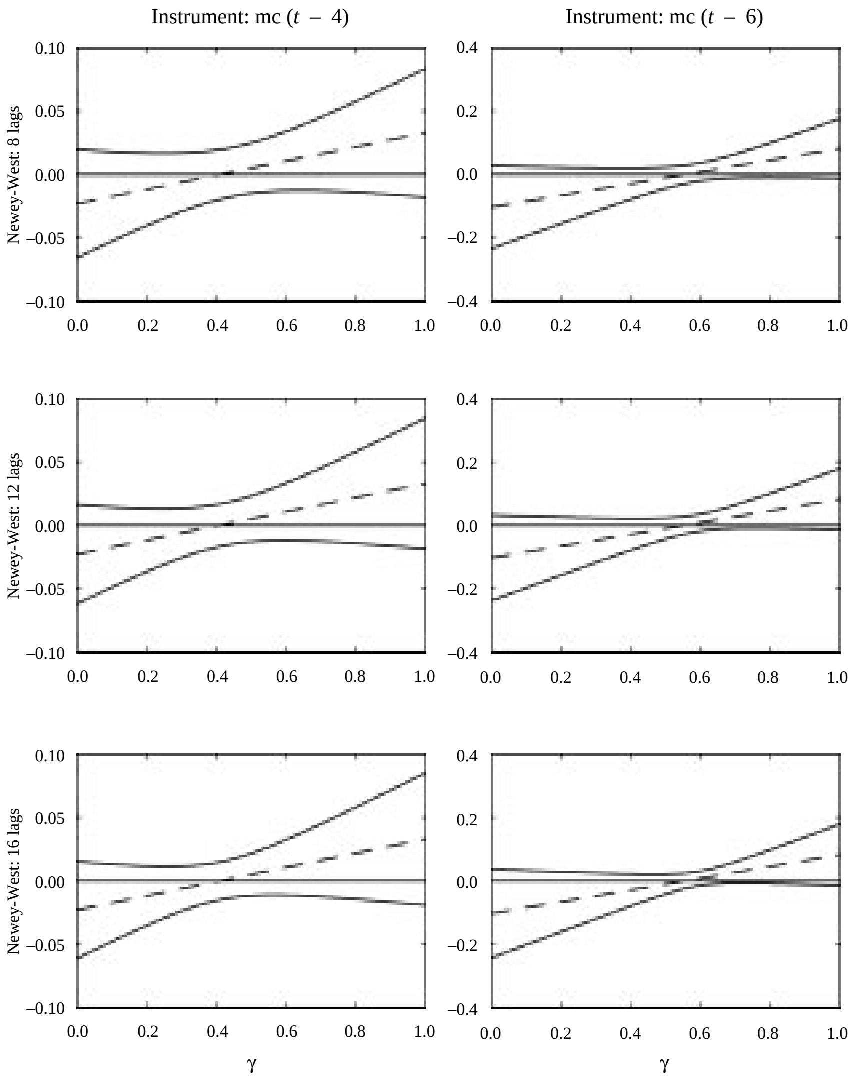
6. A few notable exceptions are Jondeau and Le Bihan (2001) and Lindé (2001) who consider full information maximum likelihood approaches.

7. The data are quarterly for the United States over the period 1960Q1–1997Q4. Inflation is the annualized change in the logarithm of the GDP deflator, and real marginal costs are measured as deviations from the sample mean of the logarithm of labour income share in the non-farm business sector.

Figure 1
Effects of different instrumental variables



Notes: The dashed line in each graph shows the instrumental variable (IV) estimates of $\lambda(\gamma)$ in the model $\pi_t(\gamma) = \lambda(\gamma) mc_t + \varepsilon_{t+1}$, where the instrument used is either the fourth lag (left panel) or the sixth lag (right panel) of real marginal cost. Newey-West standard errors are used to construct the 95 per cent confidence bands, using either 8, 12, or 16 lags.

2.2 Closed-form estimation à la Rudd-Whelan

As shown in Galí and Gertler (1999), the hybrid Phillips curve has the following closed form, conditional on the expected path of real marginal cost:

$$\pi_t = \delta_1 \pi_{t-1} + \frac{\lambda}{\delta_2 \gamma_f} \sum_{k=0}^{\infty} \delta_2^{-k} E_t[mc_{t+k}], \tag{12}$$

where δ_1 and δ_2 are, respectively, the stable and unstable roots of the hybrid Phillips curve given by:

$$\delta_1 = \frac{1 - \sqrt{1 - 4\gamma_b \gamma_f}}{2\gamma_f}, \qquad \delta_2 = \frac{1 + \sqrt{1 - 4\gamma_b \gamma_f}}{2\gamma_f}. \tag{13}$$

An alternative to the standard GMM approach is to directly estimate the closed-form representation, as done in Rudd and Whelan (2001) and Galí, Gertler, and López-Salido (2001a). Under rational expectations, the closed form defines the following orthogonality conditions:

$$E_t\left[\left(\pi_t - \delta_1 \pi_{t-1} - \frac{\lambda}{\delta_2 \gamma_f} \sum_{k=0}^{\infty} \delta_2 mc_{t+k}\right) Z_t\right] = 0, \tag{14}$$

where Z_t is a vector of instrumental variables.

With this approach, it is necessary to use a truncated sum to approximate the infinite discounted sum of real marginal costs. Based on an assumed value for the discount factor β, Rudd and Whelan use 12 leads of real marginal cost to construct the discounted stream of real marginal costs. Galí, Gertler, and López-Salido, on the other hand, use 16 leads and differ by estimating the discount factor instead of fixing its value arbitrarily. In both cases, however, there is loss of degrees of freedom because of the need to truncate the sum, which can be important given the relatively small sample size (typically about 30 years of quarterly data). Furthermore, given the way the measure of the discounted stream of future marginal cost is constructed, there is a generated regressor problem. To see this, consider the limiting case of pure forward-looking behaviour. In that case, the closed form, under rational expectations, becomes

$$\pi_t = \lambda \sum_{k=0}^{\infty} \beta^k mc_{t+k} + u_{t+1}, \tag{15}$$

where the new error term, u_{t+1}, is related to the original expectational error term, ε_{t+1}, by

$$u_{t+1} = \varepsilon_{t+1} + \gamma_f \pi_{t+1} - \lambda \sum_{k=1}^{\infty} \beta^k mc_{t+k}, \tag{16}$$

and from which the generated regressor problem is apparent. Since u_{t+1} in equation (16) is serially correlated (into to the indefinite future), it is essential that the efficiency of the GMM estimator and the consistency of the associated standard errors be evaluated. Clearly, this problem is also present in the hybrid Phillips curve. Estimation in the presence of generated regressors leads, in general, to inefficient estimates that require adjustments to obtain consistent estimates of their standard errors (see Pagan 1984, 1986; Murphy and Topel 1985; and McAleer and McKenzie 1991a, b). Galí, Gertler, and López-Salido (2001b) recognize this problem, but no attempt is made to evaluate it.

Another problem associated with the closed form is that it involves locally almost unidentified (LAU) parameters such that use of Wald-type confidence intervals is invalid. The problem here is that the ratio $\lambda/(\delta_2\gamma_f)$ has a discontinuity at every point of the parameter space where $\gamma_f = 0$. From Dufour (1997), it is then known that one can find a value of this ratio such that the distribution of the Wald statistic will deviate arbitrarily from any "approximating distribution" (such as the standard normal distribution). This suggests that Wald-type inference on structural parameters that appear in NKPC models in ratio form is, in general, an issue for any of the usual estimation approaches. Other techniques, such as confidence sets based on the inversion of likelihood ratio tests, would yield valid inference on the LAU structural parameters. Note that Wald-type inference remains valid for the "non-LAU" reduced-form parameters.

2.3 Estimation strategy

Our estimation strategy differs in three important ways from other empirical studies of the NKPC. First, we use bias-corrected estimators as proposed by Newey and Smith (2001). Second, an automatic lag-selection procedure proposed by Newey and West (1994) is adopted to compute estimates of the variance-covariance matrix of the moment conditions.

As shown by several studies, the small sample properties of method-of-moments estimators depends crucially on the number of lags used in the computation of this variance-covariance matrix. Moreover, our estimator of the variance-covariance matrix uses the sample moments in mean deviation in order to increase the power of the overidentifying restrictions test as suggested by Hall (2000). A more powerful specification test is clearly desirable, since it addresses the issues raised by Dotsey (2002) who found that the conventional specification test used in Galí and Gertler (1999) lacks power. Third, an alternative estimator is used for the non-linear specification. This estimator has the advantage that it does not depend on the normalization of the moment conditions, in contrast to the conventional GMM estimator.

We begin by presenting an alternative estimator to the conventional two-step GMM estimator: the CUE, introduced by Hansen, Heaton, and Yaron (1996). We then present bias-corrected linear IV, GMM, and CUE, as proposed by Newey and Smith (2001).

The optimal two-step GMM estimator of Hansen (1982) based on the moment condition

$$E[g(z_t, \beta_0)] = 0$$

is defined as

$$\hat{\beta} = arg \min_{\beta \in B} \frac{1}{T}\sum_{t=1}^{T} g(z_t, \beta)'\hat{\Omega}(\tilde{\beta})^{-1}\frac{1}{T}\sum_{t=1}^{T} g(z_t, \beta),$$

where $\tilde{\beta}$ is a first-step estimator usually obtained with the identity matrix as weighting matrix, and where $\hat{\Omega}^{-1}$ is a consistent estimator of the inverse of the variance-covariance matrix of the moments conditions.

The CUE is analogous to GMM, except that the objective function is simultaneously minimized over β and $\hat{\Omega}(\beta)$. This estimator is given by

$$\hat{\beta} = arg \min_{\beta \in B} \frac{1}{T}\sum_{t=1}^{T} g(z_t, \beta)'\hat{\Omega}(\beta)^{-1}\frac{1}{T}\sum_{t=1}^{T} g(z_t, \beta).$$

This estimator has important advantages compared with the conventional two-step estimator. First, unlike GMM, the estimator does not depend on the normalization of the moment conditions. As shown by Galí and Gertler (1999), the results obtained for the New Phillips curve and the hybrid version depend on the normalization adopted for the GMM estimation procedure. Second, Newey and Smith (2001) have shown that the asymptotic bias of CUE does not increase with the number of moment conditions. Hansen, Heaton, and Yaron (1996) show that in small samples, the CUE has smaller bias for IV estimators of asset-pricing models with several overidentifying restrictions compared with that of GMM.

Several approaches exist to correct for biases, including the jackknife, the bootstrap, subsampling, and analytical methods. Newey and Smith (2001) proposed an analytical bias correction for GMM and CUE estimators based on asymptotic bias formulas. Those formulas are derived from a stochastic expansion to study higher than $1/\sqrt{T}$ order properties of GMM and Generalized Empirical Likelihood estimators. The bias-corrected CUE is used for estimation of the non-linear specification. The bias formulas from Newey and Smith (2001) are adapted to a dynamic context.[8] This analytical bias correction is much simpler computationally than resampling methods, especially in non-linear models.

3 Measuring the Labour Share with Canadian Data

The labour share is given by:

$$\text{labour share} = \frac{WN}{PY}, \tag{17}$$

where WN is nominal labour income and PY is nominal output. From this basic definition, several measures of labour share can be constructed using available Canadian data.

A natural measure of the labour share is simply the ratio of total compensation of employees in the economy divided by the national income, i.e.,

lshare = wages and salaries/total GDP.

There are, however, some conceptual issues with the appropriate measure to use in order to be consistent with the model's theoretical framework. First, the measure should be net of indirect taxes, since these accrue to the government and do not constitute compensation to employees. A measure of labour share adjusted for the effects of taxes is then constructed as equation (17), but now the denominator is total GDP less indirect taxes less subsidies on factors of production and on products.

Next, an adjustment should be made to account for the remuneration of the self-employed. Given available data, the income of non-farm unincorporated businesses can be added to the numerator of equation (17) in order to account for that part of the remuneration of the self-employed that constitutes a return to labour rather than to capital. These two effects can be accounted for jointly, yielding a measure of labour share adjusted for taxes and self-employment.

8. Detailed derivations of the formulas are presented in Guay and Luger (2002).

The first three measures of labour share just described are constructed using income-based GDP from the National Income and Expenditure Accounts, which is measured at market prices. Alternatively, measures based on factor costs can be considered in determining the labour share. These can be constructed by using the wages and salaries disaggregated by industry, and the GDP at factor cost also at the industry level. By using industry-level data, sectors where the theory does not apply can be excluded from the measures of the labour share. For example, the public sector can be excluded, since the concepts of labour and capital shares arguably apply only to the market sector of the economy. The farm sector can also be excluded because of the very large subsidies that farmers receive. This preferred measure is then constructed as

lshare_{nfb} = (all industries, wages and salaries – farm wages and salaries – public wages and salaries)/(all industries GDP – farm GDP – public GDP).

The levels of the different measures of the labour share are shown in Figure 2, where they seem to move in a similar fashion over the sample period. Figures 3 and 4 show each measure in percentage deviation from its mean, together with the inflation series (based on the GDP deflator).

It becomes clear from these figures that the various measures of the labour share have very different relationships to inflation. Dynamic cross-correlations are presented in Figures 5 and 6, where it is obvious that the fourth measure described above is potentially the most promising as the explanatory variable for Canadian inflation.

Note how the third measure co-moves negatively with most leads and lags of inflation. Figure 7 shows the dynamic cross-correlations between inflation and taxes less subsidies on factors of production and on products (top panel), and between inflation and income of non-farm unincorporated businesses (bottom panel).

The strong negative co-movements seen in these figures explain why the third measure adjusted for taxes and self-employment is inconsistent with the new Phillips curve. One possible explanation for the negative co-movement between taxes (less subsidies) and leads and lags of inflation is that the period of high oil prices in the 1970s and early 1980s was also accompanied by high subsidies on imported oil. On the other hand, the negative co-movements between income of the self-employed and leads and lags of inflation might simply be due to the substantial upward trend in self-employment vis-à-vis the downward movements in inflation. Finally, the autocorrelation functions of the different measures of the labour share are

Figure 2
Alternative measures of the labour share

presented in Figures 8 and 9. The fourth measure displays the strongest persistence, a well-known feature of the inflation process.

4 Results for Canada

4.1 Baseline model estimates

We first present estimates for the reduced form of the NKPC equation (5) given by:

$$\pi_t = \kappa\lambda mc_t + \beta E_t \pi_{t+1},$$

where $\kappa = 1/(1-\eta\mu)$. If one follows Yun (1996) and Goodfriend and King (1997), then $\kappa = 1$, whereas following Sbordone (2001), $\kappa = \frac{\alpha}{(1-\alpha)}\mu$.

This reduced-form specification is estimated over the sample period 1970Q1–2000Q4. As mentioned, inflation is based on the GDP deflator, and mc_t is the real marginal cost in log-deviation from its mean, calculated as the labour share of non-farm business. Several sets of instruments are used to investigate the robustness of the estimation results. They are: [1] two lags of inflation and real marginal cost, [2] three lags of inflation and real marginal cost, [3] four lags of inflation and real marginal cost, and [4] four

Figure 3
Inflation and different measures of the labour share

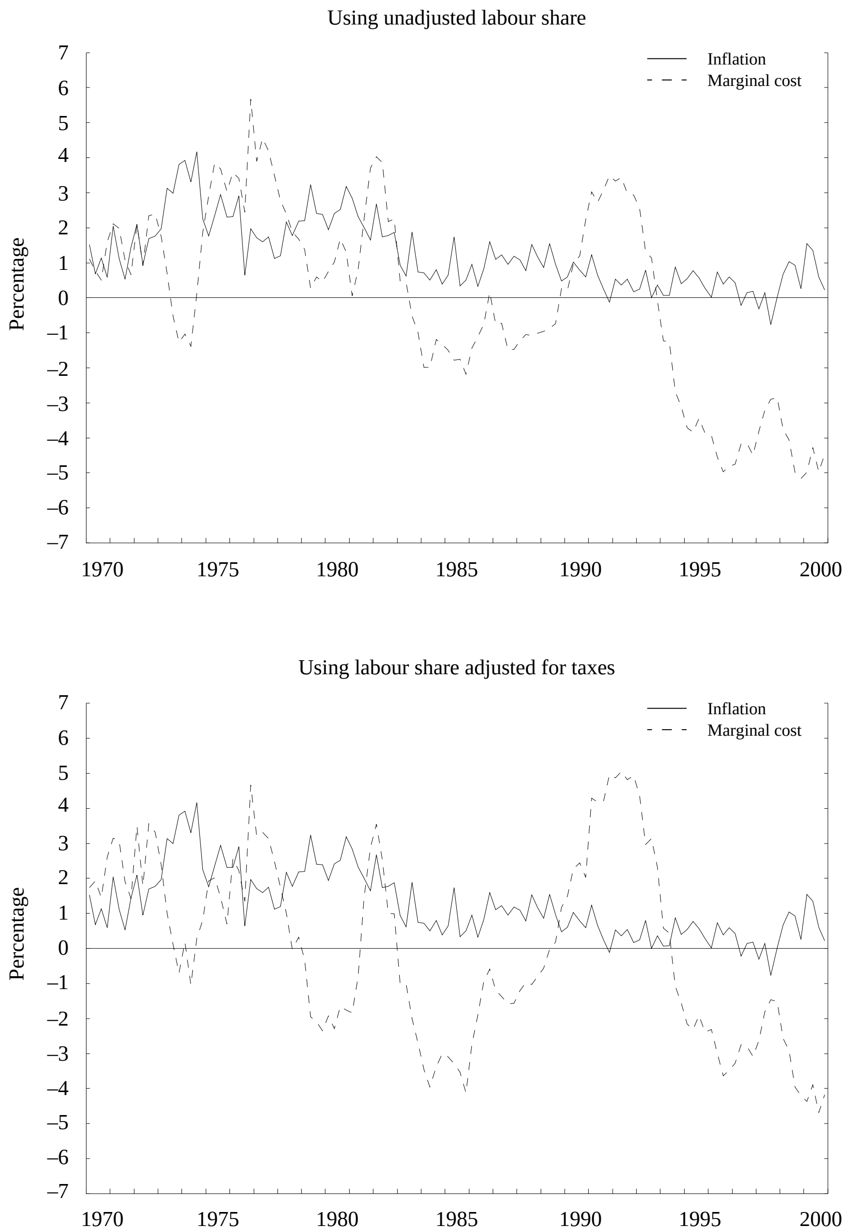


Figure 4
Inflation and different measures of the labour share

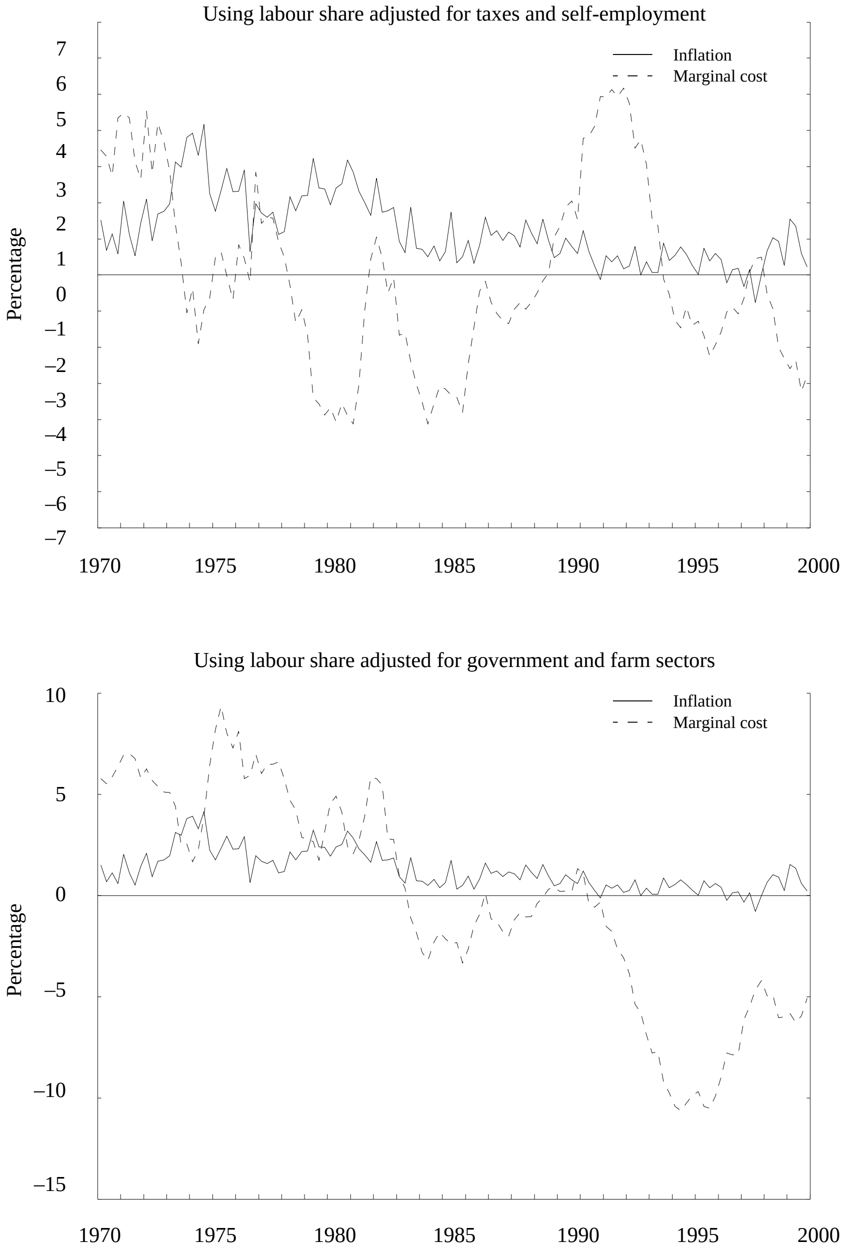


Figure 5
Dynamic cross-correlations: 1970 to 2000

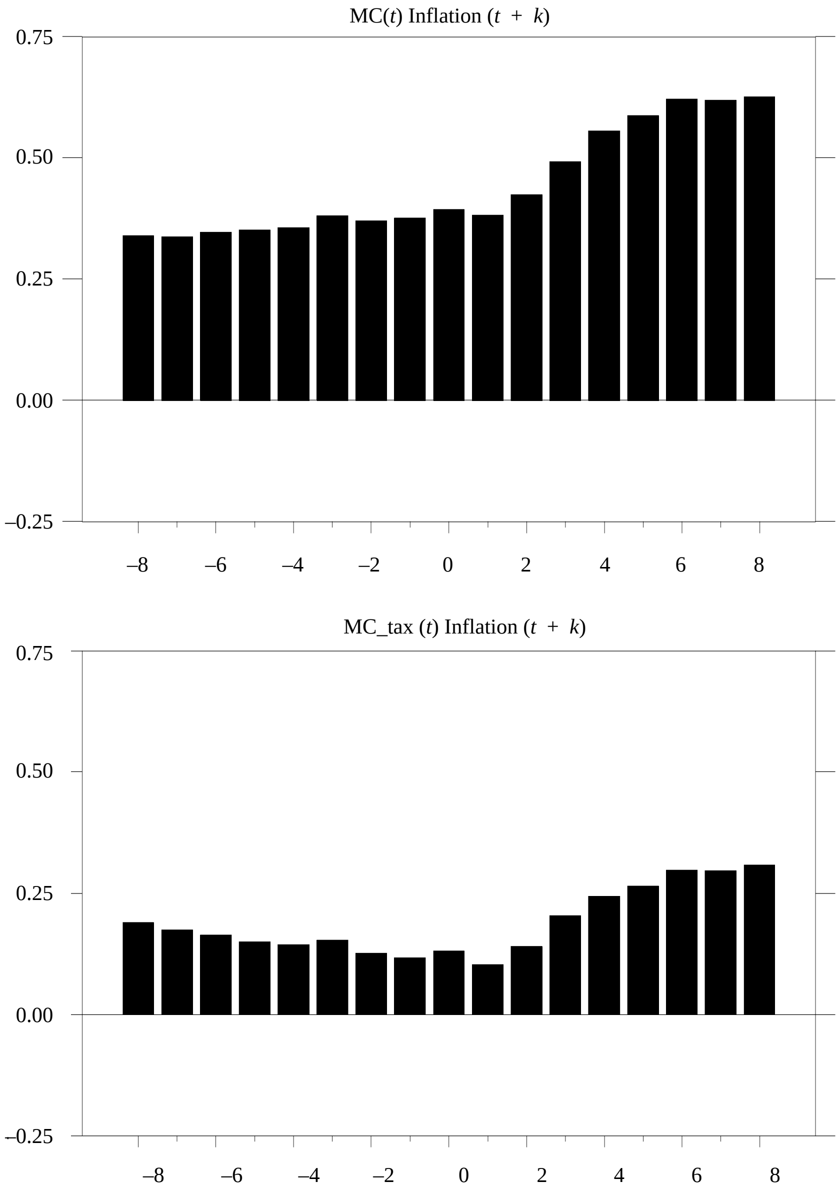


Figure 6
Dynamic cross-correlations: 1970 to 2000

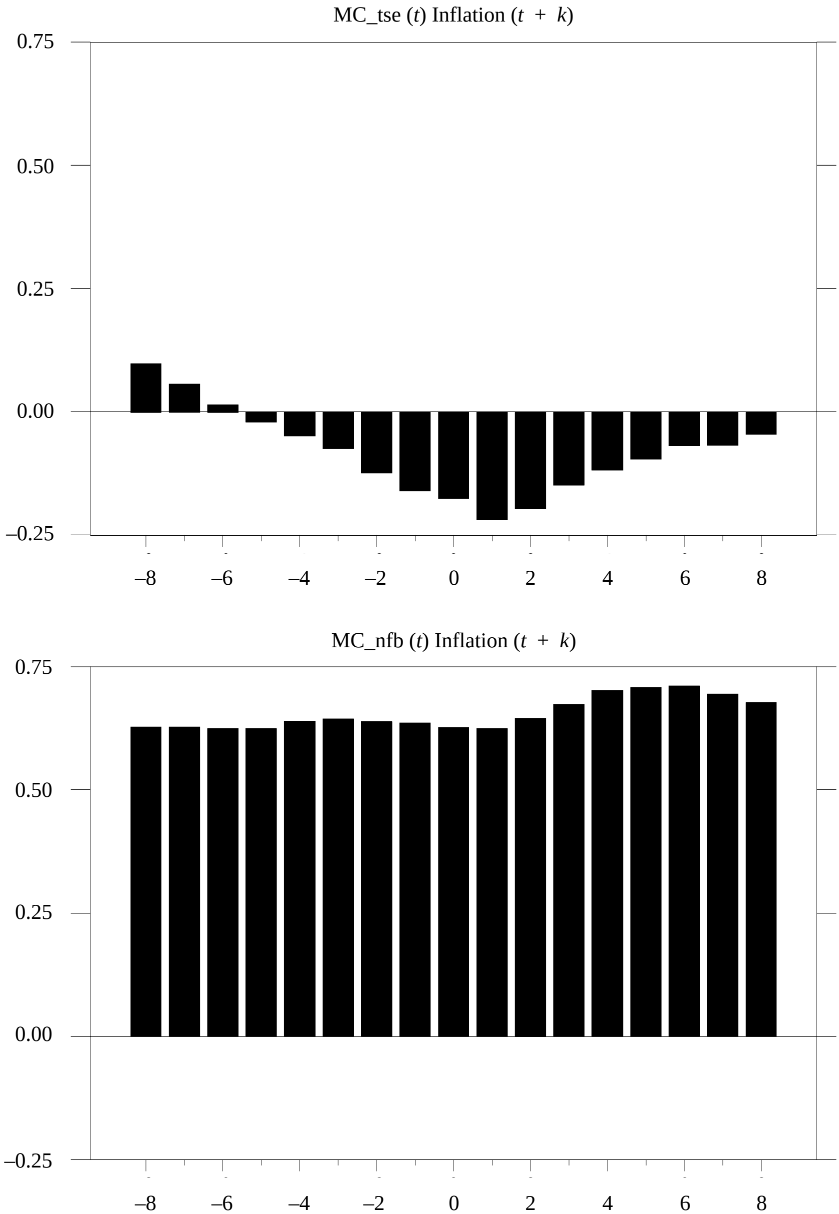


Figure 7
Dynamic cross-correlations: 1970 to 2000

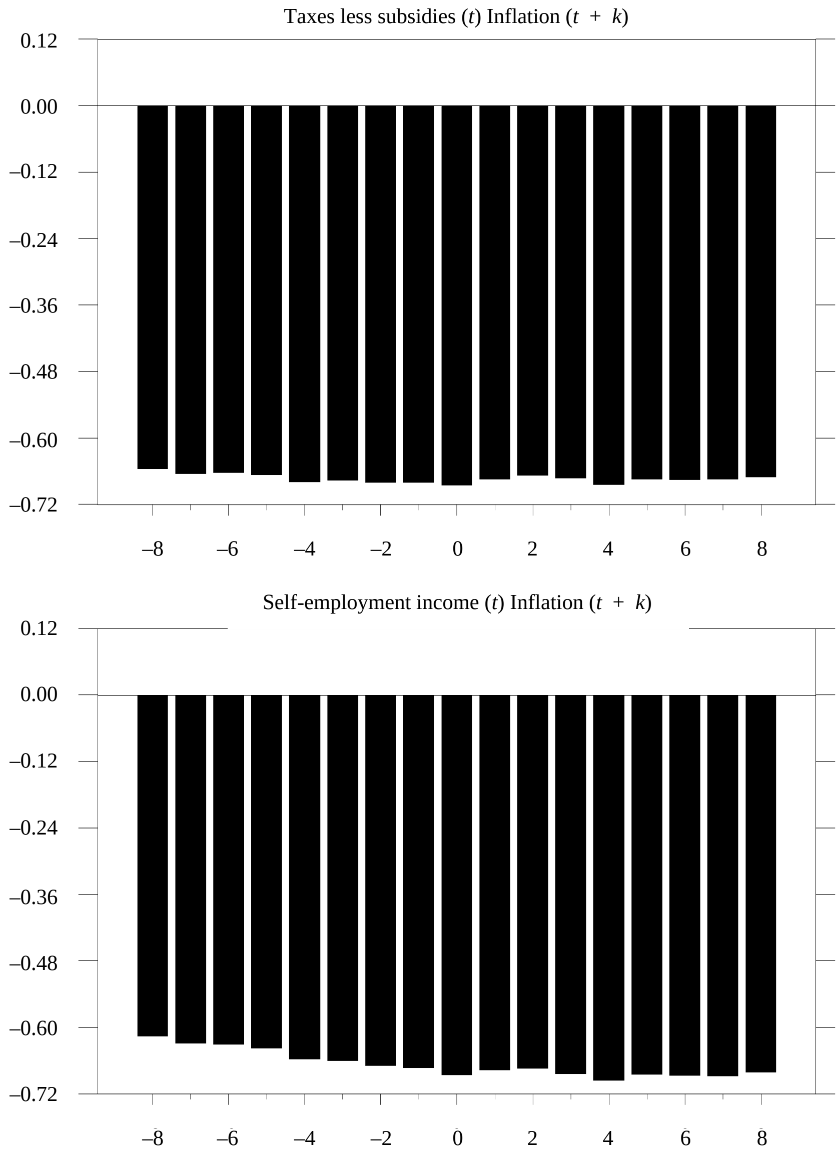


Figure 8
Autocorrelation function

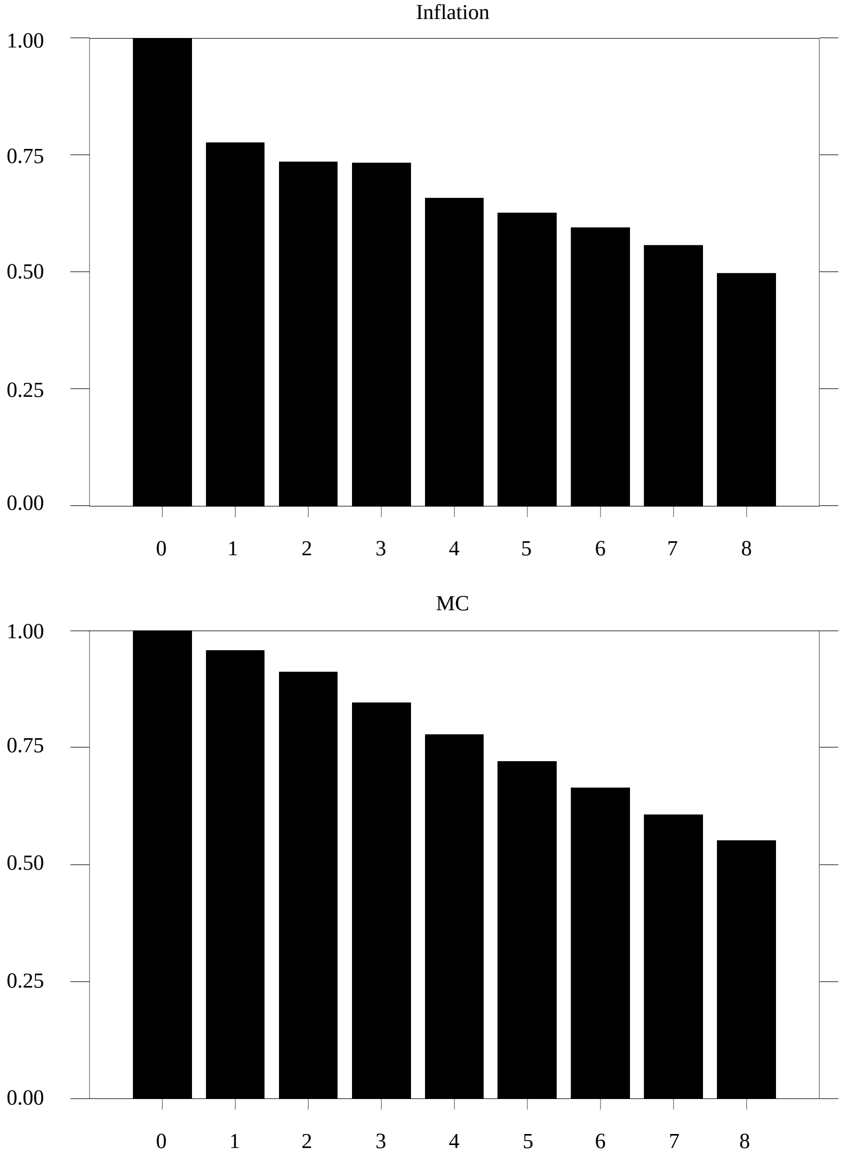
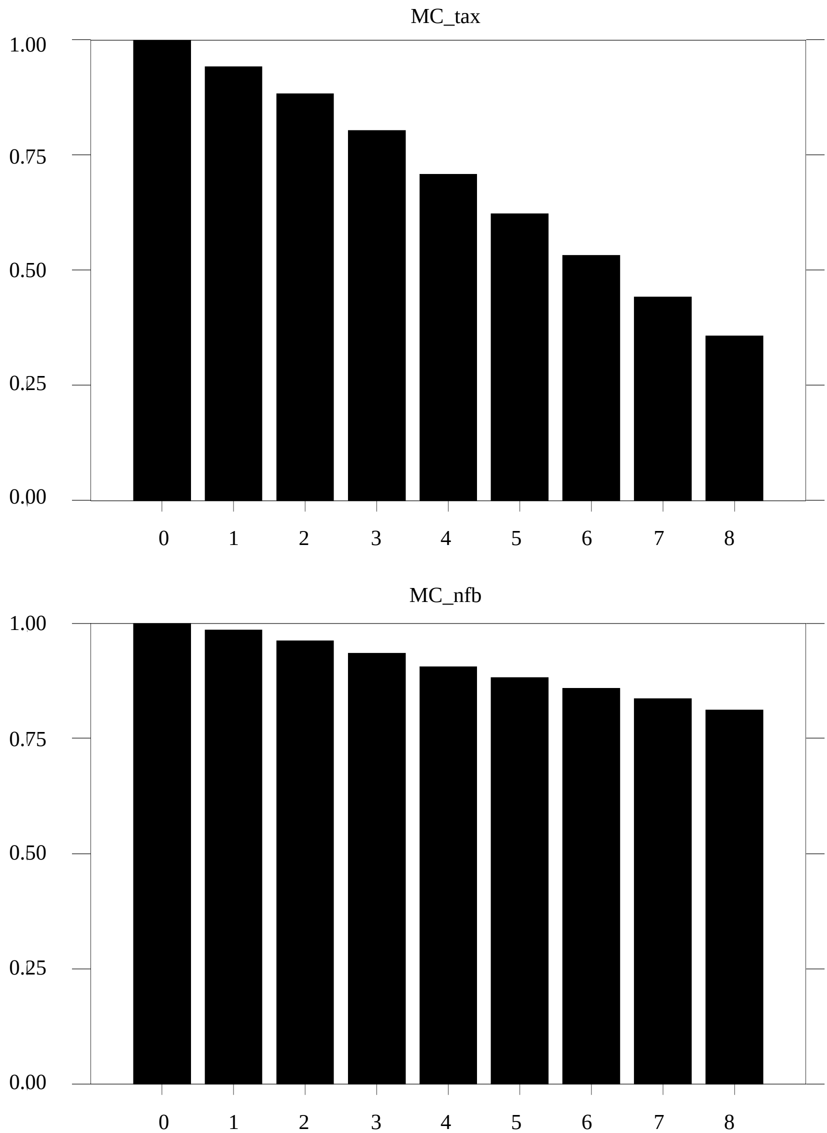


Figure 9
Autocorrelation function



lags of inflation, real marginal cost, and nominal wages. Instruments dated $t-1$ and earlier are used to mitigate possible correlations with the measurement error of real marginal cost.

We depart from earlier studies by excluding output-gap measures from the instrument sets. Two measures of output gap are usually retained as instruments. One is based on quadratically detrended output. With standard unit-root tests (such as the Augmented Dickey-Fuller), the presence of a unit root in Canadian output cannot be rejected. Under the maintained hypothesis of a unit root, quadratically detrended output is then also characterized by a unit root. Unfortunately, the asymptotic properties of IV estimators in the presence of non-stationary instruments are not known. As a result, usual inference procedures are likely to be invalid. The other measure of output gap usually used is based on the Hodrick-Prescott filter. The output gap is then a combination of lags, leads, and contemporaneous values of output. Such measures of the output gap violate the basic GMM orthogonality conditions and are likely to be correlated with the measurement error of real marginal cost.

The GMM estimator for this linear specification corresponds to the IV estimator (two-stage least squares), which we correct for bias, as explained above. We also use a heteroscedasticity and autocorrelation-consistent matrix estimator for the sample moments in deviations from the mean to increase the power of the overidentifying restrictions test, as suggested by Hall (2000) and Bonnal and Renault (2001). The automatic lags selection procedure proposed by Newey and West (1994) is adopted. Table 2 reports the results for $\kappa = 0.13$. This value is proposed by Gagnon and Khan (2001) for Canada following the assumptions in Sbordone (2001). It is important to understand that the inference results based on the reduced form do not depend on κ. While the scaling of the parameter λ depends on κ, its statistical significance does not, since the value of κ is a fixed constant that cancels out from the *t*-statistic.

The results are not encouraging for the NKPC. The slope coefficient on marginal cost is never significant whatever the set of instruments. For three cases, the coefficient has the wrong sign, and the discount factor is greater than one in all cases. Finally, the overidentifying restrictions are rejected with the instrument sets, which include four lags of inflation and real marginal cost. It appears that the New Phillips curve is misspecified, and richer dynamics would seem necessary to capture the persistence of Canadian inflation.

Table 2
Reduced-form estimates

Instrument set	β	λ	*J*-stat.
[1]	1.017	–0.0024	2.83
	(0.000)	(0.947)	(0.419)
[2]	1.010	–0.0084	8.03
	(0.000)	(0.803)	(0.155)
[3]	1.037	0.0010	23.20
	(0.000)	(0.977)	(0.002)
[4]	1.011	–0.0010	26.97
	(0.000)	(0.742)	(0.005)

Note: The *p*-values (in parentheses) corresponding to the estimates of β and λ are for the null hypotheses that these parameters are zero.

4.2 Hybrid model estimates

Estimation of the hybrid specification is based on the following structural form:

$$\pi_t = \lambda\left(\frac{1}{(1-\eta\mu)}\right)mc_t + \gamma_f E_t\pi_{t+1} + \gamma_b\pi_{t-1},$$

where

$$\lambda = \left(\frac{(1-\omega)(1-\theta)(1-\theta\beta)}{\theta}\right)\phi^{-1},$$

$$\gamma_f = \beta\theta\phi^{-1},$$

$$\gamma_b = \omega\phi^{-1},$$

$$\phi = \theta + \omega[1-\theta(1-\beta)].$$

We consider three estimators. First, the corresponding reduced form is estimated by instrumental variables (a two-stage least squares), and structural parameter estimates are then derived from the reduced-form estimates. Second, GMM estimation is performed based on the following orthogonality conditions:

$$E_t\left[\left(\pi_t - \beta\theta\phi^{-1}\pi_{t+1} - \left(\frac{(1-\omega)(1-\theta)(1-\theta\beta)}{\theta}\right)\phi^{-1}\kappa mc_t - \omega\phi^{-1}\pi_{t-1}\right)Z_t\right] = 0.$$

Note that GMM estimation results from the same orthogonality conditions depend on the chosen normalization (Galí and Gertler 1999). On the other hand, the CUE, by construction, is invariant to the choice of normalization. Following Newey and Smith (2001), analytical bias-corrected versions of these three estimators can be computed.

Based on instrument set [4], Table 3 reports reduced-form and structural parameter estimates setting $\kappa = 1$ and $\kappa = 0.13$. Also reported is the average price duration, *D* (in quarters), corresponding to the estimate of θ.

The estimates are fairly similar across methods of estimation. Forward-looking behaviour is dominant relative to the backward-looking component. The fraction of the backward-looking price-setters differs from zero and is near one-third. The discount factor is still greater than one across specifications and estimation methods. The estimates of the probability of changing price imply an unrealistic duration of price stickiness. The duration lies between 12 quarters and a value as high as 48 quarters. The slope coefficient on marginal cost now has the right sign in all cases. However, it is never significantly different from zero. Finally, the overidentifying restrictions are not rejected, but only marginally.[9] Thus, the results based on the Cobb-Douglas production technology suggest that the real marginal cost is not a significant determinant of inflation, which refutes the theoretical predictions. These results stand in contrast to Gagnon and Khan (2001), who find evidence supporting the New Phillips curve in Canada. In particular, they never reject the hybrid specification.

Table 4 reports the results for a Cobb-Douglas production function with overhead labour. In this case, the real marginal cost is given by:

$$mc_t = s_t + bh_t,$$

where

$$b = \frac{\bar{H}/H}{1 - \bar{H}/H}.$$

The series for hours worked is constructed as the number of employees multiplied by the average hours worked per quarter.[10] The resulting series is stationary around a stable mean. In contrast with the series used by Sbordone (2001) and Gagnon and Khan (2001), no detrending is needed.

9. The overidentifying restrictions test rejects for all cases with the instrument set [3] at the usual 5 per cent level.

10. The average hours worked per quarter are calculated by multiplying the average hours worked per week by 13.

Table 3
Hybrid Phillips curve estimates: Cobb-Douglas

κ	Method	θ	β	λ	ω	γ_f	γ_b	D	*J*-stat.
1	IV	0.951	1.008	0.0004	0.337	0.743	0.261	20.53	17.09
		(0.041)	(0.022)	(0.003)	(0.092)	(0.055)	(0.055)	(17.44)	
		[0.000]	[0.000]	[0.871]	[0.000]	[0.000]	[0.000]	[0.242]	[0.072]
	GMM	0.969	1.001	0.0004	0.389	0.713	0.287	32.01	17.28
		(0.081)	(0.025)	(0.002)	(0.106)	(0.057)	(0.056)	(82.72)	
		[0.000]	[0.000]	[0.853]	[0.000]	[0.000]	[0.000]	[0.700]	[0.068]
	CUE	0.979	1.002	0.0002	0.398	0.713	0.287	48.66	17.11
		(0.250)	(0.040)	(0.005)	(0.120)	(0.068)	(0.065)	(79.15)	
		[0.000]	[0.000]	[0.967]	[0.001]	[0.000]	[0.000]	[0.934]	[0.072]
0.13	IV	0.913	1.007	0.0033	0.323	0.743	0.261	11.40	17.10
		(0.020)	(0.022)	(0.020)	(0.104)	(0.055)	(0.055)	(26.19)	
		[0.000]	[0.000]	[0.870]	[0.004]	[0.000]	[0.000]	[0.664]	[0.072]
	GMM	0.917	1.001	0.0034	0.368	0.713	0.287	12.03	17.28
		(0.201)	(0.024)	(0.020)	(0.119)	(0.056)	(0.056)	(29.57)	
		[0.000]	[0.000]	[0.850]	[0.003]	[0.000]	[0.000]	[0.685]	[0.068]
	CUE	0.928	1.002	0.0024	0.376	0.712	0.288	14.04	17.11
		(0.471)	(0.036)	(0.003)	(0.167)	(0.071)	(0.066)	(93.01)	
		[0.051]	[0.000]	[0.944]	[0.025]	[0.000]	[0.000]	[0.880]	[0.072]

Notes: Standard errors are in parentheses; *p*-values are in square brackets.

Table 4
Hybrid Phillips curve estimates: Cobb-Douglas and overhead labour

κ	Method	θ	β	λ	ω	γ_f	γ_b	D	*J*-stat.
1	GMM	0.879	1.001	0.0056	0.478	0.645	0.352	8.26	28.59
		(0.029)	(0.039)	(0.004)	(0.078)	(0.045)	(0.040)	(2.01)	
		[0.000]	[0.000]	[0.146]	[0.000]	[0.000]	[0.000]	[0.000]	[0.012]
	CUE	0.858	0.926	0.010	0.533	0.585	0.393	7.02	25.26
		(0.062)	(0.159)	(0.012)	(0.127)	(0.109)	(0.078)	(3.05)	
		[0.000]	[0.000]	[0.422]	[0.000]	[0.000]	[0.000]	[0.022]	[0.029]
0.13	GMM	0.718	1.001	0.044	0.391	0.648	0.352	3.548	28.59
		(0.068)	(0.033)	(0.029)	(0.066)	(0.044)	(0.040)	(0.853)	
		[0.000]	[0.000]	[0.138]	[0.000]	[0.000]	[0.000]	[0.000]	[0.012]
	CUE	0.607	0.902	0.122	0.355	0.582	0.377	2.54	23.93
		(0.113)	(0.126)	(0.099)	(0.084)	(0.116)	(0.080)	(0.730)	
		[0.000]	[0.000]	[0.224]	[0.000]	[0.000]	[0.000]	[0.001]	[0.047]

Notes: Standard errors are in parentheses; *p*-values are in square brackets.

Finally, the instrument set corresponds to the fourth one augmented by four lags of hours worked.

We first tried to estimate the parameter *b*. Unfortunately, the estimates were never significant. Instead, we report the results obtained with GMM and CUE for a value of *b*, calibrated as in other empirical studies of the New Phillips curve. Following Rotemberg and Woodford (1999), *b* is calibrated to 0.4. The estimates of price-stickiness duration are now more plausible, especially for $\kappa = 0.13$. The discount factor is now less than one in the case of estimation by CUE. The forward-looking component still dominates, but now the fraction of backward-looking price-setters is more important for $\kappa = 1$. Here again, the slope coefficient on marginal is never significant, and the specification is rejected in all cases.

We also tried to estimate the specification including the adjustment cost of labour based on equation (6). Labour is measured as explained above and the instrument set is the same as for estimation of the previous model specification. The parameter ϕ was never significant across estimation methods. We then calibrated this parameter following the estimates reported by Ambler, Guay, and Phaneuf (1999). The estimation results were not significantly different from the ones based on a Cobb-Douglas production function without adjustment cost.

4.3 Open economy

We proceed to estimate the hybrid specification with real marginal cost augmented by the terms of trade as in equation (7). The terms of trade are calculated as the logarithm of the import price minus the logarithm of the domestic GDP deflator. Instrument set [4] augmented with four lags of the terms of trade is used.

In this case as well, the relative coefficient to the terms of trade is never significant. Table 5 reports estimates obtained with CUE for $\kappa = 0.13$. The parameter of openness, Θ, has a point estimate, which seems to be in accordance with the degree of openness of the Canadian economy, but it is not statistically significant. The other results are similar to previous ones and again the specification is decisively rejected. The addition of the terms of trade in the instrument set results in an important increase in the value of the overidentifying restrictions test statistic. This suggests that the terms of trade could be an important explanatory variable for Canadian inflation.

The results are similar when we replace the terms of trade in the specification of real marginal cost by the real exchange rate. The parameter relative to the real exchange rate is estimated imprecisely, and the overidentifying restrictions test clearly rejects the model in all cases.

Table 5
Hybrid Phillips curve estimates: Open economy with terms of trade

Method	θ	β	λ	ω	Θ	γ_f	γ_b	D	*J*-stat.
CUE	0.939	0.999	0.0011	0.528	0.303	0.640	0.359	16.49	35.86
	(0.467)	(0.066)	(0.019)	(0.254)	(9.02)	(0.041)	(0.030)	(127.20)	
	[0.047]	[0.000]	[0.952]	[0.040]	[0.999]	[0.000]	[0.000]	[0.897]	[0.000]

Notes: Standard errors are in parentheses; *p*-values are in square brackets.

5 Discussion

The estimation strategy advocated in this paper allows us to obtain estimates of New Phillips curves, which do not depend on the normalization of the moment conditions. Furthermore, the implementation of bias-corrected estimators mitigates the well-known problem in IV methods of a bias effect that increases with the number of moment conditions. When applied to Canadian data, the bias-corrected estimator results in more importance being given to the forward-looking relative to the backward-looking component in the hybrid version New Phillips curve compared with the GMM estimates obtained by Gagnon and Khan (2001).

In contrast with other empirical studies,[11] the specification test based on overidentifying restrictions rejects the New Phillips curve and its hybrid version for almost all specifications considered in this paper. The estimation of the weighting matrix is crucial for the small sample properties of Hansen's (1982) specification test, especially when the number of moment conditions is important relative to the number of observations.[12] These studies fixed at arbitrary values the number of lags used in kernel estimation of the weighting matrix. In this paper, we adopt a data-dependent automatic lag selection procedure, and the estimation of the weighting matrix is based on sample moments in deviation. This approach improves the power of the overidentifying restrictions test in small samples.[13]

11. Balakrishnan and López-Salido (2002); Gagnon and Khan (2001); Galí and Gertler (1999); Galí, Gertler, and López-Salido (2001a); Galí, Gertler, and López-Salido (2001b); and Galí and López-Salido (2001).
12. For some of these studies, the ratio of the number of moment conditions to the number of observations equals one-third.
13. Similar remarks hold for the econometric investigation of the New Phillips curve for U.S. inflation (see Guay and Luger 2002).

Conclusions

The rejection of alternative specifications of the New Phillips curve suggests that a richer dynamic structure in the explanatory variables will be needed to capture the dynamics of Canadian inflation. In the case of the United States, Kurmann (2002) also finds considerable uncertainty between the observed persistent movements in inflation and what is predicted by a New Phillips curve model. His results and those of this paper represent an important step back from the conclusions of previous authors who argue that New Phillips curve models are a good representation of inflation dynamics. These new results suggest that, at the theoretical level, richer versions of the structural model from which the New Phillips curve is derived would need to be developed. Mankiw and Reis (2002) proposed a "sticky-information"-based Phillips curve that can generate inflation dynamics similar to what is observed in the data. However, assessing the empirical relevance of that model raises several other econometric issues that go beyond the scope of this paper.[14]

14. Khan and Zhu (2002) report estimates of the "sticky-information"-based Phillips curve. However, their inference method suffers from a generated regressor problem of the type mentioned above.

Appendix
Data Description

When applicable, the final series is in quarterly frequency, seasonally adjusted, at annual rates, and in millions (of dollars or persons), unless otherwise indicated. The series codes are from Statistics Canada's CANSIM database.

1. Total GDP deflator: constructed from the following series.
   - Nominal GDP = V498086
   - Constant dollar GDP = V1992259
   - Chained dollar GDP = V1992067

2. Labour income share
   - lshare = wages and salaries/total GDP = V498076/V498074
   - $lshare_{tax}$ = wages and salaries/(total GDP – indirect taxes less subsidies on factors of production and on products)
     = V498076/(V498074 – V1992216 – V1997473)
   - $lshare_{tse}$ = (wages and salaries + income of non-farm unincorporated business)/(total GDP – indirect taxes less subsidies on factors of production and on products)
     = (V498076 + V498081)/(V498074 – V1992216 – V1997473)
   - $lshare_{nfb}$ = (all industries wages and salaries – farm wages and salaries – public wages and salaries)/(all industries GDP – farm GDP – public GDP)

Note: $lshare_{nfb}$ is constructed from CANSIM Table 379–0006 (GDP at factor cost), 382–0001 (old table for wages and salaries), and 382–0006 (new table for wages and salaries).

3. Import prices: constructed from the following series.
   - Nominal imports = V498106
   - Constant dollar imports = V1992253
   - Chained dollar imports = V1992063

4. Hours worked
   - Average hours worked per week, all industries = LSA2050 (Bank of Canada series code)

5. Employment
    - Total employment, 15 years old and above = D767608 and V2062811
    - Private sector employment = total employment – V2066969
6. Population
    - Population, 15 years old and above = D767284 and V2091030
7. Nominal exchange rates
    - Canadian dollar/U.S. dollar closing rate = B3414 (monthly frequency, a quarter is the average of three months)

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