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Empirical Evidence on the Cost of Adjustment and Dynamic Labour Demand

by

Robert A. Amano

Bank of Canada Banque du Canada

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Robert A. Amano
Research Department
Bank of Canada
Ottawa, Ontario, CANADA K1A 0G9
Tel.: (613) 782-8827
Fax: (613) 782-7163
E-mail: bamano@bank-banque-canada.ca

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ABSTRACT

In this paper the author examines whether there is significant evidence of the effect of adjustment costs on Canadian labour demand. This is an important question, as sluggish adjustment of labour demand resulting from significant adjustment costs may be one factor that could help explain some of the unemployment persistence found in Canadian data.

The author uses a linear-quadratic model and attempts to estimate the relative adjustment costs of labour demand as well as its rate of adjustment towards long-run equilibrium. In contrast to others who have examined the dynamic behaviour of labour demand, the author estimates the structural parameters using the Euler equation and employs a limited-information approach that does not require an explicit solution for the model's control variables in terms of the forcing processes.

The empirical estimates imply that adjustment costs are about four times more important than disequilibrium costs and that it takes over three and a half years for 90 per cent of labour demand adjustment to be completed. Therefore the author concludes that significant adjustment costs are an important feature of Canadian labour demand and that sluggishness due to these costs may be one explanatory factor in unemployment persistence.

RÉSUMÉ

Dans le présent document, l'auteur cherche à établir s'il existe des preuves empiriques significatives confortant l'hypothèse que les coûts d'ajustement de la demande de travail au Canada agissent sur cette dernière. C'est là une importante question, car si cette hypothèse se confirmait, cela pourrait contribuer à expliquer en partie la persistance du chômage relevée dans les données canadiennes.

À l'aide d'un modèle quadratique linéaire, l'auteur tente d'estimer les coûts d'ajustement relatifs de la demande de travail aussi bien que le rythme d'ajustement de cette dernière vers l'équilibre à long terme. Contrairement aux autres chercheurs qui se sont intéressés au comportement dynamique de la demande de travail, l'auteur estime les paramètres structurels à l'aide de l'équation d'Euler et d'une méthode du maximum de vraisemblance à information limitée qui n'exige pas que les variables de contrôle du modèle soient explicitement résolues en fonction des variables d'impulsion exogènes.

Les estimations empiriques impliquent que les coûts d'ajustement vers l'équilibre sont environ quatre fois plus élevés que les coûts obtenus en l'absence d'équilibre et que l'ajustement de la demande de travail s'achève dans une proportion de 90 % au bout d'un peu plus de trois ans et demi. Par conséquent, l'auteur conclut que les coûts d'ajustement élevés de la demande de travail au Canada constituent une importante caractéristique de celle-ci et que la lenteur d'ajustement due à ces coûts est une cause possible de la persistance du chômage.

1 INTRODUCTION

Perhaps one of the most important Canadian macroeconomic problems for the 1990s is the relatively high and persistent rate of unemployment. In response, there have been calls for policy actions to help remedy this situation. However, from a policy perspective it is important to first identify factors that may help explain such unemployment persistence. Numerous studies have focussed on “supply-side” factors such as mismatching, sectoral shifts, demographics, displacement behaviour and unemployment insurance (see Poloz 1994, for a review of the Canadian evidence). In marked contrast, relatively little attention has been paid to sluggish adjustment of labour demand as a potential reason for Canadian unemployment persistence.

There are several reasons for expecting labour demand to be slow to adjust. The most obvious reason is that it is costly to hire and fire employees.[1] Hiring costs include expenditures on items such as advertising, time spent on interviewing and training, and the loss of output while the new employee learns the job. Oi (1962) finds that hiring costs at the International Harvester Company amounted to about 142 hours’ pay at the company’s average hourly rate. Downsizing costs include items such as severance payments and the effect on productivity of a decline in the morale of the remaining employees. Nickell (1979) and Burgess (1988) link slow adjustment of labour to the imposition of a variety of labour market policies that make it more difficult for firms to shed labour. Other reasons that may give rise to slow labour adjustment are those associated with the institutional structure of the economy. These would include distortions such as unionization and benefits a firm must pay for new employees (such as unemployment insurance, pension and workers’ compensation contributions, and other fringe benefits) over the cost of overtime for its current labour force. In short, there are several factors that could make labour costly

1. See Nickell (1986) for a lucid description of the size and structure of adjustment costs in labour demand.

to adjust.

Although many papers have allowed for sluggish labour demand adjustment, only a relatively small number focus on it. Previous work on labour demand as a factor in unemployment persistence has concentrated on experiences in the European countries and the United States (see, for example, Alogoskoufis and Manning 1988). Not surprisingly, most of this research finds evidence of a significant degree of sluggishness in the European countries but not in the United States. A study of Canadian labour demand is particularly interesting, as it represents an intermediate case between Europe and the United States in terms of unemployment experience, and the Canadian data appear to be little explored. Significant evidence of high adjustment costs may help account for some of the persistence in Canadian unemployment. Such evidence would suggest that policies that lead to the reduction of real wages or stimulation of demand will not yield large employment gains in the short run. For instance, conjectures that monetary policy should generate a "little" inflation to lower real wages in order to quickly adjust labour demand would be drawn into question (see Summers 1991).

In this paper, I take a preliminary empirical look at this important labour demand question. More specifically, the properties of the linear-quadratic (LQ) model with integrated processes are exploited to obtain a measure of the relative importance of disequilibrium to adjustment costs as well as the rate of adjustment toward equilibrium. The LQ framework is a popular framework for investigating the dynamic behaviour of economic agents. The model has been used to explain, *inter alia*, the demand and supply for money (Amano and Wirjanto 1993, 1995), the demand and supply for labour (Kennan 1988a), the demand for imports (Amano and Wirjanto 1994), the demand for labour and capital (Meese 1980), private consumption (Laxton and Tetlow 1992) and price adjustment (Cozier 1989). The LQ framework has also been used in large-scale macroeconomic models to model dynamic behaviour (see, for example, Coletti et al. 1995).

The popularity of the LQ model likely reflects the fact that it has several attractive features. In particular, it allows for a wide range of dynamic behaviour. In the context of labour demand, the LQ model, despite its simplicity, encompasses a class of dynamic models often used by researchers (see Gregory, Pagan and Smith 1990). Examples include the standard partial-adjustment and error-correction models estimated by Rosen (1968), Topel (1982) and Jenkinson (1986), among others.

Second, the LQ model gives rise to linear decision rules in the variables. This is a particularly convenient feature from the perspective of empirical labour demand, since the variables often used in labour demand studies appear to be characterized by nonstationary processes, and the LQ model has well-understood properties for these nonstationary variables (see Gregory, Pagan and Smith 1990, henceforth GPS). As shown later, this assumption allows me to exploit the recently developed theory of cointegration to estimate the parameters of the Euler equation. Previous research examining labour demand in an LQ framework has tended to assume that the variables under consideration are stationary (see, for example, Sargent 1978 and Meese 1980).[2] Our unit-root and stationarity tests suggest that this assumption may be a source of misspecification.

Third, the LQ approach requires a minimization of the expected discounted present value of an expression that includes quadratic adjustment costs. The LQ model also incorporates forward-looking elements into the decision process and in addition provides some microeconomic foundations for aggregate dynamics. Much of the existing literature that examines sluggish adjustment of labour demand does not incorporate these important elements into the analysis (see, for example, Alogoskoufis and Manning 1988).

The LQ model is estimated using an Euler equation approach. This approach is based on the insight that the optimization problem is recursive so that the same time path

2. It should be noted that these papers were written before the ready application of unit-root and cointegration tests appeared in the economic literature.

for labour demand is chosen whether the decision is based on the whole future stream of the expected marginal product of labour or on an optimal trade-off of labour demand between this period and next period. Since the informational requirements of predicting the next period are far less stringent than those for predicting the entire future, the Euler equation approach substantially eases the econometrician's task.

The paper is structured as follows. Section 2 describes the linear-quadratic model and derives some of its implications. The estimation strategy is outlined in Section 3, while the empirical results are given in Section 4. Section 5 concludes the paper.

2 THE LINEAR-QUADRATIC MODEL

This section describes the LQ model and derives some of its implications in the nonstationary context. We generalize the static formulation by assuming that labour demand is set to minimize, subject to rational expectations, an intertemporal loss function with quadratic costs of adjustment. One may view these structures as a result of aggregation over firms in the economy or, alternatively, as providing local linearizations of the first-order conditions. The firm is assumed to control the level of labour (n_t) and solve the problem of minimizing the expected present value of adjustment and disequilibrium costs:

$$\min_{\{n_i\}} \mathrm{E}_t \sum_{i=t}^{\infty} \beta^{i-t} [\gamma (n_i - n_i^*)^2 + (n_i - n_{i-1})^2] \tag{1}$$

for $i \geq t$, subject to a law of motion between the target level of labour, n^*, and some observable economic variables. In equation (1), E_t is the expectations operator conditional on the firm's information at time t (I_t), $\beta \in (0, 1)$ is the subjective discount factor, and the parameter $\gamma > 0$ is a weighting factor that determines the relative importance of disequilibrium to adjustment costs. Note that γ is the inverse of the usual cost of adjustment. Adjustment costs correspond to those mentioned in the introduction, whereas disequilibrium costs arise when firms are not at their optimal levels of employment.

In general, we assume that the following law of motion for the target variable holds

$$n_t^* = X'_t \alpha + v_t \tag{2}$$

where v_t is a white noise process known to the firms, that is $v_t \in I_t$, but unknown to the econometrician whose information set is $H_t \subset I_t$, X_t is a (kx1) row vector of forcing variables and α is a *(k*x1*)* column vector of unknown parameters.

The necessary first-order condition for the minimization of (1) is given by the following Euler equation:

$$\Delta n_t = \beta E_t \Delta n_{t+1} - \gamma (n_t - n_t^*) \tag{3}$$

and the corresponding transversality condition is

$$\lim_{T \to \infty} E_t [\beta^T \{\gamma (n_T - n_T^*) + \Delta n_T\}] = 0 \tag{4}$$

The forward solution to (3) is given by

$$n_t = \lambda n_{t-1} + (1-\lambda)(1-\beta\lambda) E_t \sum_{i=t}^{\infty} (\beta\lambda)^{i-t} n_i^* \tag{5}$$

where $\lambda < \beta^{-1/2}$ is the smallest stable root of the Euler equation obtained from the first-order condition and satisfies the condition

$$\beta\lambda^2 - (1+\beta+\gamma)\lambda + 1 = 0 \tag{6}$$

Notice that λ also represents the so-called speed of adjustment to the target level of employment. In the following section, we present a methodology that allows us to consistently estimate this speed of adjustment term.

It follows from equations (2) and (5) that the control variable n_t will inherit any stochastic trends in the forcing variables. For the purpose of illustration, assume that X_t is an independent random walk, that is,

$$(1-L)X_t = e_t \tag{7}$$

where $E_{t-1}e_t = 0$. Substituting equation (2) into (5) and (7) yields

$$(1-\lambda L)n_t = (1-\lambda)X'_t\alpha + (1-\lambda)v_t \tag{8}$$

Since the root λ lies inside the unit circle, it follows from equation (8) that the endogenous variable n_t must be integrated of order one and that the white noise error term v_t is I(0). The latter implies that n_t and X_t are cointegrated with cointegrating vector $(1, \alpha)$. GPS show that similar results also hold when the forcing variables follow more complicated I(1) processes. This result implies that if the forcing variables follow I(1) processes, then the cointegration restriction between n_t and X_t is given by the LQ model.

To obtain an Euler equation that can be estimated, first substitute equation (2) into (3) to obtain

$$\Delta n_t = \beta E_t \Delta n_{t+1} - \gamma(n_t - X'_t\alpha) + \gamma v_t \tag{9}$$

and then replace $E_t\Delta n_{t+1}$ by its realization $(\Delta n_{t+1} + u_{t+1})$, where u_{t+1} is a purely expectational error, such that $E_t u_{t+1} = 0$, and rewrite equation (9) as

$$\Delta n_t = \beta\Delta n_{t+1} - \gamma(n_t - X'_t\alpha) + \eta_{t+1} \tag{10}$$

where $\eta_{t+1} = \beta u_{t+1} + \gamma v_t$, such that $E_t\eta_{t+1} = 0$. The disturbance, η_t, is thus a composite error term that can be rewritten as a first-order moving-average process, provided the structural error term v_t is a white noise process. Notice that equation (10) may be viewed as a "forward-looking" error-correction model. Since, as noted earlier, the LQ model implies that n_t and the forcing variables X_t are cointegrated (in the sense of Engle and Granger 1987), Dolado, Galbraith and Banerjee (1991), henceforth DGB, have suggested a two-step procedure for estimating the parameters in (10).

3 THE ESTIMATION STRATEGY

In this section, the estimation strategy for equation (10) and for the speed of adjustment term, λ are described. I begin with the former. In the first step, consistent estimates of the long-run parameter (α) may be obtained from a cointegrating regression

$$n_t = X'_t\alpha + \xi_t \tag{11}$$

where $\xi_t = (1-\lambda L)^{-1}[\lambda\gamma v_t - \lambda\alpha e_t]$. Notice that since the smallest stable root λ satisfies the condition in (6), as the adjustment cost gets large (that is, γ becomes small), the stable root approaches unity and ξ_t is nearly integrated and hence highly persistent. This implies that if adjustment costs are high, then evidence consistent with cointegration will be difficult to detect. Note also that since any bias in the least-squares (LS) estimates of equation (11) are super-consistent, it is possible to substitute these estimates into equation (10) and ignore any sampling uncertainty in the estimate of α when the remaining parameters in the Euler equation are estimated (see Stock 1987).[3]

However, it is important to note that the super-consistency property does not, by itself, ensure that the estimates of α will have good finite-sample properties (see Banerjee et al. 1986). This is due to the fact that the LS estimates of α are not asymptotically efficient. The inefficiency arises because the asymptotic distribution depends on nuisance parameters because of serial correlation in the error term and because of the endogeneity of the regressor matrix X_t induced by Granger-causation from innovations in n_t to

3. Super-consistency refers to the property that cointegrating estimates will converge to their true values at a faster rate, T, than the usual LS rate of $T^{1/2}$.

innovations in X_t. This dependence on nuisance parameters invalidates conventional inferential procedures. Owing to these problems, it is desirable to use a procedure that is asymptotically optimal under more general conditions. Therefore estimation approaches developed by Phillips and Hansen (1990), Park (1992), and Stock and Watson (1993) are used; these estimators are designed to eliminate nuisance parameter dependencies and possess the same limiting distribution as full-information maximum-likelihood estimates. The latter implies that the resulting estimates will be asymptotically optimal. The application of three different estimators also allows me to examine the robustness of the results.

These approaches can be used to estimate the forward-looking error-correction term $\hat{u}_t = n_t - X'_t\hat{\alpha}$, where $\hat{\alpha}$ is a T-consistent and asymptotically efficient estimate of the long-run parameters. This, in turn, allows me to rewrite equation (10) as

$$\Delta n_t = \beta\Delta n_{t+1} - \gamma\hat{u}_t + \eta_{t+1} \tag{12}$$

Since all variables in (12) are I(0), DGB suggest estimating the discount rate β and the ratio of disequilibrium to adjustment cost γ by a generalized instrumental variable method; Hansen's (1982) generalized method of moments (GMM) estimator is used. This estimator should allow us to control for the effect of an MA(1) process in the composite error term on the standard errors. If the structural error term v_t is serially uncorrelated, then lags of Δn_t and ΔX_t at time *t*-1 or earlier are valid instruments for GMM estimation. However, in order to allow for the possibility that v_t follows an MA(1) process, perhaps owing to the effects of aggregation, the model using lags of Δn_t and ΔX_t at time *t*-2 and earlier is also estimated. To the extent that there are more instruments than parameters to be estimated, the validity of the model is tested using Hansen's (1982) J-test for over-identifying restrictions.

Finally, I turn to the issue of consistently estimating the speed of adjustment term,

λ. Begin by assuming that the law of motion for X_t is integrated of order one — an assumption that cannot be rejected (see the next section). With this assumption, one can use the Wiener-Kolmogorov prediction formula to replace the expectation in the forward solution (5) and derive an estimating equation:

$$\Delta n_t = (\lambda - 1)(n_{t-1} - X'_{t-1}\alpha) + (1-\lambda)\Delta X'_t\alpha + (1-\beta\lambda)(1-\lambda)\mu_t \tag{13}$$

The error-correction form (13) can be estimated using non-linear LS estimation to obtain a consistent measure of λ (see Phillips and Loretan 1991).

4 THE EMPIRICAL RESULTS

4.1 Pretests for integration and cointegration

To implement the two-step procedure, it is necessary to specify the forcing variables X_t that influence the firm's target level of labour (n_t^*). I follow Layard and Nickell (1985) and specify the long-run labour demand as a linear function of the level of real aggregate demand and real producer wages:

$$n_t^* = \alpha_0 + \alpha_1 y_t + \alpha_2 w_t + v_t \tag{14}$$

where n_t is the number of employees, y_t is a real output measure and w_t is a measure of real producer wages calculated as a ratio of labour income (including supplementary income) per paid person-hour and the producer price index.[4] We use quarterly data from 1967Q1 to 1993Q4 and use them in natural log form; the details regarding the data are presented in the Data appendix.[5]

There are several issues to note about equation (14). First, the measure of labour

4. We also used a measure of labour income that did not include supplementary income. Our results did not significantly change. This is not surprising, as the measures of labour income are correlated at over 99 per cent.
5. The choice of the sample is dictated by the availability of labour income data.

input that probably best corresponds to n_t in equation (14) is total hours worked. However, employment data are used for two main reasons: (i) the principal motivation for the analysis is the observation that unemployment is highly persistent, and by focussing on employment I can more directly address this link between adjustment costs and unemployment dynamics; and (ii) fluctuations in total hours worked will in general reflect changes in labour input on both the intensive and extensive margins, and the nature of the adjustment costs on these two margins is probably quite different. For instance, adjustment costs arising from overtime are likely to be smaller than those associated with hiring new workers. To address unemployment persistence, I focus on isolating the adjustment costs associated with changing employment.[6]

Second, equation (14) embodies the effects of the supply side of the labour market by including the real output term. If output is a function of potential output or the non-accelerating inflation rate of unemployment, then supply-side influences such as changes in productivity and/or the cost of capital should be translated into movements in real output. Third, (14) is consistent with a generalized constant-elasticity-of-substitution technology production function, where $-\alpha_2$ is the elasticity of substitution between aggregate labour and capital. In the following paragraphs, I test formally whether equation (14) represents a well-specified long-run or target equation.

I begin by examining the time-series properties of each series using both the augmented Dickey and Fuller (1979) test and a modified version of the Phillips and Perron (1988) Z_α test proposed by Stock (1991). The latter test, denoted the MZ_α test, appears to have better finite-sample properties than both the augmented Dickey-Fuller (ADF) and

6. Nevertheless, to determine whether the distinction between total hours and employment is important in the long-run analysis, I test for the presence of cointegration using total hours as the dependent variable in equation (14). I find slightly weaker evidence of cointegration for total hours than for employment (reported below), which is likely attributable to a declining trend in hours worked from the beginning of the sample to about 1977; thereafter, hours worked appears to be stable around a mean of 37 with a standard deviation of 0.4 hours.

Phillips-Perron tests (see Stock 1991 and Perron and Ng 1994). The ADF and MZ_α tests allow us to test formally the null hypothesis that a series is I(1) against the alternative that it is I(0). However, it is well known that these unit-root tests have weak power in the presence of persistent roots. In order to guard against the possibility that our inability to reject the null hypothesis of a unit root simply reflects a lack of power, a test recently developed by Leybourne and McCabe (1994), which has stationarity as its null, is also applied. Monte Carlo evidence suggests that the Leybourne and McCabe $\hat{s}_\beta$ test encompasses the better known test for stationarity developed by Kwiatkowski et al. (1992) in terms of both size and power. The test statistics for the variables in levels and in first differences are reported in Table 1 (p. 21) and are easy to summarize. All the tests suggest that the variables under consideration are well characterized as nonstationary or I(1) processes in levels and stationary or I(0) in first differences.

As argued in the previous section, an implication of the LQ model is that if the forcing processes y_t and w_t are I(1), then these variables should form a cointegrating relationship with n_t. Evidence of cointegration would suggest that equation (14) captures all the permanent components in labour demand. I test whether this is supported by the data by applying the ADF test suggested by Engle and Granger and the MZ_α test proposed by Stock (1991) to the LS residuals from equation (14). The cointegration test results are as follows. The test statistic from the ADF test is -2.70 (with a data-dependent lag of 3), whereas that from the more powerful MZ_α test is -27.47.[7] The latter provides evidence of cointegration at the 10 per cent level. As we mentioned earlier, as the adjustment costs get larger, the cointegrating regression residual term becomes nearly integrated and hence highly persistent. Thus, if labour adjustment costs are expected to be high, then the marginal evidence of cointegration from relatively weak tests is not surprising (see Gregory

7. The cointegration critical value for the ADF test is calculated from the response surface estimates in MacKinnon (1991), whereas that for Stock's MZ_α test is taken from Haug (1992). Details on the methods used to calculate these test statistics are in the footnote to Table 1.

1994). To control for this problem we extend the $\hat{s}_{\beta}$ test to a cointegration framework.[8] The $\hat{s}_{\beta}$ test statistic is 0.07, which implies that we are unable to reject the null of cointegration even at the 10 per cent level.[9] In all, I tentatively conclude that the variables under study form a valid cointegrating relationship.

As noted above, even though the LS estimates are super-consistent, they will not be asymptotically efficient nor will their asymptotic distribution be standard. To control for these problems, Table 2 (p. 21) reports long-run parameter estimates using the procedures developed by Phillips and Hansen (1990), Park (1992) and Stock and Watson (1993). I find the parameter estimates for y_t and w_t from all three estimators to be statistically significant and to have a priori expected signs. I also find that the estimates are not statistically different from each other. Specifically, the parameter estimates for domestic activity and real producer wages are found to be about 0.8 and -0.3, respectively. The former is broadly similar to the results in Bean, Layard and Nickell (1986), while the latter lies within the range found in a representative survey of the literature by Hamermesh (1986) and that estimated by Kennan (1988b) for Canada. Given the assumed underlying production function, the parameter estimate for the real producer wage implies an elasticity of substitution between labour and capital of about 0.3.

Figure 1 (p. 23) plots the actual (n_t) and target (n_t^*) levels of employment, while Figure 2 (p. 23) displays the gap between the target and actual levels of employment. These figures give another indication of the plausibility of specifying the long-run demand for labour as in Table 2. From the figures it is evident that over the steady growth period of the late 1960s and 1970s, the actual and target levels of employment are generally within 1 per

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8. For this test to be valid the residuals must come from a cointegrating regression where the parameter estimates are both consistent and efficient. Since LS estimation of the cointegrating regression is only consistent, the residuals for the $\hat{s}_{\beta}$ test are taken from a regression using an estimation procedure developed by Stock and Watson (1993).
9. The critical values are taken from Shin (1994).

cent of each other. The only exception is the 1974-75 economic slowdown; over this two-year period a persistent negative gap of slightly over 1 per cent occurs between target and actual employment. In marked contrast to the earlier period, the 1980s represents a period of sharp decline followed by rapid growth in economic activity. The results over the period 1980 to 1983 suggest that in the absence of adjustment costs, firms would have shed considerably more labour, as the target level of employment is continuously lower than the actual level. The negative gap bottoms out at 3.6 per cent in 1982Q1 and closes thereafter. In the expansionary period of the late 1980s, in contrast, the desired level of labour surpassed the actual level by an average of 1.2 per cent. Over the recession of the early 1990s, once again, a persistent difference occurs between the target and actual levels of employment. In contrast with that of the 1982 recession, the gap is not as sharp nor as deep, but it is almost as persistent. Finally, in 1993Q3 and thereafter the labour market appears to go into a state of excess demand; interestingly, the emergence of excess labour demand in 1993Q3 coincides with the beginning of the downward trend in the Canadian unemployment rate.

For the purposes of estimating the Euler equation in the second stage of our procedure, it is important that the long-run parameter estimates be structurally stable over the sample period. Structural stability is tested using a series of parameter constancy tests for I(1) processes recently proposed by Hansen (1992) — the *Lc*, *MeanF* and *SupF* tests. All three tests have the same null hypothesis of parameter stability but differ in their alternative hypothesis. Specifically, the *SupF* is useful in testing whether there is a sharp shift in regime, while the *Lc* and *MeanF* tests are useful for determining whether or not the specified model captures a stable relationship. The results of the *Lc*, *MeanF* and *SupF* tests are 0.38, 3.71 and 6.16, respectively. These values imply that we are unable to reject the null hypothesis of stability for any of the tests at conventional levels of significance. We note that Hansen (1992) suggests that these tests may also be viewed as tests for the null of cointegration against the alternative of no cointegration. Thus, these test results also corroborate the previous conclusion of cointegration among the variables under study.

This evidence of structural stability seems at odds with studies that find an interaction between labour demand and the increased generosity of the Canadian unemployment insurance (UI) system implemented in 1977 (see Corak 1994a for cross-sectional evidence).[10] To further investigate this claim three different versions of (14) are reestimated in an attempt to test for this effect. More specifically, equation (14) is augmented with (i) a dummy variable that takes the value of one after 1977Q3 — the approximate date of the introduction of the more generous "regional extended benefits"; (ii) an interaction term between the dummy variable in (i) and the wage variable; and (iii) an interaction term between real output and the dummy variable.[11] Given a specific date for the possible change in labour demand behaviour, the test should be a more powerful test of the structural stability hypothesis. The reestimation results indicate that one is unable to find any evidence in favour of an interaction between labour demand and the UI system, even at the 70 per cent level. This suggests that the interaction effect found by Corak at the microeconomic level may be lost when one works at the aggregate level. Indeed, Corak (1994b) argues that the impact of UI on the aggregate unemployment rate, and presumably aggregate employment, is likely not very significant.

The evidence from this section suggests that equation (14) is a well-specified model (in a cointegration sense) of long-run or target aggregate labour demand and that the corresponding parameter estimates are T-consistent, asymptotically efficient and stable. In the next section, these parameter estimates are used to form a measure of the forward-looking error-correction term, $\hat{u}_t$, which will be used to estimate the Euler equation (12).

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10. Corak (1994a) finds that some firms will use unemployment insurance as a type to buffer when product demand temporarily falls.
11. These regressions are estimated using the approach developed by Stock and Watson (1993). See the footnote to Table 2 for details.

The long-run equation is also used to estimate the speed of adjustment of labour demand to its steady state.

4.2 Results for the Euler equation

In this section, I report the parameter estimates of the LQ model using a limited-information procedure that is based on the model's Euler equation. In contrast, earlier empirical studies often estimated the parameters of the LQ model using a full-information approach that requires an explicit solution for the model's control variables in terms of the forcing processes. Under full-information maximum-likelihood (FIML) estimation, the process assumed to generate the forcing variables must be specified and estimated jointly with the law of motion and with certain cross-equation restrictions. Provided that the model is correctly specified, the FIML estimator will be more efficient than that based on the Euler equation approach. However, the limited-information approach adopted in this paper provides consistent parameter estimates under more general conditions. In addition, in a Monte Carlo study based on stationary forcing variables, West (1986) finds that even under the assumption of no misspecification, full-information estimation is only moderately more efficient than limited-information estimation.

The structural parameters in equation (12) are estimated using Hansen's (1982) GMM procedure. The instruments include a constant, and lags of Δn_t, Δy_t and Δw_t. Instrument sets lagged one period will yield consistent estimates of β and γ (subject to identification) given the assumption about the composite error term η_t, whereas sets lagged two periods will yield consistent estimates even if the structural error term v_t follows an MA(1) process. Based on the finite-sample results in West and Wilcox (1994), a relatively large number of lagged instruments are included.[12] The error-correction term $\hat{u}_t$ is constructed using Stock and Watson's (1993) cointegrating regression procedure.

12. See the footnote to Table 3 for details.

In the first instance, an attempt is made to estimate both the discount rate and the adjustment parameter by directly estimating the Euler equation. These results are reported in Table 3 (p. 22).[13] The point estimates of β, although significantly different from zero at the 1 per cent level, are in most cases larger than one. In these cases, however, the hypothesis that $\beta < 1$ can not be rejected at the 5 per cent level of significance. As for the estimates of the adjustment parameter, they lie within the range of 0.2 to 0.3, with all values being significant at the 5 per cent level. Recall that γ is the ratio of disequilibrium and adjustment cost parameters, so these estimates suggest that adjustment costs are about fourfold more important than disequilibrium costs in determining the dynamics of the demand for labour. Finally, note that the J-tests are unable to reject the validity of the over-identifying restrictions imposed by the estimation, for any of the instrument sets considered. These J-test results are consistent with a model that is correctly specified. In contrast, previous research on labour demand within the rational expectations and LQ model framework has tended to reject the over-identifying restrictions (see Pfann and Palm 1993).

Although the results in Table 3 are relatively favourable, I next follow the standard practice of fixing the parameter β and then estimating the adjustment parameter from the Euler equation. This is done for two reasons. First, the results in GPS demonstrate the difficulties in identifying β when the forcing variables X_t are generated by an I(1) process. This may be one reason I tend to get point estimates of β greater than one. Second, by estimating γ over a range of reasonable values for the discount parameter, one can discern an indication of the sensitivity of γ to different settings of β. Surprisingly, the estimates of γ lie within the relatively narrow range of 0.20 and 0.27 (Table 4) and are statistically significant at the 5 per cent level. Notice that the estimates of the adjustment parameter

13. These results and those in Table 4 appear to be quite robust to different combinations of the instruments.

appear to be relatively insensitive to the value of the discount parameter and to the instrument sets considered. Again, none of the J-tests reject the over-identifying restrictions, even at the 10 per cent level, for any of the instrument sets or the discount factor considered, which suggests again that the model is correctly specified.[14]

The evidence from this section strongly supports the hypothesis that adjustment costs are an important feature of Canadian labour demand. The results are surprisingly robust and suggest that adjustment costs are about four times as important as disequilibrium costs in determining the demand for labour. To the extent that I can infer firm-level behaviour from aggregate data, this finding suggests that the representative firm will attempt to spread its labour force adjustments over time. To get an idea of the time necessary for the representative firm to adjust its labour demand, the error-correction model (13) can be used to estimate a value of λ. Equation (13) is generalized to include second-order leads and lags of ΔX_t in an attempt to capture any dynamics remaining in the model and then estimate the error-correction model using non-linear LS. I find that, even after I control for the effects of the supply side using the real output term, $\hat{\lambda} = 0.84$ (with a standard error of 0.04). This implies that the half-life of a shock is about one year and that it takes slightly over three and a half years for 90 per cent of the adjustment to be completed.

This estimate appears to be within the range of previous work examining labour demand in an explicit dynamic theory framework. Rose and Selody (1985), working with Canadian data and the context of a large macroeconomic model, find quantitatively similar results. Sargent (1978) and Meese (1980), using U.S. data, find the value of λ to be about 0.95, whereas Nickell (1984) estimates λ to be approximately 0.85 for U.K. manufacturing

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14. As West and Wilcox (1994) point out, the normalization that is used may be important. Therefore I renormalize on Δn_{t+1} in equation (12) and reestimate the parameter γ. The results suggest that the adjustment parameter estimates and the J-test statistics are not sensitive to this renormalization.

employment data. In other related work, Abraham and Housman (1993) estimate the speed of adjustment for different industries in three European countries. They find that the speed of adjustment term ranges from 0.84 to 0.92 in Germany, 0.91 to 0.94 in France and 0.81 to 0.95 in Belgium. Hamermesh (1993) uses pooled microeconomic time-series data from U.S. manufacturing firms and estimates λ to be about 0.84. Finally, Nickell (1986) does some "back of the envelope" calculations for the United States and finds that the speed of adjustment term should lie between 0.83 and 0.96 for white-collar workers and between 0.67 and 0.92 for blue-collar workers. Overall, the results appear to fall within the bounds suggested by previous research and suggest that sluggish adjustment of labour may be one potential cause of Canadian unemployment persistence. It should be noted, however, that most of these studies use estimation approaches that do not carefully consider the time-series properties of the data, so a direct comparison with the current estimates is difficult.

5 CONCLUDING REMARKS

This paper has examined whether adjustment costs are a significant feature of Canadian labour demand. A simple LQ model with integrated regressors is used to estimate the relative importance of adjustment costs as well as the rate of adjustment towards equilibrium. In contrast to many previous studies that have examined the dynamic behaviour of labour demand, I incorporate a forward-looking element into the decision process and carefully consider the times-series properties of the data. Moreover, unlike other studies using the LQ model, the structural parameters are estimated via the Euler equation, using a limited-information approach that does not require an explicit solution for the model's control variables in terms of the forcing processes. A different assumption is also made about the data generation process of the variables, an assumption which appears to be supported by the data. It is important to emphasize that the LQ construct is used as a characterization of labour demand adjustment at the macroeconomic level. Accordingly, no claims about adjustment behaviour at the level of the firm are made.

The estimated parameters are consistent with the underlying dynamic theory and suggest that there are significant adjustment costs in labour demand. Specifically, adjustment costs are found to be about fourfold more important than disequilibrium costs in determining the demand for labour and that it takes over 3 1/2 years for 90 per cent of labour demand adjustment to be completed. These results suggest that costly adjustment of labour demand may be an important contributing factor to the observed persistence of Canadian unemployment.

There are several extensions to this paper that may be worth pursuing. The most obvious is to test explicitly whether the assumption of symmetric labour demand adjustment is a reasonable approximation for aggregate employment data. Another potentially useful extension is to explore the distinction between costs that arise from adjusting the number of workers and those that arise from adjusting total hours worked. As previously mentioned, I use the number of workers in the current analysis, as it facilitates the interpretation between sluggish labour adjustment and unemployment persistence. The nature of adjustment costs for total hours worked may be interesting to look at, since adjustment costs arising from both the extensive and intensive margins are likely to be different from those associated with hiring or firing employees. Finally, it would be interesting to apply the same methodology used in this paper to other countries. If adjustment costs in labour demand are indeed important sources of unemployment persistence, one would expect to find evidence of higher adjustment costs for European labour demand relative to those for Canada and the United States.

Data appendix

This appendix presents the data definitions and reference numbers (provided in parentheses) for the variables used in this study. Unless otherwise noted, the series are drawn from the CANSIM data base. The employment variable (n_t) is defined as the number of paid employees (LFSA270 — Labour Force Survey), and real domestic activity (y_t) is proxied by real gross domestic product (D20463). The wage variable, w_t, is constructed as follows:

$$w = \frac{YW/(52 \cdot HAW \cdot n)}{PGDPFC}$$

where *YW* is labour income (D20088 — D20091), *HAW* is average weekly hours in the commercial and non-commercial sectors (LFSA2050 — Labour Force Survey) and *PGDPFC* is the gross domestic product (GDP) price deflator at factor cost. The latter is constructed as

$$PGDPFC = \frac{PGDP}{1 + (TILGS/(YGDP - TILGS - ENARS))}$$

where *PGDP* is the GDP price deflator (D20011/D20463), *TILGS* is total indirect taxes less subsidies (D20008), *YGDP* is nominal GDP (D20011) and *ENARS* is the national accounts expenditure residual (D20029).

Tables

Table 1:
Tests of the time-series properties of the data
augmented Dickey-Fuller, Stock's MZ_α and Leybourne and McCabe's $\hat{s}_\beta$ tests[a]

Variable	ADF Lags	ADF t-statistic	MZ_α-statistic	$\hat{s}_\beta$-statistic
n_t	1	-1.43	0.58	0.84***
y_t	1	-1.05	0.53	1.59***
w_t	0	-2.59	0.59	0.49***
Δn_t	0	-5.46***	-22.58**	0.07
Δy_t	0	-7.38***	-60.84***	0.05
Δw_t	0	-9.32***	-23.70**	0.07

a. Henceforth, "***", "**", "*" indicate significance at the 1, 5 and 10 per cent levels, respectively. The ADF critical values are calculated from MacKinnon (1991), while the PP critical values are taken from Fuller (1976). All test regressions include a trend term. We use the lag length selection procedure advocated by Ng and Perron (1994) with a 5 per cent critical value. The initial number of AR lags is set equal to the seasonal frequency plus 1 or 5. For the MZ_α test, the spectral density is estimated with an AR(4) spectral estimator (see Stock 1991) and its critical values are taken from Fuller (1976). Finally, the $\hat{s}_\beta$ test statistic is calculated using an ARMA(2,1) representation, and its critical values are those in Kwiatkowski et al. (1992).

Table 2:
Estimation of the static labour demand equation[a]

Variable	SW	PPH	CCR
Constant	-6.790*** (0.158)	-7.090*** (0.466)	-7.587*** (0.248)
y_t	0.748*** (0.031)	0.789*** (0.058)	0.787*** (0.023)
w_t	-0.257*** (0.064)	-0.335** (0.119)	-0.300*** (0.045)

a. Standard errors are in parentheses. The SW estimates are based on second-order leads and lags and a Newey and West (1987) consistent variance-covariance estimator with the truncation parameter set equal to the seasonal frequency. The PPH estimates are based on a VAR(1) prewhitening quadratic kernel procedure of Andrews and Monahan (1992). The CCR estimates are those from the third stage as suggested by Park and Ogaki (1991).

Table 3:
Estimates of the Euler equation[a]

	1	2	3	4
β	1.114*** (0.104)	1.077*** (0.099)	1.075*** (0.141)	0.913*** (0.110)
γ	0.272** (0.106)	0.300*** (0.086)	0.210** (0.096)	0.223*** (0.084)
J-test	13.624	15.344	11.732	14.351

a. Standard errors are in parentheses. The models are estimated using Hansen's (1982) GMM estimator with a Bartlett kernel suggested by Newey and West (1987). The second-stage estimates of the weighting matrix are estimated using a lag length of one to allow for an MA(1) error process. The instrument sets include constant Δn_{t-i}, Δy_{t-i} and Δw_{t-i}. The first column of results corresponds to lags from 1 to 5; the second 1 to 6; the third 2 to 6; and the fourth 2 to 7.

Table 4:
Estimates of the adjustment term for preset values of beta[a]

	1	2	3	4
β = 0.990 J-test	0.232** (0.097) 14.389	0.266*** (0.079) 15.906	0.203** (0.089) 12.069	0.247*** (0.077) 13.179
β = 0.975 J-test	0.226** (0.096) 14.535	0.259*** (0.079) 16.052	0.200** (0.087) 12.263	0.242*** (0.076) 13.318
β = 0.950 J-test	0.216** (0.094) 14.796	0.248*** (0.079) 16.313	0.195** (0.085) 12.600	0.224*** (0.075) 13.572
β = 0.900 J-test	0.195** (0.091) 15.358	0.225*** (0.076) 16.876	0.185** (0.083) 13.299	0.218*** (0.073) 14.144

a. See footnote, Table 3.

Figure 1: Actual versus target level of employment

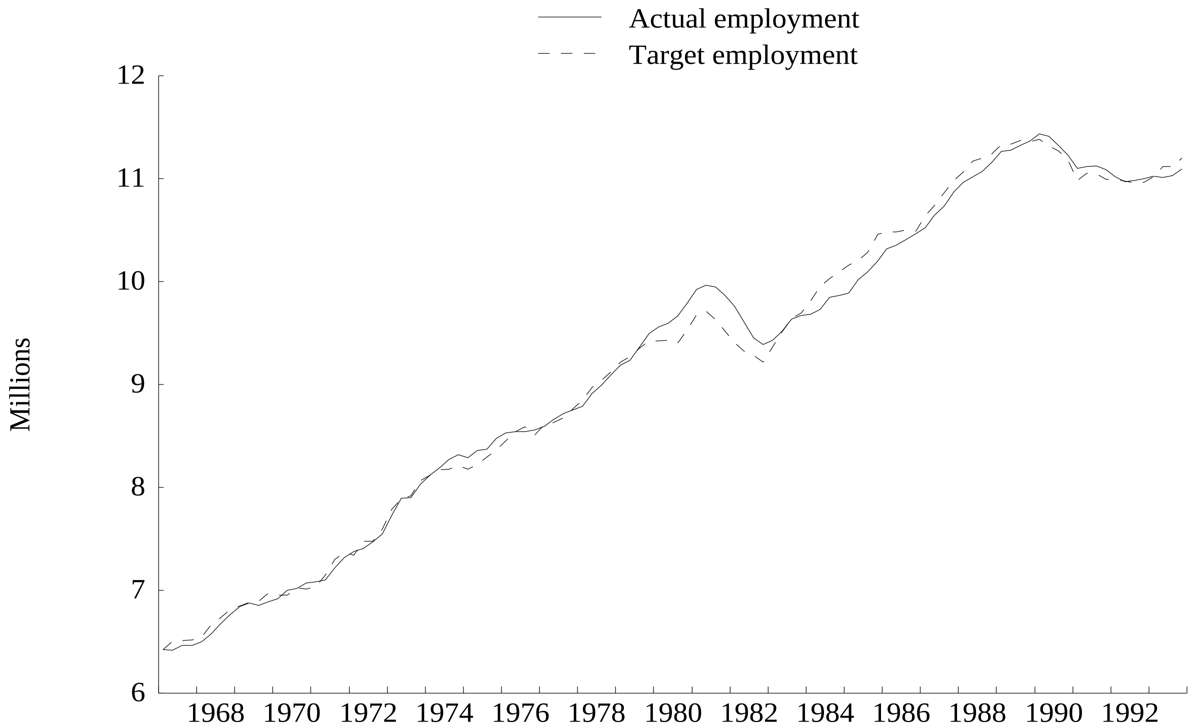
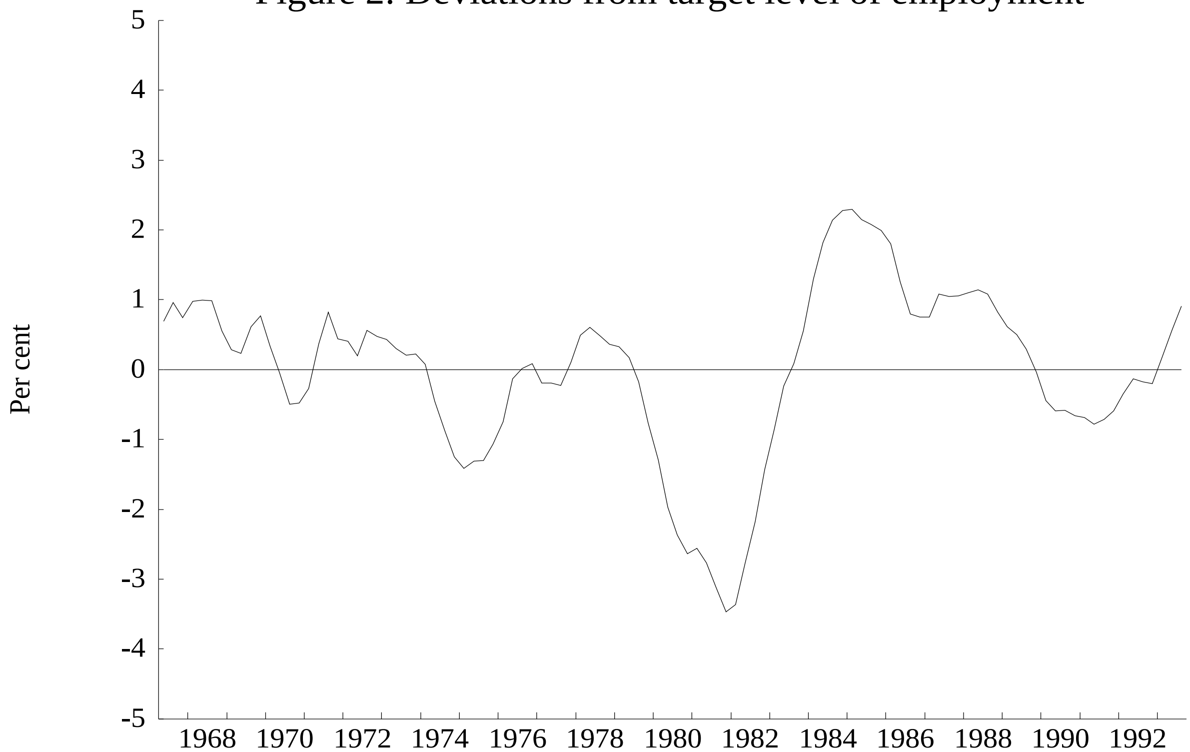


Figure 2: Deviations from target level of employment



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