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## Abstract

The existing macroeconometric evidence lends support to the wage Phillips curve by showing a negative relation between the *rate of change* in wages and the unemployment rate, conditional on lagged price inflation. Most theoretical models of wage setting, however, generate a “wage curve,” described by a negative relation between the *level* of the real wage and unemployment. Real wage dynamics have important implications for how shocks affect aggregate consumer price inflation, and for the determination of the natural rate of unemployment. This paper examines the dynamics of the aggregate wages in Canada, and tests whether real wages and unemployment have a long-term *level* relationship. The results indicate that a simple aggregate wage Phillips curve continues to describe the behaviour of aggregate wages in Canada quite well. The micro evidence, however, does not unequivocally support one specification against the other; rather, what seems to emerge is more complex wage dynamics better described in an error-correction specification. Wage changes reflect the short-run movement in the unemployment rate, while they adjust towards a long-run equilibrium level, as could be described in a wage curve model.

*JEL classification: J31*
*Bank classification: Inflation and prices*

## Résumé

Les résultats de modèles macroéconométriques donnent à penser que la courbe de Phillips des salaires peut être représentée par une relation négative entre le *taux de variation* des salaires et le taux de chômage, à condition d’inclure un terme relatif à l’inflation retardée. La plupart des modèles théoriques de détermination des salaires, cependant, produisent une « courbe des salaires » définie par une relation négative entre le *niveau* des salaires réels et le taux de chômage. La dynamique des salaires réels a des implications importantes quant aux répercussions des chocs sur l’augmentation globale des prix à la consommation et pour la détermination du taux de chômage naturel. Dans cette étude, l’auteur examine la dynamique globale des salaires au Canada et vérifie si les *niveaux* des salaires réels et du chômage sont liés en longue période. Il constate qu’une simple courbe de Phillips agrégée des salaires décrit toujours assez bien le comportement global des salaires au Canada. Les observations microéconomiques ne valident toutefois pas sans équivoque une spécification au détriment de l’autre; elles font plutôt émerger une dynamique des salaires plus complexe, dont rend mieux compte une structure de correction des erreurs. Les variations salariales réagissent aux fluctuations à court terme du taux de chômage, tout au long de leur ajustement vers un niveau d’équilibre à long terme du genre de celui que décrit un modèle relatif à la courbe des salaires.

*Classification JEL : J31*
*Classification de la Banque : Inflation et prix*

# 1. Introduction

The relationship between wages and unemployment has been the subject of numerous studies in the macroeconomic and labour literature. Two competing models have emerged on the wage-unemployment trade-off: (i) supported by macroeconometric evidence, a wage Phillips curve with a negative relation between the *rate of change* in wages and the unemployment rate, and (ii) supported by theoretical models of wage setting, a "wage curve" described by a negative relation between the *level* of the real wage and unemployment.

Understanding the wage-unemployment trade-off is important from a policy perspective, as it has implications for how shocks affect the evolution of wages and prices. Consequently, wage dynamics determine the role of productivity shocks and supply shocks in the context of a price Phillips curve that describes the trade-off between aggregate consumer price inflation and the unemployment rate or the output gap. The role of productivity in determining the natural rate of unemployment is also affected by how wages are determined.

From a monetary policy perspective, the interest in this trade-off is twofold. First, a better understanding of wage dynamics in itself adds to the knowledge of the functioning of the economy and provides an empirical guide for the modelling of wage determination in the central bank's models. Second, given the importance of the Phillips curve in the formulation of policy and the significant supply shocks to which the economy is subject, the assumptions made on the specification of the effects of these shocks on price inflation in a Phillips curve framework need to be tested.

The intuition behind the effect of wage dynamics is simple: a quick response of wages to changes in productivity growth would reduce substantially the effect of the productivity shocks on price inflation. Furthermore, if workers bargain in terms of real wage change, then supply shocks that erode the level of real wages will only temporarily raise price inflation. In contrast, if workers bargain in terms of the level of real wages, then such shocks will continue to impact wage bargaining and price inflation in later periods.[1]

This paper reports on a preliminary investigation of the Phillips curve specification for wages versus the wage curve specification in Canada using micro-level data.[2] We examine the dynamics of the aggregate wage, and we test whether real wages and unemployment in Canada have a long-term

1. In Canada, for example, it has been noted that, following oil price shocks, aggregate real wages have diverged significantly from what could be considered as their equilibrium level.
2. Earlier work by Crawford and Dupasquier (1993) and Cozier (1991) on wage dynamics provides some insights into this question for Canada.

*level* relationship. Micro data provide a more appropriate testing ground for comparing Phillips curve and wage curve specifications. We extend our analysis using household data drawn from Statistics Canada's Survey of Consumer Finances (SCF) and firm-level data from collective bargaining agreements in the unionized sector. These micro data have the additional advantages (beyond what is methodological) of broad representative coverage of the Canadian labour market, detailed information on the individual level (in the case of the SCF), and an error-free measurement of the wage rate (in the case of union settlements). The Canadian experience could be useful for understanding the role of labour market institutions in determining wage dynamics. The Canadian labour market is similar to that of the United States in many ways, but it has a much higher unionization rate and more widespread collective bargaining practices, similar to those in many European countries.

Our results highlight the importance of using micro data: a simple aggregate wage Phillips curve continues to describe quite well the behaviour of aggregate wages in Canada, as it does for the United States. Our estimates show that while there could be a long-term level relationship between wages and unemployment, the adjustment to that level is very slow. Controlling for the composition of the work force and firm characteristics does not alter these results. However, the micro evidence suggests that wage dynamics are better described by an error-correction model. The province-level wage estimates do point to a significant convergence of wages to an equilibrium-level relation. The estimates on the firm level are even more supportive of a faster adjustment in the Canadian union sector. The results also reflect the important role of these institutions in the determination of wage dynamics.

This paper is organized as follows. Section 2 describes the traditional Phillips curve and the wage curve, and the implication for the determination of the natural rate of unemployment. Section 3 briefly reviews the empirical literature. Section 4 describes the data and presents the empirical results. Section 5 concludes.

## 2. The Phillips Curve and the Wage Curve

The traditional derivation of the accelerationist price Phillips curve has two main building blocks: (i) a wage Phillips curve—a negative relationship between the expected real wage change and the unemployment rate, and (ii) a markup pricing rule and adaptive expectations.[3]

3. Expectation formation has been the subject of recent ongoing debate on the Phillips curve (e.g., the New Keynesian paradigm). In this paper, we stay within the traditional framework of the accelerationist models.

Algebraically, the wage Phillips curve can be expressed as

$$w_t - p_t^e = w_{t-1} - p_{t-1} + \alpha - \beta u_t + \Delta x_t + \varepsilon_t, \tag{1}$$

where *w* and *p* are the wage level and price level, respectively; $p^e$ is the expected price level; $\Delta x$ is productivity growth; and the unemployment rate is *u*. If inflation expectations are backward-looking and firms set prices as a markup over unit labour costs, then price expectations and the markup equations will take the following form:

$$p_t^e - p_{t-1} = p_{t-1} - p_{t-2}, \tag{2}$$

and,

$$p_t = w_t - x_t + \mu_t, \tag{3}$$

where $\mu$ and $x$ are the markup and productivity measures, respectively. Rewriting equation (1) using the above expectation formation assumption and the markup pricing equation, we get the standard accelerationist Phillips curve:

$$\Delta p_t = \Delta p_{t-1} + \alpha - \beta u_t + \Delta\mu_t + \varepsilon_t. \tag{4}$$

Equation (4) implies that the non-accelerating inflation rate of unemployment (NAIRU) in the absence of supply shocks is constant and equal to $\alpha/\beta$.

Unlike the specification in equation (1), almost all theoretical models of wage setting generate a negative relation between the *level* of the real wage and unemployment, given the reservation wage and the level of productivity.[4] A simple representation of these models is:

$$w_t - p_t^e = \rho b_t + (1-\rho)x_t + \alpha_0 - \beta u_t + \varepsilon_t. \tag{5}$$

All variables are defined as before. The term *b* is the reservation wage, while *x* is a measure of productivity or rents generated by the worker-firm specific match. The parameter $\rho$ ranges from 0 to 1. In some efficiency wage models, such as the shirking model of Shapiro and Stiglitz, productivity does not affect wages directly. In bargaining models (rent sharing), $\rho$ is typically less than 1, since wages depend on the surplus from the match. Institutional factors such as the dependence of unemployment benefits on previous wages suggest that the reservation wage would

4. Efficiency wage models (Shapiro and Stiglitz 1984) and bargaining models (Mortensen and Pissarides 1994) are typical examples in the literature.

depend on the lagged wage. Fairness models of wage determination also suggest that workers' aspirations in job search and wage bargaining are shaped by their previous earnings. Productivity increases that influence non-labour income or earning opportunities in the informal sector (or home production) are also determinants of the reservation wage. For simplicity, we assume a naive parameterization of the reservation wage,[5]

$$b_t = a + \theta(w_{t-1} - p_{t-1}) + (1-\theta)x_t, \tag{6}$$

where $\theta \in [0,1]$. The reservation wage is assumed to be homogeneous of degree 1 in the real wage and productivity; otherwise, technological progress will lead to a persistent trend in the unemployment rate. Replacing this reservation wage in equation (5) gives:

$$w_t - p_t^e = \alpha + \theta\rho(w_{t-1} - p_{t-1}) + (1-\theta\rho)x_t - \beta u_t + \varepsilon_t, \tag{7}$$

where $\alpha = \rho a + \alpha_0$. Formally, we can derive the price Phillips curve from equation (7) using the expectation hypothesis (equation (2)) and the price markup (equation (3)). Price inflation will then take the following form:

$$\Delta p_t = \Delta p_{t-1} + \alpha + \mu_t - \theta\rho\mu_{t-1} - \theta\rho\Delta x_t - \beta u_t + \varepsilon_t. \tag{8}$$

Note that equation (8) is a generalization of the price dynamics specified in equation (4), the traditional Phillips curve. In fact, under the assumption that productivity has no effect, either directly ($\rho = 1$) or indirectly through the reservation wage ($\theta = 1$), on expected real wages, the effect of supply shocks in the two specifications becomes identical. In this case, the reservation wage is simply the lagged wage level, and equation (7) becomes a relationship between wage growth and the unemployment rate. Any supply shocks will have temporary effects on prices, as they will not affect wage bargains. However, in the general case, when $\theta\rho < 1$, the shocks to the level of energy prices, interest rates, or payroll taxes, all of which have a direct influence on the markup, $\mu$, will have a persistent effect on inflation.

## 2.1 The role of productivity

In the classical price Phillips curve specification, productivity growth plays no role. The negative effect of productivity growth on price inflation given wages (equation (3)) is fully offset by the positive effect on wage inflation. In our more general specification, productivity growth matters.

5. Although fairly simple, this specification captures the essential arguments advanced in the literature on aspiration wages. Different specifications have similar qualitative results for our model.

Because the reservation wage adjusts slowly to productivity, wage growth will not adjust one-to-one to productivity. As a result, productivity acceleration will cause a favourable shift in the inflation-unemployment trade-off as described in equation (8). We can derive the equilibrium rate of unemployment consistent with non-accelerating inflation:

$$u_t = \left(\frac{1}{\beta}\right)(\alpha + \mu_t - \theta\rho\mu_{t-1} - \theta\rho\Delta x_t) . \tag{9}$$

The equilibrium unemployment rate as described in equation (9) shows how productivity gains could reduce the natural rate of unemployment. The value of $\theta\rho$ will have an effect on the dynamics of the NAIRU. We can also derive the equilibrium natural unemployment rate in the absence of supply shocks and with no productivity increases ($\mu_t = \mu$ and $x_t = x$):

$$u = \left(\frac{1}{\beta}\right)(\alpha + (1 - \theta\rho)\mu) . \tag{10}$$

Finally, with a constant markup and productivity growth, the accelerationist Phillips curve specification described in equation (4) is consistent with any value of $\theta\rho$. Under this assumption, this parameter could not be identified from an aggregate inflation regression. The two price-inflation specifications, equation (4) and equation (8), are indistinguishable, as all values of $\theta\rho$ generate observably identical behaviour for aggregate price inflation.[6]

The rest of this paper empirically examines the behaviour of wage dynamics in Canada. We develop an empirical strategy to identify and then estimate $\theta\rho$ using various data sets from the Canadian labour markets. We use aggregate data to facilitate comparison with the existing macro literature. We also use micro-level data to address some of the shortcomings of the aggregate analysis. The estimated parameters are then compared across data sets and with other countries' existing estimates.

## 3. The Literature

Since different wage dynamics have a direct effect on the specification of the price Phillips curve, one could test competing hypotheses using either wage or price data. We will review some of the papers that use labour-market data from the United States and Europe to test the hypotheses. We will also review two papers that studied the recent behaviour of price inflation in the United States

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6. Roberts (1997) has also demonstrated the consistency of the wage curve with an alternative "New-Keynesian" version of the Phillips curve.

and examined directly the effect of supply shocks on consumer price inflation and the role of productivity in the Phillips curve.

### 3.1 Wages and unemployment: micro vs. macro evidence

The original relationship between wages and unemployment, as documented by Phillips based on data for the United Kingdom, relates the growth in nominal wages to the unemployment rate. However, the modern accelerationist specification restates the Phillips curve as a relationship between the rate of change of expected real wages and the unemployment rate. Equation (1) is an algebraic description of that hypothesis. The Phillips curve relationship has received a lot of attention in the economic literature, and has been subject to various modifications. There is wide macroeconometric support for the Phillips curve in the United States. Most time-series estimates suggest that a simple specification, such as equation (1), is a good representation of the aggregate data on U.S. wage inflation, price inflation, and the unemployment rate. For OECD Europe, a modified specification of equation (7), with an error correction but a strong autocorrelation of wages, seems to better fit the aggregate data. Blanchard and Katz (1999) argue that, on average, the estimated $\theta\rho$ in Europe is around 0.75. Turner, Richardson, and Rauffet (1996) estimate a wage specification that is similar to equation (7) (with richer dynamics) for several OECD countries using semi-annual aggregate wage data and unemployment series.[7] Their results are in line with the Blanchard and Katz (1999) assessment.

One key shortcoming of the Phillips curve is its lack of theoretical underpinning. Textbook labour market models on wage determination generate a wage dynamics equation that differs from the empirical specification described by the accelerationist Phillips curve. Furthermore, Blanchflower and Oswald (1994) argue that estimation using aggregate data may spuriously bias the effects of lagged wages on current wages. Consequently, the use of micro data is more appropriate to test the Phillips curve specification. Their cross-country empirical analysis suggests that the correct relationship between wages and unemployment is better described by a “wage curve” that relates the level of expected real wages to the unemployment rate.[8]

Blanchflower and Oswald (1994) present a wide range of evidence on the appropriate functional form of the relation between wages and local unemployment rates. They conclude that log wages are a monotonically decreasing (and convex) function of local unemployment, and that the wage curve relation is well approximated by a simple log linear function. They argue that the negative

7. The countries are the United States, Japan, Germany, France, the United Kingdom, Italy, and Canada.
8. The countries studied are the United States, the United Kingdom, Canada, Australia, South Korea, and seven other European countries. See Card (1995) for a book review of Blanchflower and Oswald.

relationship between wages and unemployment is identical across countries and stable over time. Their evidence is based on estimating a variation of equation (7) using pooled cross-sections of micro data and controlling for year and region effects. They argue that the wage curve is an equilibrium locus of wages and unemployment that essentially replaces the market-level labour-supply function. They reject the Phillips curve specification by showing evidence of no significant autoregression in real wages. In particular, using data from the Current Population Survey (CPS) between 1979 and 1987 in the United States, and from the General Household Survey Series (GHSS) between 1973 and 1990 in Britain, they show that the estimated coefficient on the lagged real wage in equation (7) is close to zero, in both the United States and Britain (see their Tables 4.27 and 6.20, for example).

The strong rejection of the wage Phillips curve based on micro evidence generated a debate in the empirical labour literature. Card (1995) questions the validity of the results shown by Blanchflower and Oswald. He cites two sources of bias against the Phillips curve specification in a model like the one described in equation (7): (i) technical problems associated with the presence of both a lagged dependent variable and a regional fixed effect, and (ii) the possibility of serial correlation in the residual. Although Card does not produce any estimates himself, he argues that it is premature to reject the Phillips curve and calls for more empirical research on the dynamic relation between wages and unemployment.

Blanchard and Katz (1997) echo the concerns suggested by Card (1995) and point to further shortfalls of the Blanchflower and Oswald results. Using administrative data on average weekly earnings of covered workers from the U.S. unemployment insurance system, data on average hourly earnings for manufacturing-production workers from a large-scale-establishment survey, and data on hourly earnings from the Merged Annual Outgoing Rotation Groups in the CPS, their regressions constantly yield estimates of the coefficient on the lagged wage that is close to one. They claim that their results differ from those of Blanchflower and Oswald because the latter use annual earnings from the March CPS as the basic wage measure. This wage measure mixes up actual changes in wages with changes in hours of work. This systematic measurement error and the sampling error, owing to the relatively small March CPS samples used to measure annual state average wages, bias down the autoregressive coefficient in a regression of current log wage on lagged log wage and other covariates.

Blanchard and Katz (1999) offer an explanation for the difference between Europe and the United States. Consider equation (7): $\theta\rho$ could be different than 1 if either $\rho < 1$ or $\theta < 1$. It is not sufficiently clear to justify the difference in $\theta$ between Europe and the United States, unless one is willing to introduce hard-to-measure non-market labour activities like the underground economy or home production. Regarding $\rho$, the role of unions and the labour-market regulations (stringent hiring and firing) in Europe could explain these differences in wage-setting behaviour. Abowd et al. (1998), using employer-employee longitudinal data from France and the United States, show evidence of a greater direct effect of productivity on wages in France than in the United States. Canadian labour markets might display behaviour similar to that in Europe. For example, Abowd and Lemieux (1993), using wage-settlements data, show significant rent-sharing activities in the union sector in Canada.

## 3.2 Relative price shocks and inflation dynamics

An alternative way to estimate $\theta\rho$ is to investigate the direct effect of relative price movements on inflation behaviour. Whelan (1999) estimates an augmented version of equation (8) for the United States. As an illustration, consider the consumer price index (CPI) and assume that a proportion $(1-\psi)$ of the consumption bundle is priced as a markup over domestic unit labour costs $(p_t = \mu_t + w_t - x_t)$, while the rest consists of imported goods (or others like food and energy) for which prices are set exogenously to $p_t^x$:

$$\begin{aligned} p_t^c &= \psi p_t^x + (1-\psi) p_t \\ &= p_t + \psi(p_t^x - p_t). \end{aligned}$$

The implicit markup of the consumption price index over domestic unit labour costs becomes $\mu_t^c = p_t^c - w_t + x_t = \mu_t + \psi(p_t^x - p_t)$.

Now, using the consumption prices and substituting the consumption markup into equation (8), we get:

$$\Delta p_t^c = \Delta p_{t-1}^c + \alpha + \mu_t - \theta\rho\mu_{t-1} - \theta\rho\Delta x_t - \beta u_t + \psi(p_t^x - p_t) - \theta\rho\psi(p_{t-1}^x - p_{t-1}) + \varepsilon_t.$$

Using the exogenous variation of the relative prices to identify $\theta\rho$, the estimated parameters are derived as the ratio of the coefficients of the lagged and contemporaneous values of the relative price terms.[9] While Whelan's results suggest an estimate of $\theta\rho$ that is statistically different from 1, the point estimate is around 0.8, which is closer to the traditional Phillips curve than the wage curve assumption of $\theta\rho = 0$. Note, however, that the evolution of the estimated time-varying NAIRU differs. The two series (one based on $\theta\rho = 1$ and one based on $\theta\rho = 0.8$) tell a similar story of a declining NAIRU in the 1990s: the decline varies in magnitude, with the series corresponding to $\theta\rho = 0.8$ falling by only half as much as the series corresponding to $\theta\rho = 1$

9. The relative price indicators are energy and import prices.

since the mid-1980s. Whelan (1999) concludes that the evidence points against the traditional Phillips curve formulation, with important implications for the recent behaviour of the NAIRU in the United States.

## 3.3 Productivity growth and the Phillips curve

The behaviour of the U.S. inflation and unemployment in the 1990s led some economists to argue that the “new economy” is responsible for the improvement in the inflation-unemployment trade-off. Ball and Moffit (2001) present a model in which worker’s aspirations for wage increases adjust slowly to shifts in productivity. Their model yields a price Phillips curve with the gap between productivity growth and aspiration wage growth as an additional explanatory variable:

$$\Delta p_t = \Delta p_{t-1} + \alpha - \delta(\Delta x_t - \Delta b_t) - \beta u_t + \varepsilon_t .$$

The authors spend considerable time explaining what determines the workers’ aspiration wages. The following reservation wage dynamics were adopted to emphasize how reservation wages adjust over time to the most recent real wage increases:

$$\Delta b_t = \beta \Delta b_{t-1} + (1-\beta)(\Delta w_{t-1} - \Delta p_{t-1}) .$$

Although Ball and Moffit discuss the wage-formation influence on the specification of the Phillips curve, they do not fully integrate a model of the labour market in the derivation of the Phillips curve. However, they show that there is empirical evidence for this specification using the U.S. data. This model could explain, for example, why the productivity slowdown in the 1970s, combined with relatively high wage demands, caused an unfavourable Phillips curve shift. By the same token, the acceleration of productivity in the 1990s and the sluggishness in wage growth caused a favourable shift in the Phillips curve.

In summary, the literature is not conclusive. Using wage measures, the macro data support the Phillips curve in the United States and a slightly modified error-correction model in Europe. The micro evidence casts doubt on even the Phillips curve specification for the United States. The importance of the different wage dynamics in the derivation of the price Phillips curve has also been recognized. The evidence from the price Phillips curves remains mixed. Recent studies in the United States suggest a modified version of the Phillips curve. The proposed specification of the price inflation-unemployment trade-off seems to help to explain some puzzles regarding the dynamics of price inflation and unemployment in the last decade in the United States. Institutional factors could partly explain the differences between the United States and continental Europe. With respect to Canada, the lack of adequate wage data is responsible for the absence of recent evidence from the Canadian labour markets. This paper fills that void. We use several micro and macro data sources on wages and unemployment to explore wage dynamics in Canada.

# 4. Data

The aggregate data are mainly annual time series of wages, prices, real output, and the unemployment rate in Canada. These series are drawn from public-use files available from CANSIM (URL: http://www.statcan.ca/english/CANSIM). At the micro level, either individual wages from annual household surveys or wage settlements from collective bargaining agreements in the unionized sector are used. The following sections describe the data and some of the adjustments we had to make.

## 4.1 Aggregate data

We construct a series of annual aggregate wages from 1978 to 1999. Before 1984, we used total labour compensation in the business sector (Statistics Canada, *I602002*) plus payroll taxes paid by the business sector (Statistics Canada, *D18212*, multiplied by the share of the business sector in the aggregate gross domestic product). For the period 1984 to 1999, the wage measure used is the fixed-weight average hourly earnings series for the Survey of Employment, Payroll, and Hours (SEPH). This measure is constructed by Statistics Canada using fixed weights by industry and region.[10]

The labour productivity series in the non-farm business sector (Statistics Canada, *I602502*) is defined as the ratio of real GDP at factor cost to the total hours worked in the non-farm business sector. Trend productivity is proxied by the cyclically adjusted measure of labour productivity, calculated as the predicted value from a regression of hourly output on a quadratic trend for the 1961–99 period. For prices, we used the all-items CPI and the CPI excluding the effect of indirect taxes. The unemployment rate of the working-age population is drawn from Statistics Canada's Labour Force Survey.

## 4.2 Micro data

On the firm level, we used contract data from the wage-settlements file covering collective bargaining agreements in the Canadian private sector for the period 1977–99. The data set includes 4,474 contracts with 500 and more workers. Each of these settlements corresponds to an

10. The survey covers all industries except agriculture, fishing and trapping, private households services, religious organizations, and military personnel of defence services.

identified bargaining unit, allowing for the construction of panel data for 993 bargaining units. The available information for each contract includes its start and end dates, the number of employees covered by the agreement at the start date, the industry affiliation, and the geographic location. The wage measure is the base wage in each month of the contract, typically paid to the lowest skill group covered by the collective bargaining agreement.

Table A.1 in Appendix A provides some summary statistics on the settlements sample. The sample spans a 23-year period, with a relatively high number of contracts in the first half. The average duration for the multi-year contracts is about 2 1/2 years, although it varies somewhat by year, with relatively longer contracts in the 1990s. The nominal-wage rate shows double-digit growth in the late 1970s and early 1980s, and then remains fairly stable in the second half of the 1980s. Nominal-wage growth, however, slowed down substantially in the 1990s. Average employment coverage shows no trend or adjustment to the cycle.

On the individual level, we assembled 16 annual microdata files from the SCF to construct a consistent wage series over the years 1981–97. The SCF contains information on annual income, as well as personal and labour-related characteristics of individuals aged 15 years and over. In particular, information is available on wages and salaries, income from self-employment in the previous year, labour force status, number of weeks worked in the previous year, full-time/part-time status last year, number of hours worked in the reference week, occupation and industry, years of experience and seniority, and educational attainment. Other demographic characteristics, such as age, gender, marital status, language spoken, immigration status, and geographic location, are also available.[11]

Two potential drawbacks arise when using average weekly earnings from the SCF as a measure of the wage rate over the business cycle. First, average weekly earnings may vary because of changes in the underlying (hourly) wage rate or because of changes in hours worked per week. Second, changes in the composition of the work force may understate the cyclicality of real wages, since the skill level of the work force tends to decrease during expansions and increase during recessions. In the rest of this paper, we will use the wage series adjusted for human capital, other socio-economic characteristics, and hours, obtained by computing “regression-adjusted” measures of the wage rate. More specifically, we use ordinary least squares (OLS) to estimate the following wage equation:

11. Table A.3 presents the distribution of workers across provinces, industries, and sectors. About 65 per cent of the (weighted) observations are concentrated in Quebec and Ontario, while more than half of the individuals work in the manufacturing, trade, and service industries. About 19 per cent of the sample is in the public sector. For further details about the SCF data, see Farès and Lemieux (2000).

$$w_{it} = \beta X_{it} + \sum_{t=1}^{16} \delta_t Year_t + \varepsilon_{it}, \tag{11}$$

where $w_{it}$ is log average weekly earnings of individual $i$ in year $t$ (earnings are deflated by total annual CPI); $X_{it}$ includes various observable characteristics such as age, education, sex, marital status, language spoken, tenure, industry dummies, province dummies, full-time dummy, and actual hours of work;[12] and $Year_t$ is a dummy variable for each year in the sample. The estimated coefficients of the unrestricted year dummies, $\hat{\delta}_t$, $t = 1...16$, can then be interpreted as the regression-adjusted measures of the wage rate; i.e., the predicted yearly wage rate of an individual with a fixed set of characteristics. The last two columns in Table A.2 and Figure A.1 show these constructed wage series. The graph indicates that adjusting for the compositional shifts in the work force results in different dynamics for wages, particularly owing to the business cycle effect. For example, at the end of the 1980s adjusted wages grew faster than unadjusted wages, while between 1992 and 1994 adjusted wages remained virtually unchanged when unadjusted wages had continued their modest rise. We interpret this as evidence for a composition bias in aggregate wage measures.

We also used equation (11) to compute an aggregate wage measure in the union sector, based on wage settlements. In this case, $X_{it}$ includes contract characteristics such as duration, number of workers covered, an indicator for the cost-of-living adjustment (COLA), industry, and region controls. Similarly, to construct adjusted measures of wages at the provincial level, we estimate a model with a full set of province-year interactions:

$$w_{ijt} = \beta X_{ijt} + \sum_{j=1}^{10} \sum_{t=1}^{16} \delta_{jt} Prov_j * Year_t + \varepsilon_{ijt}, \tag{12}$$

where $Prov_j$, for $j = 1,..,10$, is a set dummy variable for provinces. The estimated province-year effects $(\hat{\delta}_{jt})$ again can be interpreted as regression-adjusted measures of the wage rate in a province, $j$, in year $t$ (i.e., the wage in different provinces and different years for an individual with a specified set of characteristics).

12. Actual hours are drawn either from the SCF for the survey week or for similar workers in the Labour Force Survey for the actual year. See Appendix A for more detailed information.

# 5. Empirical Results

In this section, we estimate the model as described in equation (7). Our objective is to test the hypothesis supporting the traditional Phillips curve specification, $\theta\rho = 1$. We use aggregate wage data as a benchmark, and we compare the results with those for the United States. We then estimate the model using provincial-level wage and unemployment data, as well as data from union contracts.

## 5.1 The macroeconometric evidence

Following the formulation in Blanchard and Katz (1997), we use a modified version of equation (7). More precisely, we consider the following empirical specification:

$$\Delta w_t = \alpha_w + \Delta p_{t-1} - \lambda(w_{t-1} - p_{t-1} - x_{t-1}) - \beta u_t + \varepsilon_t, \tag{13}$$

where $\lambda = 1 - \theta\rho$. Inflation expectations are assumed to be backward-looking. This error-correction specification implies that if $\lambda > 0$, then the real wage adjusts to a level determined by trend productivity and the unemployment rate. The speed of adjustment will depend on the magnitude of $\lambda$. The standard wage Phillips curve specification can be seen as imposing the restriction $\lambda = 0$ (or $\theta\rho = 1$).

Table 1 reports the estimated results of equation (13) for the United States and Canada. The U.S. results are taken from Blanchard and Katz (1997), while the Canadian estimates are our own. The coefficient on lagged inflation is constrained to be equal to 1 in the spirit of the accelerationist Phillips curve. Although oversimplified, this empirical specification fits the data quite well in both countries. The table highlights the success of the standard macroeconometric wage equation (linking wage inflation to the unemployment rate) in the United States as well as in Canada. The estimated slope of the Phillips curve is negative and significant. The parameter estimates indicate that each time the unemployment rate increases by 1 percentage point, wage growth decreases between 0.85 and 0.99 per cent in both countries.

The point estimate of $\lambda$ in the United States is negative, which implies the absence of any long-term level relation between the real wage and the unemployment rate. For Canada, the estimated $\lambda$ has the right sign, but it is very small in magnitude and not statistically significant.

**Table 1: Wage Inflation and Unemployment [a]**

| | Dependent variable = $\Delta w_t$ | | | |
|---|---|---|---|---|
| | **United States[b] 1970–95** | | **Canada[c] 1963–99** | |
| Constant | 7.20 (1.40) | 6.75 (1.31) | 0.079 (0.009) | 0.053 (0.065) |
| $\Delta p_{t-1}$ | 1.00[d] | 1.00[d] | 1.00[d] | 1.00[d] |
| $U_t$ | -0.99 (0.21) | -0.95 (0.19) | -0.85 (0.11) | -0.88 (0.15) |
| $(w_{t-1} - p_{t-1}) - x_{t-1}$ | | 0.11 (0.05) | | -0.017 (0.043) |
| $\bar{R}^2$ | 0.67 | 0.72 | 0.57 | 0.56 |

a. Standard errors are shown in parentheses.
b. From Blanchard and Katz (1997), Table 1, page 61.
c. The definitions of the Canadian data are compatible with the U.S. data used in Blanchard and Katz (1997).
d. The coefficient is constrained to equal 1.

For Canada, the sample includes periods of high inflation as well as more recent episodes of low inflation combined with the introduction of the Canadian inflation targets. The sample also includes periods of unusual supply shocks, like the oil-price shock in the 1970s. As a result, the parameter estimates might be sensitive to the assumption on the expectation formation, and might not be robust over different sample periods. To address these issues, Table 2 shows the estimated parameters from equation (13) when the coefficient on lagged inflation is not constrained to equal 1, and over different sample horizons.[13] All the columns in the table show an estimated coefficient on lagged inflation of less than 1, perhaps reflecting changes in inflation expectations. The accelerationist hypothesis might be too constraining. Consequently, in the tables that follow we present the estimates with and without this constraint being imposed. Table 2 displays two other interesting patterns: (i) the coefficient on the unemployment rate decreases in the more recent sample periods, indicating a possible flattening in the Phillips curve, and (ii) although all the

13. We also experimented with different deflators as alternatives to the CPI. We used the GDP deflator as well as a convex combination of the CPI and the GDP deflator, with different weights. The estimated results are quite robust to these sensitivity checks.

estimates of $\lambda$ are not statistically significant, this estimate changes sign and becomes large in magnitude in recent periods.

**Table 2: Canadian Aggregate Wage Inflation and Unemployment[a]**

| | Dependent variable = $\Delta w_t$ | | | | |
|---|---|---|---|---|---|
| | **1963–99** | | **1970–99** | **1978–99** | **1982–97** |
| Constant | 0.085<br>(0.009) | 0.22<br>(0.08) | 0.23<br>(0.09) | -0.09<br>(0.11) | -0.34<br>(0.37) |
| $\Delta p_{t-1}$ | 0.81<br>(0.15) | 0.68<br>(0.10) | 0.64<br>(0.12) | 0.79<br>(0.12) | 0.69<br>(0.14) |
| $U_t$ | -0.81<br>(0.11) | -0.58<br>(0.17) | -0.79<br>(0.23) | -0.54<br>(0.21) | -0.23<br>(0.30) |
| $(w_{t-1} - p_{t-1}) - x_{t-1}$ | | 0.089<br>(0.054) | 0.082<br>(0.058) | -0.088<br>(0.068) | -0.21<br>(0.20) |
| $\bar{R}^2$ | 0.78 | 0.79 | 0.81 | 0.76 | 0.65 |

a. Source: Blanchard and Katz (1997), Table 1, page 61.

We have already mentioned in section 4 how an aggregate wage series could mask some important changes in the composition of the workforce. As a check, using equation (11) we construct an aggregate adjusted wage series from the SCF and the wage-settlements data. We re-estimate equation (13) using the constructed series and present the results in Table 3.

As expected, the use of adjusted wages increases their measured cyclicality. The estimated slope of the Phillips curve reported in Table 3 is larger than what we previously reported. This difference reflects the extent of the composition bias.[14] The constraint on the lagged inflation coefficient does not seem to affect these results. On its own, the accelerationist hypothesis cannot be rejected in the error-correction specification, column 4, where the estimated coefficient is shown to be not statistically different from 1. The estimates of $\lambda$ are not altered significantly by the use of adjusted wage data. In the unconstrained specification, this estimate is similar to the one presented in Table 2 for the same sample period. This coefficient is also imprecisely estimated.

We perform the same exercise using wage-settlements data from the union sector. We construct an aggregate wage growth measure adjusted for the industry and regional location as well as for

14. See Farès and Lemieux (2000) for more details on the composition bias in the SCF.

other observable contract characteristics, such as its duration and whether it includes an escalator clause. Table 4 presents the results.

**Table 3: Adjusted Wage Inflation and Unemployment, SCF**

| | Dependent variable = $\Delta w_t$ | | | |
|---|---|---|---|---|
| | SCF[a] 1982–97 | | | |
| Constant | 0.15 (0.03) | -2.16 (0.75) | 0.13 (0.02) | -1.39 (2.00) |
| $\Delta p_{t-1}$ | 1.00[b] | 1.00[b] | 0.62 (0.12) | 0.86 (0.33) |
| $U_t$ | -1.62 (0.31) | -1.05 (0.31) | -1.32 (0.27) | -1.13 (0.36) |
| $(w_{t-1} - p_{t-1}) - x_{t-1}$ | | -0.28 (0.09) | | -0.19 (0.25) |
| $\bar{R}^2$ | 0.62 | 0.76 | 0.68 | 0.67 |

a. Source: SCF. Wages are adjusted to individual observable characteristics (see the text for further description). The other variables are the same as described in Table 1.
b. The coefficient is constrained to equal 1.

The slope of the Phillips curve is estimated to be smaller in magnitude in the union sector than in the more representative SCF data. The difference is remarkable, as a 1 percentage point increase in the unemployment rate leads to a 1.1 per cent decrease in the SCF wages, but only a 0.4 per cent decrease in the union wages. Unions appear to play an important role in shielding workers' wages from the effect of market fluctuations. The slope of the Phillips curve is once again estimated to be smaller in the 1982–97 period than in the full-sample estimates.

The estimated coefficient of lagged inflation is statistically less than 1, again casting doubt on the accelerationist specification. Regarding the error-correction model, the estimated parameters in Table 4 concur with what Tables 2 and 3 presented: namely, the estimates of $\lambda$ are generally small and insignificant, and the sign of the estimate becomes positive in the 1982–97 period.

**Table 4: Adjusted Wage Inflation and Unemployment, Wage Settlements[a]**

| | Dependent variable = $\Delta w_t$ | | | | |
|---|---|---|---|---|---|
| | 1978–99 | | | | 1982–97 |
| Constant | 0.07<br>(0.01) | -0.13<br>(0.33) | 0.07<br>(0.01) | 0.56<br>(0.46) | -0.71<br>(0.94) |
| $\Delta p_{t-1}$ | 1.00[b] | 1.00[b] | 0.88<br>(0.06) | 0.81<br>(0.08) | 0.78<br>(0.08) |
| $U_t$ | -0.71<br>(0.14) | -0.72<br>(0.14) | -0.73<br>(0.13) | -0.72<br>(0.13) | -0.40<br>(0.22) |
| $(w_{t-1} - p_{t-1}) - x_{t-1}$ | | -0.03<br>(0.06) | | 0.09<br>(0.08) | -0.14<br>(0.17) |
| $\bar{R}^2$ | 0.55 | 0.53 | 0.92 | 0.92 | 0.87 |

a. Sources: Wage-settlement data are from the collective bargaining agreements in the Canadian union sector. Wages are adjusted to contract-specific observables (see the text for further description). The other variables are the same as described in Table 1.

b. The coefficient is constrained to equal 1.

Taken as a whole, the macro results in the above tables do not constitute a strong rejection of the traditional specification of the wage Phillips curve. There might be some doubts regarding the accelerationist hypothesis on its own, but conditional on lagged inflation the wage Phillips curve seems to fit the aggregate data in Canada quite well. Adjusting wages for composition shifts in the workforce does not alter these results. The evidence from the union sector is in line with the aggregate results. There is some suggestive evidence in favour of the error-correction specification in the 1982–97 period, but the coefficients are imprecisely estimated. In fact, given the slow adjustment of wages to their equilibrium level implied by the estimated $\lambda$, one could argue that the Phillips curve provides a simple and satisfactory model for wage behaviour over the relevant policy horizon.

## 5.2 The microeconometric approach

Blanchflower and Oswald (1994) raise concerns about the non-robust nature of the macro estimates and question the reliability of the estimated parameters using aggregate time-series analysis. Several misspecification issues could be cited to underline the weakness of the estimation. For example, the formation of inflation expectations could differ from the simple backward-looking or autoregressive specification used. The measurement of underlying productivity as well as the omission of possibly highly autocorrelated wage determinants are all possible criticisms of the described estimation.

Micro-level data provide a more appropriate testing ground for comparing Phillips curve and wage curve specifications. The possibility of strongly autocorrelated unobservables that affect wages as well as difficult-to-measure variables such as inflation expectations can all be addressed properly in a panel-data regression model. We look at the behaviour of wages and unemployment rates across regions within Canada over time. The typical empirical approach using regional data starts with equation (13) and assumes that the expected price inflation and productivity variables relevant for wage setting are independent of the region and can be captured by time dummies. The following regression is then used:

$$\Delta w_{it} = \alpha_i + \gamma_t - \lambda w_{i,\,t-1} - \beta u_{it} + \varepsilon_{it}\,, \tag{14}$$

where $i$ indexes the region, and $\alpha_i$ and $\gamma_t$ are two sets of unrestricted region and year dummies, respectively. We use equation (12) to construct adjusted provincial wage series from the SCF. The first two columns of Table 5 report the results of estimating equation (14) using the adjusted provincial wage panel.

The first specification corresponds to the standard Phillips curve where $\lambda$ is constrained to zero, while in the second column $\lambda$ is estimated. The estimated slope, $-\beta$, is negative and significant. The magnitude of this estimate is less than what is reported in Table 3 using aggregate data. This could be partly a result of including year dummies in the provincial specification. These dummies would absorb most of the time variation, leaving $\beta$ to be identified mainly through the cross-provincial variation in the unemployment rate.[15]

In column 2 of Table 5, note the improved fit of the regression when $\lambda$ is unconstrained. This parameter is positive and precisely estimated. It is also twice as large as previously reported in the aggregate analysis. The point estimate suggests that the wage adjustment to the equilibrium level is complete in less than five years. These results indicate that the aggregate estimate of $\lambda$ might be biased towards zero, which justifies the large empirical support of the traditional Phillips curve when using aggregate time series.

15. This result is common to the estimates of the slope of disaggregate Phillips curves. Cross-regional migration could be responsible for the smaller impact of local unemployment on local wages.

**Table 5: Provincial Wage Inflation and Unemployment[a]**

| | Dependent variable = $\Delta w_{it}$ | | | | | |
|---|---|---|---|---|---|---|
| | **Adjusted hourly wages[b]** | | **Unadjusted weekly earnings[b]** | | **Unadjusted annual earnings[b]** | |
| Constant | 0.04<br>(0.02) | 0.01<br>(0.02) | 0.03<br>(0.02) | 0.01<br>(0.02) | 0.03<br>(0.02) | 4.00<br>(0.68) |
| $U_{it}$ | -0.34<br>(0.18) | -0.20<br>(0.17) | -0.32<br>(0.20) | -0.26<br>(0.18) | -0.12<br>(0.19) | -0.18<br>(0.17) |
| $w_{i,\,t-1}$ | | -0.28<br>(0.05) | | -0.36<br>(0.06) | | -0.38<br>(0.06) |
| No. of observations | 160 | 160 | 160 | 160 | 160 | 160 |
| $\bar{R}^2$ | 0.42 | 0.50 | 0.34 | 0.47 | 0.43 | 0.54 |

a. Source: Provincial wages are calculated using the SCF data base for the 1981–97 period.

b. All regressions include an unrestricted set of year dummies and an unrestricted set of province dummies. Provincial weights are used in the regression.

Using yearly state-level data from the United States, and region-level data from the United Kingdom, Blanchflower and Oswald (1994) estimate $\lambda$ to be close to 1. Once region and time effects are included, their estimate suggests that the wage adjusts to the unemployment rate rapidly with no interesting dynamics. Blanchard and Katz (1997), however, report that regressions using various microdata from the United States between 1980 and 1991 consistently yield estimates of $\lambda$ close to 0 (0.03 to 0.09). Blanchard and Katz suggest that the basic source of difference between their results and the results reported by Blanchflower and Oswald (1994) is the different wage measure. They argue that a systematic measurement error from the use of an inappropriate wage measure and a sampling error in the lagged wage measure both lead to an upward-biased estimate of $\lambda$.

Using the SCF wage data, we construct different wage measures and test whether alternative wage measures could lead to different estimates of $\lambda$. Columns 3 and 4 in Table 5 show the estimated parameters for weekly earnings that are not adjusted for the individual characteristics or hours of work. Columns 5 and 6 show the results using annual earnings (similar earnings measures were used by Blanchflower and Oswald). The results indicate that different wage measures lead to different estimates of $\lambda$. In no case, however, are these estimates close to one.

These findings partly explain the higher estimate (in absolute value) for $\lambda$ when an unadjusted wage series is used, but they do not fully explain the difference in the estimates between Europe and the United States. We can shed more light on this issue using wage settlements in the union sector in Canada. The role of unions could be looked at in this context; moreover, given the panel structure of the data, fixed-effect estimators can be used to control for unobservable firm characteristics.

Table 6 shows the results of estimating equation (14) using firm-level data from the union sector. Columns 1 and 2 show that the unemployment rate has a negative and significant effect on wage growth at the firm level. Again, the estimated slope is not as large as shown in the macro results.[16] Surprisingly, the absolute value of the estimated $\lambda$ in column 2 is very small, albeit significant. The presence of an unobserved firm-specific effect that is (positively) correlated with the level of the lagged wage could lead to a (downward) biased estimate of $\lambda$. First-differencing the data within the same bargaining unit would get rid of this fixed effect (FE). The results are shown in column 3. The estimated $\lambda$ is much larger in magnitude. It is statistically significant and comparable to estimates from the province-level wage regressions using the SCF.

Finally, common to dynamic panel-data models, first-differencing would create a correlation between the lagged wage level and the error term. This could lead to a biased FE estimator. Instrumental variable (IV) techniques could be used to correct the bias. Column 4 reports the IV estimates using further lags of the wage level as instruments for the previous wage.[17] The parameters are less precisely estimated when IV methods are applied; however, the point estimates of $\lambda$ are not significantly different. This estimate is sensitive to the choice of instruments. Moreover, Hausman's specification test does not imply a strong rejection of the FE model as specified in column 3. As in the case of the province-level wage data, the magnitude of this parameter implies that wages (at the firm level) converge to their equilibrium level within five years.

16. In Table 6, the estimated standard errors of the coefficient on the unemployment rate are biased, owing to the different level of aggregation of the wage measures and the unemployment rate. The standard errors of the estimated coefficient of the lagged wages are not subject to this bias, and consequently no attempt is made to correct for it.

17. Using further lags within the same bargaining unit would limit the sample to include only those units that have at least four consecutive contracts reported in the data.

**Table 6: Firm-level Wage Inflation and Provincial Unemployment**

| | Dependent variable = $\Delta w_{it}$ | | | |
|---|---|---|---|---|
| | **OLS**[a] | | **FE**[a,b] | **FE-IV**[a,c] |
| Constant | 0.053<br>(0.004) | 0.10<br>(0.00) | | |
| $U_{it}$ | -0.36<br>(0.03) | -0.34<br>(0.03) | -0.28<br>(0.04) | -0.28<br>(0.16) |
| $w_{i,\,t-1}$ | | -0.020<br>(0.003) | -0.31<br>(0.02) | -0.27<br>(0.36) |
| No. of observations | 3217 | 3217 | 2767 | 2293 |
| $\bar{R}^2$ | 0.75 | 0.75 | 0.69 | |

a. All regressions include controls for the contract duration, a dummy for COLA, an unrestricted set of year, and region and industry dummies.

b. Fixed-effect: all variables are first-differenced to eliminate the time-invariant firm-specific effect.

c. Instrumental variables for the (first-differenced) real wage at the end of the previous contract include 23-year effects, the change in unemployment rate, and the one lagged change in the real wage at the end of the previous contract. The *p*-value of Hausman's overidentification test is 0.252. The statistic is distributed as chi-squared with 1 degree of freedom.

In all, the results from our micro analysis suggest that wage dynamics as usually described in the standard wage Phillips curve model are misspecified. Our estimates indicate that neither a Phillips curve specification nor a wage curve specification on its own could describe the wage behaviour in Canada. Wages do appear to adjust toward a long-run equilibrium level, but with important short-run dynamics. An error-correction model with a speed of adjustment in the 0.2 to 0.3 range seems to better fit the dynamics of regional wages in the last two decades. Firm wage levels from the union sector appear to have similar dynamics.

## 6. Conclusion

Wage dynamics have important consequences on price inflation and on the equilibrium unemployment rate. More precisely, they determine how productivity growth and supply shocks affect aggregate consumer price inflation, and have a bearing on the determination of the NAIRU. The empirical literature disagrees on the nature of wage dynamics. Macroeconometric evidence lends considerable support to the traditional Phillips curve, while some micro studies reject the Phillips curve specification in favour of a wage curve specification.

This paper has tested these two different specifications. We estimated an error-correction model of wages in Canada. The model nested the Phillips curve and the wage curve specifications and allowed us to test which model better fit the data. Using aggregate data, our macro results are in line with most previous macro studies that support the traditional specification of the wage Phillips curve. These results are robust to changes in the composition of the labour force, as well as to changes in the industrial composition of the sample of firms.

The micro evidence, however, is less straightforward. The results do not unequivocally support one specification against the other; rather, what seems to emerge is more complex wage dynamics better described in an error-correction specification. Wage changes reflect short-run movements in the unemployment rate, while they adjust towards a long-run equilibrium of the sort that can be described in a wage curve model.

Our results suggest that the price inflation-unemployment trade-off as described in the traditional accelerationist Phillips curve model might be misspecified. Productivity growth appears to play an important role in shifting this curve. The benign inflation and unemployment behaviour in the 1990s could reflect favourable productivity shocks in the last decade. While this paper does not test this hypothesis directly, future extensions of this research should go in this direction, similar to the recent work by Ball and Moffit (2001) in a U.S. context.

# References

Abowd, J. and T. Lemieux. 1993. "The Effects of Product Market Competition on Collective Bargaining Agreements: The Case of Foreign Competition in Canada." *Quarterly Journal of Economics* 108: 983–1014.

Abowd, J., F. Kramarz, D. Margolis, and K. Troske. 1998. "The Relative Importance of Employer and Employee Effects on Compensation: A Comparison of France and the United States." Cornell University. Photocopy.

Ball, L. and R. Moffit. 2001. "Productivity Growth and the Phillips Curve." NBER Working Paper No. 8421.

Blanchard, O.J. and L. Katz. 1997. "What We Do Know and Do Not Know About the Natural Rate of Unemployment." *Journal of Economic Perspectives* 11(1): 51–72.

———. 1999. "Wage Dynamics: Reconciling Theory and Evidence." *American Economic Review* 89(2): 69–74.

Blanchflower, D. and A. Oswald. 1994. *The Wage Curve*. Cambridge: MIT Press.

Card, D. 1995. "The Wage Curve: A Review." *Journal of Economic Literature* 33(2): 785–99.

Cozier, B. 1991. *Wage and Price Dynamics in Canada*. Technical Report No. 56. Ottawa: Bank of Canada.

Crawford, A. and C. Dupasquier. 1993. "Can Inflation Serve as a Lubricant for Market Equilibrium?" In *Economic Behaviour and Policy Choice Under Price Stability*, 49–80. Proceedings of a conference held by the Bank of Canada, October 1993.

Farès, J. and T. Lemieux. 2000. "Downward Nominal Wage Rigidity: A Critical Assessment and Some New Evidence for Canada." In *Price Stability and the Long-Run Target for Monetary Policy*, 3–31. Proceedings of a conference held by the Bank of Canada, June 2000.

Mortensen, D.T. and C. Pissarides. 1994. "Job Creation and Job Destruction in the Theory of Unemployment." *Review of Economic Studies* 61(3): 397–415.

Roberts, J. 1997. "The Wage Curve and the Phillips Curve." Board of Governors, Federal Reserve Board Finance and Economic Discussion Series Paper No. 1997–57.

Shapiro, C. and J. Stiglitz. 1984. "Equilibrium Unemployment as a Worker Discipline Device." *American Economic Review* 74(3): 433–44.

Turner, D., P. Richardson, and S. Rauffet. 1996. "Modelling the Supply Side of the Seven Major OECD Economies." OECD Economics Department Working Paper No. 167.

Whelan, K. 1999. "Real Wage Dynamics and the Phillips Curve." Board of Governors, Federal Reserve Board Finance and Economic Discussion Series Paper No. 2000–2.

# Appendix A: Data Description

## A.1 Survey of Consumer Finances

The SCF provides large samples of around 40,000 workers for each year between 1982 and 1997, with the exception of 1983, when the survey was not conducted. Public-use samples are also available for heads of households and spouses every other year for the 1970s. Data for all workers are available starting only in 1981. The survey was discontinued after 1997. For all available years, the SCF was conducted in April as a supplement to the Labour Force Survey (LFS), and it asked a battery of questions about income in the previous year, in addition to the usual LFS questions that pertain to the reference week.

The wage measure we use is average weekly earnings, expressed in 1991 dollars. Earnings are defined as the sum of wages and salaries from all types of civilian employment. Included are gross cash wages and salaries received in the reference year from all jobs, before deductions for pension funds, hospital insurance, income taxes, Canada Savings Bonds, etc. Tips and net commissions are also included; taxable allowances and benefits provided by employers are not. For each individual in a given sample year, average weekly earnings are calculated as the ratio of annual wages and salaries, excluding income from self-employment and rental property, to the total weeks worked in that year. We also restrict the sample to workers aged 20 to 65.

An hourly wage rate cannot be computed directly, since the SCF does not provide direct information on the number of hours worked per week in the previous year. Fortunately, several indirect measures of hours worked per year can be used to control for variation in hours. As stated earlier, the SCF collects information on hours worked during the reference week and on whether the worker worked full-time during the previous year. We have also computed direct measures of actual hours worked per week by detailed category of worker, using the monthly micro-data files from the LFS, from 1981 to 1997. Matching these hours measures to workers in the SCF provides an additional proxy for weekly hours of work in the previous year. Our strategy is to use regression methods to “adjust” average weekly wages for changes in weekly hours of work, as proxied by these different measures.

## A.2 Tables

**Table A.1: Wage-Settlements Characteristics by Year**

| Year | Number of contracts | Average duration | $\Delta w_t$ [a] | $w_{t-1}$ [b] | Average employment |
|---|---|---|---|---|---|
| 1977 | 37 | 24.49 | 8.15 | 5.80 | 1,774 |
| 1978 | 310 | 22.83 | 8.64 | 6.04 | 1,345 |
| 1979 | 228 | 25.76 | 11.31 | 6.55 | 1,558 |
| 1980 | 242 | 26.70 | 12.05 | 7.24 | 1,259 |
| 1981 | 194 | 27.68 | 12.57 | 8.26 | 1,667 |
| 1982 | 205 | 24.43 | 9.58 | 9.35 | 1,417 |
| 1983 | 207 | 25.96 | 5.09 | 10.77 | 1,528 |
| 1984 | 283 | 27.87 | 3.60 | 13.09 | 1,846 |
| 1985 | 198 | 28.70 | 3.50 | 11.37 | 1,345 |
| 1986 | 226 | 27.09 | 3.17 | 14.54 | 1,735 |
| 1987 | 213 | 30.70 | 3.71 | 12.49 | 1,495 |
| 1988 | 232 | 29.65 | 4.88 | 15.10 | 1,982 |
| 1989 | 149 | 30.54 | 5.27 | 13.66 | 1,587 |
| 1990 | 246 | 30.36 | 5.61 | 15.71 | 2,078 |
| 1991 | 174 | 28.97 | 3.81 | 16.63 | 1,191 |
| 1992 | 192 | 31.26 | 2.27 | 18.73 | 1,749 |
| 1993 | 176 | 32.32 | 1.37 | 16.84 | 2,299 |
| 1994 | 156 | 35.10 | 1.45 | 17.16 | 1,444 |
| 1995 | 169 | 33.51 | 1.35 | 20.39 | 1,914 |
| 1996 | 160 | 34.58 | 2.15 | 16.40 | 1,612 |
| 1997 | 152 | 39.46 | 2.10 | 17.56 | 1,452 |
| 1998 | 197 | 39.44 | 1.95 | 19.65 | 1,868 |
| 1999 | 128 | 36.56 | 2.78 | 19.18 | 1,670 |
| Overall | 4,474 | 29.74 | 5.28 | 13.32 | 1,640 |

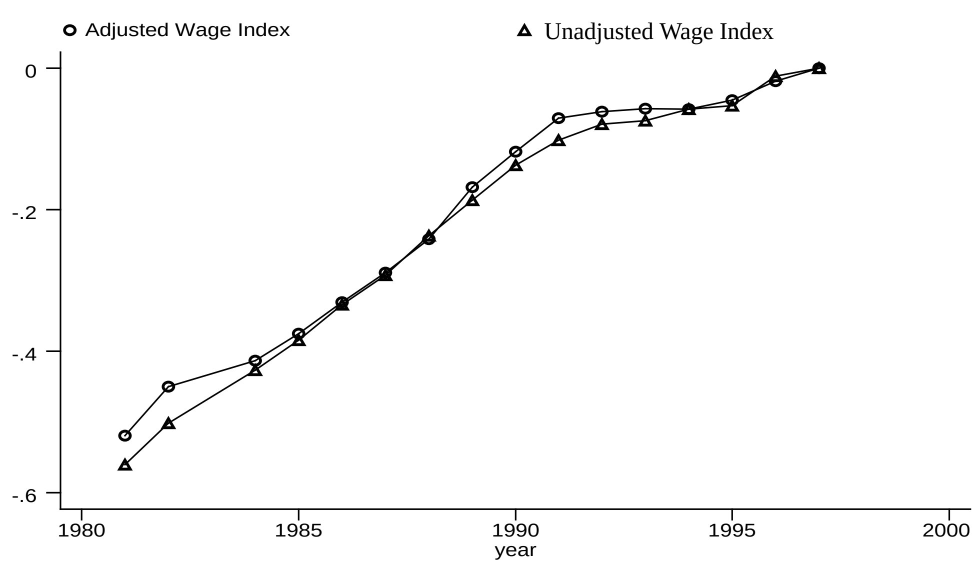
a. Average annual nominal wage growth.
b. Wage level at the end of the previous contract.

**Table A.2: Descriptive Statistics by Year**

| | Aggregate Data | | | Survey of Consumer Finances | | | |
|---|---|---|---|---|---|---|---|
| | | | | Unadjusted wages | | Adjusted wages | |
| Year | Inflation rate[a] | Unempl. rate | $\Delta w_t$ [b] | $\Delta w_t$ [c] | $w_t$ [d] | $\Delta w_t$ [c.] | $w_t$ [d.] |
| 1981 | 12.13 | 7.58 | 10.28 | | -.560 | | -.518 |
| 1982 | 10.73 | 10.97 | 9.64 | 5.85 | -.501 | 6.86 | -.450 |
| 1984 | 3.93 | 11.31 | 4.40 | 7.49 | -.426 | 3.72 | -.412 |
| 1985 | 3.46 | 10.68 | 3.91 | 4.22 | -.384 | 3.76 | -.375 |
| 1986 | 3.26 | 9.66 | 3.56 | 5.01 | -.334 | 4.48 | -.330 |
| 1987 | 4.02 | 8.83 | 3.63 | 4.20 | -.292 | 4.15 | -.288 |
| 1988 | 3.27 | 7.77 | 3.69 | 5.59 | -.236 | 4.75 | -.241 |
| 1989 | 4.33 | 7.56 | 5.33 | 5.02 | -.186 | 7.36 | -.167 |
| 1990 | 4.21 | 8.13 | 5.40 | 4.88 | -.137 | 5.02 | -.117 |
| 1991 | 3.36 | 10.33 | 5.04 | 3.59 | -.101 | 4.70 | -.070 |
| 1992 | 1.06 | 11.15 | 3.74 | 2.22 | -.078 | 0.93 | -.061 |
| 1993 | 1.63 | 11.36 | 1.97 | 0.48 | -.074 | 0.41 | -.056 |
| 1994 | 1.40 | 10.38 | 1.49 | 1.64 | -.057 | -0.04 | -.057 |
| 1995 | 1.96 | 9.44 | 2.19 | 0.49 | -.052 | 1.25 | -.044 |
| 1996 | 1.56 | 9.65 | 2.57 | 4.16 | -.010 | 2.69 | -.017 |
| 1997 | 1.59 | 9.12 | 0.84 | 1.09 | 0 | 1.78 | 0 |

a. Total CPI excluding indirect tax effects.
b. Average annual nominal wage growth. From the total labour compensation in the business sector and fixed-weight average hourly earnings from SEPH.
c. Average annual nominal wage growth.
d. Nominal wages are normalized to 0 in 1997.

Figure A.1 Constructed Aggregate Wage Indices



Adjusted vs. Unadjusted Wages (In Levels)

**Table A.3: Provincial and Industrial Distribution, SCF 1981–97**

| Sample composition (Percentage) | | | |
|---|---|---|---|
| **Province** | | **Industry** | |
| Newfoundland | 1.91 | Agriculture | 1.31 |
| Prince Edward Island | 0.46 | Other primary | 2.39 |
| Nova Scotia | 3.11 | Manufacturing (non-durables) | 9.00 |
| New Brunswick | 2.61 | Manufacturing (durables) | 8.67 |
| Quebec | 25.37 | Construction | 5.74 |
| Ontario | 38.44 | Transportation, communication | 8.02 |
| Manitoba | 3.62 | Wholesale trade | 4.71 |
| Saskatchewan | 2.86 | Retail trade | 11.77 |
| Alberta | 9.44 | Finance, insurance, real estate | 5.95 |
| British Columbia | 12.19 | Community services | 19.16 |
| | | Personal services | 7.94 |
| **Sector** | | Business and miscellaneous | 7.88 |
| Private | 81.35 | Public administration | 7.47 |
| Public | 18.65 | | |

Note: The estimated frequency distributions are all weighted.

Source: Statistics Canada, Survey of Consumer Finances. Cross-sectional files from 1981 to 1997. No data are available for 1983. Sample size is 623,875.

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