The Dynamics of Income Shares and the Wage Curve-Phillips Curve Controversy[1]

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Abstract. This paper argues that the Phillips curve–wage curve controversy cannot be settled within the conventional testing frameworks and suggests an alternative test, which builds on the model of Blanchard and Katz (1997). Using long macro data for the OECD countries, the evidence gives very strong support for the Phillips curve and indicates that wage behaviour is no different among the OECD countries. This implies that adverse supply shocks, which push wages in excess of the full employment equilibrium, have only temporary effects on real product wages and therefore cannot explain the persistently high unemployment in most European countries.

JEL Classification: C23, E24, E3, J30, J60

Key words: Wage curve, Phillips curve, mean-reversion of factor shares.

1. Introduction

Since the publication of the seminal book on the wage curve of Blanchflower and Oswald (1994) there has been an explosion in wage curve estimates using regional data.[2] Most of the studies find support for the Blanchflower and Oswald (1994) hypothesis that wages adjust almost instantaneously to changes in unemployment. These results challenges to the standard macroeconomic framework, where the Phillips curve has traditionally represented the supply side of the economy and ensured that income and unemployment automatically gravitate toward a unique equilibrium following supply shocks. Abandoning the Phillips curve implies that supply shocks have persistent effects on output and unemployment.

Blanchard and Katz (1997, 1999) have disputed the notion of the universality of the wage curve as found by Blanchflower and Oswald, and argue that Blanchflower and Oswald's (1994) estimates, which are based on regional data are biased against the Phillips curve for the US. Blanchard and Katz (1997) suggest an alternative method that tests

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whether labour's share in total income is mean reverting using an error-correction model. Similar tests to discriminate between the two hypotheses have been conducted by Johansen (1995) for Norway and OECD (1997) for the OECD countries. Based on their own and OECD's (1997) estimates, Blanchard and Katz (1999) argue that US labour market dynamics are represented by the Phillips curve, whereas most European labour markets are represented by the wage curve. Blanchard and Katz (1999) suggest that this may partly explain why, since the beginning of the 1970s, the adverse supply shocks have had permanent effects on unemployment in most of Europe, but not the US.

This paper argues that the tests of mean reversion of income shares, which are employed by Blanchard and Katz (1997), Johansen (1995), and OECD (1997) to discriminate between the wage curve and the Phillips curve, cannot be used to discriminate between the two competing hypotheses of wage behaviour. It is shown that mean reversion in income shares is consistent with both the wage curve hypothesis and Phillips curve hypothesis.

Building on the Blanchard-Katz' framework an alternative test to discriminate between the wage curve and the Phillips curve, is suggested. Furthermore, pooled cross section and time-series macro data are used to overcome the low power of the tests that are based on individual country estimates. However, the key parameters are allowed to vary across countries.

Section 2 surveys the micro evidence on the wage curve and concludes that the Blanchflower-Oswald approach is biased against the Phillips curve and is overly sensitive to specification and the choice of estimator. In Section 3 it is shown that tests of mean reversion of labour's income share, as suggested by Blanchard and Katz (1997), cannot be used to discriminate between the wage curve and the Phillips curve because both hypotheses are consistent with mean reversion in labour's income share. The macroeconomic implications of the dynamic adjustment of wages are examined in Section 4, and in Section 5 it is shown that supply shocks can have persistent effects on labour's income share under both the wage curve and the Phillips curve hypotheses, thus rendering it difficult to empirically distinguish between the two hypotheses in small samples. Section 6 tests the wage curve against the Phillips curve using a large panel data set for the OECD countries and Section 7 concludes the paper.

2 Micro evidence on the wage curve

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[2] Albaek *et al* (2000), Baltagi *et al*, 2000, Bell (1996), Bell *et al* (2000), Black and FitzRoy (2000), Bratsberg and Turunen (1996), Dyrstad and Johansen (2000), Groot *et al* (1992), Janssens and Konings (1998), Kennedy and Borland (2000), Pannenberg and Schwartze (1998), and Wagner (1994), among others.

Blanchflower and Oswald (1994) base their conclusions on estimates of the following regression:

$$w_{jt} = f_j + a\, w_{j,t-1} - b\, u_{jt} + g\, X_{jt} + TD_t^{'} d + e_{jt} \qquad (1)$$

where j stands for region j, w is the log of wages, X is regional characteristics, u is the log of the rate of unemployment, f_j are fixed effect dummies, TD_t are time-dummies, and e is a disturbance term.[3] Fixed effect dummies were not included in all of their estimates. Time-dummies are included in the regressions to capture the effects on wages of advances in productivity, increasing prices and other shocks that are common across regions.

Blanchflower and Oswald (1994) use Equation (1) to discriminate between the wage curve and the Phillips curve. If a = 1, then wage growth is a function of the log of unemployment following the Phillips curve, whereas a = 0 implies that the log of wages is a function of the log of unemployment following the wage curve. Estimating Equation (1) using regional micro data for several OECD countries, Blanchflower and Oswald report estimates of a that are less than 0.3 and conclude that the evidence supports the wage curve specification.

The results of Blanchflower and Oswald have not gone unchallenged. Blanchard and Katz (1997), Black and FitzRoy (2000) and Card (1995) argue that Blanchflower and Oswald's (1994) results are partly an outcome of the use of inappropriate data for the US. Albaek *et al* (2000) find that Blanchflower and Oswald's results for Norway are due to the exclusion of fixed effect dummies. Blanchard and Katz (1999) argue that at the state level wages are not only likely to depend on lagged wages but, due to interstate labour mobility, also on lagged aggregate wages. Since the effects of lagged aggregate wages will be captured by the time-dummies, a will be biased downwards. Using British county data, Black and FitzRoy (2000) find that the Blanchflower and Oswald results are overturned when normal hourly wages, as opposed to earnings, are used, because earnings are influenced by cyclical fluctuations in overtime hours and therefore negatively correlated with unemployment. Using hourly wages they find evidence in favour of the Phillips curve.

Bell (1996) questions the assumptions of common productivity trends and consumer prices across regions. Diverse growth in productivity and consumer prices across regions will lead to an omitted variable bias. If these omitted variables are serially correlated, then the estimates of a will be biased. Card (1995) argues that technical problems associated with the presence of lagged dependent variables and fixed effects render the Blanchflower-Oswald specification inappropriate. Whelan (1997) argues that micro wage equations do not have any observable implications for macro data on wage and price inflation and concludes that "economists are ill-served when conclusions about macroeconomics are drawn from micro data without consideration of the restrictions imposed by aggregation" (p 18). Bell *et al* (2000) find that the coefficient of lagged wages is close to one but reduced substantially when a state-specific wage trend is added to the regression for the UK. This result is of great concern because it suggests that the results are highly sensitive to the inclusion of deterministic variables, including fixed effect dummies, as also found by Albaek *et al* (2000).

Another serious problem associated with estimates of Equation (1), is that estimates of a are biased and inconsistent regardless of data and specification, thus rendering estimates of Equation (1) an even more inappropriate tool for discriminating between the wage curve and the Phillips curve. The bias comes from two sources. First, the error terms are correlated with the fixed effects. For a first order autoregressive model Nickell (1981) shows that the least squares dummy variable (LSDV) estimator results in the following bias for a reasonably large number of values of T:

$$p\lim_{N\to\infty}(\hat{a}-a) \cong -\frac{1+a}{T-1}$$

where N is number of individuals (regions) and T is the number of time periods. Hence the LSDV estimator yields downward biased estimates of a, particularly when a is close to one and T is small. Since T is low in most estimates of Equation (1), and often below 10, the downward bias is non-trivial. Monte Carlo experiments have shown that the LSDV bias is also serious in moderately large samples (Kiviet, 1995, and Judson and Owen, 1999). For $T = 20$, Judson and Owen (1999) find that $\hat{a}$ is biased by –0.104 for $a = 0.8$. Hence, LSDV estimates will tend to reject the Phillips curve specification, regardless.

A second source of bias in estimates of a comes from the correlation between the lagged dependent variable and the error term. The sign and the magnitude of the resulting bias depends on the regressors and the size of the sample, and therefore have to be determined by Monte Carlo experiments using the original data set. However, the bias is generally largest the closer a is to one, and in simple autoregressions with or without constant terms and time trends, a is biased downwards when it is one because estimates of a are skewed to the left, particularly in small samples (see Andrews, 1993, and Hamilton (1994, pp 486-501) for proofs).

[3] Blanchflower and Oswald (1994) use data for individuals. However, since the effective number of observations is number of regions times years, Equation (1) can be used as a simplification.

Using instruments for the lagged dependent variable or estimating in first differences, to remove the fixed effects, will alleviate the problem of biased estimates of a; however, Monte Carlo simulations by Kiviet (1995) show that various transformations of the data and alternative estimators, including the GMM estimator, do not the resolve the bias and consistency problems, because of the correlation between the error terms and the fixed effects in small samples.

Overall these considerations suggest that estimates of Equation (1) using regional data are sensitive to inclusion of variables, measurement of variables, and specification, and are biased against the Phillips curve hypothesis, especially for small T and when a is close to one. Furthermore, the effective number of observations in the majority of the micro data sets is below 100, thus rendering the power of the tests relatively low. Given that supply shocks have persistent effects on unemployment when a = 1, the crucial issue is not whether a is close to 0 (wage curve) or 1 (Phillips curve) but whether a is 1 or not. This requires that the estimates of a are strictly unbiased, consistent and very efficient. Hence, an alternative test is called for.

3 Factor shares and wage formation

To overcome the problems that are associated with estimates of Equation (1), Blanchard and Katz (1997, 1999) and Johansen (1995) have suggested an alternative test to discriminate between the two models based on an error-correction framework using aggregate data. They suggest that whereas the wage curve implies mean reversion in labour's income share, the Phillips curve does not. This section shows that this test needs to be extended to enable discrimination between the two models because they both predict mean reversion in labour's income share.

Consider a model that nests the Phillips curve and the wage curve as follows:

$$w_t = f + aw_{t-1} + p_t - ap_{t-1} + (y-l)_t - a(y-l)_{t-1} - bu_t + z_t^{'}k - az_{t-1}^{'}k \qquad (2)$$

where p is the log of consumer prices or the value added price deflator, y is the log of output and l is the log of hours worked, z is a vector of the log of wage push variables, and k is a vector of coefficients associated with the wage push variables. Price and productivity homogeneity is assumed and the influence of demand shocks on wages is suppressed by assuming equality between price expectations and actual prices. If a = 0, then Equation (2) collapses to a wage curve, and if a = 1, reduces to a Phillips curve.

Equation (2) can be rewritten as:

$$\Delta w_t = f + \Delta p_t + \Delta(y-l)_t - bu_t + \Delta z_t^{'}k - (1-a)[w_{t-1} - p_{t-1} - (y-l)_{t-1} - z_{t-1}^{'}k] \qquad (3)$$

This error correction model resembles the equation estimated by Blanchard and Katz (1997), Johansen (1995), and OECD (1997) in order to discriminate between the wage curve and the Phillips curve. Restricting k = 0 in the error-correction term (the bracket parenthesis), they find that the maintained hypothesis of a zero coefficient of the error correction term cannot be rejected for the US, and conclude that the Phillips curve applies to the US labour market. Based on the evidence provided by OECD (1997), Blanchard and Katz (1999) note that (1 - a) is around 0.25 for most European countries and suggest that supply shocks have persistent unemployment effects in these countries.

In the following it will be shown that the coefficient of the error-correction term cannot be used to discriminate between the wage curve and the Phillips curve regardless of whether or not k is restricted to zero. To demonstrate this two cases are considered, namely when k is restricted to zero and when it is not.

3.1 When k is restricted to zero

Rewriting Equation (3) under the assumption that k = 0 in the error-correction term, yields:

$$\Delta s_t^L = f - b\,u_t + \Delta z_t^{'}k + (a-1)s_{t-1}^L, \qquad (4)$$

where S^L = WL/YP is the share of labour in total income and s^L = ln(S^L). This first order differential equation has a solution that depends on the magnitude of a. Three cases can be distinguished.

If a = 0 (wage curve), then Equation (4) becomes:

$$s_t^L = f - b\,u_t + \Delta z_t^{'}k\ . \qquad (5)$$

If Δz_t is stationary but u_t contains a unit root, then s^L will contain a unit root. Based on Equation (3) with the restriction k = 0 in the error-correction term, it can be concluded that the Phillips curve applies, when in fact the wage equation is the correct specification. However, if the rate of unemployment is mean reverting, then it may be erroneously concluded that the wage curve is the correct representation of the labour market depending on whether the wage push factors contain a unit root.

If a = 1 (Phillips curve), then labour's income share track the following dynamic path regardless of whether k is restricted to zero in the error-correction term:

$$s_t^L = s_0^L + \sum_{i=0}^{t-1} (f - b u_{t-i} + \Delta z_{t-i}^{'} k) \quad (6)$$

where s_0^L is labour's initial income share. Since demand shocks are suppressed here, it follows that $U_t = U_t^* = \exp[(f + \Delta z_t^{'} k)/b]$, under the assumption of the Phillips curve and perfect competition in the goods market as shown in the next section, where U^* is the natural rate of unemployment. Hence, the second right hand term in Equation (6) is zero and labour's income share will only deviate from its initial value due to demand shocks. From this it can be concluded that estimates of Equation (3) under the assumption that k = 0 in the error-correction term will yield mean reversion of the error correction term when the Phillips curve applies. The opposite conclusion would have applied if the Blanchard-Katz criterion was used.

If 0 < α < 1 the solution to Equation (4) is given by:

$$s_t^L = s_0^L a^t + \sum_{i=0}^{t-1} a^i [f - b u_{t-i} + \Delta z_{t-i}^{'} k]$$

Since the initial condition s_0^L is not known, the following solution is found using backward iteration:[4]

$$s_t^L = \frac{f}{1-a} + \sum_{i=0}^{t-1} a^i [\Delta z_{t-i}^{'} k - b u_{t-i}] . \quad (7)$$

From this equation it follows that the log of labour's share will converge to a constant mean, $f/(1 - a)$, iff Δz_t and u_t are individually or jointly mean reverting. If $\Delta z_t \to 0$ and $u_t \to f$ then it follows from Equation (7) that $\lim_{t\to\infty} s_t^L \to \frac{f(1-b)}{1-a}$. Finally, labour's income share will contain a unit root if unemployment contains a unit root.

In summary, the results in this section suggest that estimates of Equations (3) and (4) with k restricted to zero in the error correction term, cannot be used to discriminate between the wage curve and the Phillips curve, and may even give misleading results. The Phillips curve implies unconditional mean-reversion of the error-correction term, whereas the wage curve implies mean reversion conditional on mean-reversion of the log of unemployment.

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[4] The first right hand side term in Equation (7) is obtained from condition $t \to \infty$ so that $s_0^L a^t + \sum_{i=0}^{t-1} a^i f = \frac{f}{1-a}$.

3.2 When *k* is *not* restricted to zero

An alternative strategy is to relax the assumption of $k = 0$, which yields the error correction term from Equation (3):

$$v_{t-1} = w_{t-1} - p_{t-1} - (y-l)_{t-1} - z_{t-1}^{'}k \tag{8}$$

where í$_{t-1}$ is the error correction term. Since labour's income share is mean reverting for both the Phillips curve and the wage equation, as shown in the Appendix, it follows that í$_{t-1}$ is mean reverting iif *z* is mean reverting. Hence, the error correction term is mean-reverting iff *z* is mean reverting regardless of whether the Phillips curve or the wage equation applies. This implies that unrestricted estimates of Equation (2) to discriminate between the Phillips curve and the wage curve, is conditional on mean reverting behaviour of *z*.

Equation (8) resembles the wage curve if it is expanded with unemployment:

$$v_{t-1} = w_{t-1} - p_{t-1} - (y-l)_{t-1} - z_{t-1}^{'}k + bu_{t-1} \tag{9}$$

The wage curve implies that this equation is mean reverting since it assumes that *z* and *u* are cointegrated. However, mean reversion of this equation is also consistent with the Phillips curve hypothesis, because it predicts that *k* and $\hat{a}$ are zero under the assumption of perfect competition in the goods market. As shown in Section 5, *k* may not be zero under the joint hypothesis of imperfect competition and the Phillips curve. Hence, the statistical significance of the coefficient of the error correction term in estimates of Equation (2) cannot be used as a test to discriminate between the Phillips curve and the wage curve even if unemployment is added as a regressor in the error correction term.

4 Implications for the natural rate and the persistence of shocks

As discussed in Blanchard and Katz (1997, 1999), the magnitude of *a* is important for the persistence of supply shocks on unemployment. Suppose that prices are set as a mark-up on unit labour costs in a perfectly competitive goods market, that expectations are borne out, and that long-run price and productivity homogeneity exist. Then Equation (3) implies the natural rate of unemployment as follows:

$$\ln U_t^* = \frac{f + (1-a)z_t^{'}k + a\Delta z_t^{'}k}{b}\,. \tag{10}$$

Under the assumption of the Phillips curve ($a = 1$), the natural rate is given by:

$$U_t^* = \exp[(f + \Delta z_t^{'} k)/b],$$

and supply shocks have only one-off effects on unemployment. Under the assumption of the wage curve (a = 0), the natural rate is given by:

$$U_t^* = \exp[(f + z_t^{'} k)/b].$$

Hence, supply shocks will permanently influence unemployment, $\partial \ln U^* / \partial z_j = k_j / b$, where j is the j'th element in the k-vector. This implies that cointegration between u_t and z_t is a necessary, but *not* a sufficient condition for the wage curve to apply. It is not a sufficient condition, because u_t and z_t are also cointegrated in the intermediate case where $0 < a < 1$. This can be seen as follows.

In the intermediate case, where $0 < a < 1$, the solution to Equation (10) is:

$$\ln U_t^* = \frac{f}{(1-a)b} + \sum_{i=0}^{t-1} \frac{a^i}{b} [a \Delta z_{t-i}^{'} k + (1-a) z_{t-1-i}^{'} k].$$

From the equation it follows that unemployment and the z-variables are cointegrated, but that autoregression will be more pronounced here than in the situation where a = 0 (wage curve). From an empirical point of view this result is rather unfortunate because it renders it very difficult to distinguish between the wage curve and the intermediate case, if the Phillips curve is rejected.

The magnitude of a is not only important for the persistence of supply shocks on unemployment and output but also important for the distributional effects of supply shocks. In the case where $0 \le a < 1$, workers will be partially compensated for higher taxes and adverse terms-of-trade shocks, and will gain in real terms from higher unemployment benefit replacement ratios. However, in the Phillips curve framework workers bear the whole burden of higher labour taxes, higher direct and indirect taxes, and higher non-tax indirect labour costs.

5 Labour's income share and the natural rate under imperfect competition

The exposition in the previous section suggests that wage push factors have permanent effects on wages under the wage curve hypothesis, and only temporary effects on wages under the Phillips curve hypothesis. This should, in principle, make it easy to discriminate between the two hypotheses. However, this section shows that the supply shocks can have persistent effects on wages in the Phillips curve framework under imperfect competition, which makes it even more difficult to empirically discriminate between the models in small samples.

There is substantial evidence suggesting that mark-ups are negatively related to supply shocks and positively related to prices charged by competitors. The pricing-to-market literature consistently finds mark-ups to be positively related to the exchange rate, measured as the domestic currency price of foreign exchange (Dornbusch, 1987). Similarly, substantial empirical and theoretical literature suggests that firms do not pass on higher costs to prices in the short run (Ball *et al*, 1988). It is therefore not surprising that estimates of the wage elasticity of prices are consistently below one in first-difference estimates of price equations. The estimated coefficient of wages is simply biased downwards because mark-ups are corrected with wages.

Allowing mark-ups to be inversely related to wage push factors yields the price equation as follows:

$$\Delta p_t = \Delta w_t - \Delta(y-l)_t - l\Delta z_t^{'}k \tag{11}$$

where l is a constant, $0 \le l \le 1$, and measures the extent to which mark-ups are reduced following an adverse supply shock. If l = 1, then mark-ups counterbalance supply shocks on a one-to-one basis, and supply shocks do not affect unemployment. If l = 0, then mark-ups are unaffected by supply shocks.

Solving Equations (2) and (11) jointly yields the natural rate as follows:

$$\ln U_t^* = \frac{f + (1-l)(1-a)z_t^{'}k + (1-l)a\Delta z_t^{'}k}{b} \tag{12}$$

which shows that the influences of supply shocks on the natural rate are muted by the factor l. Intuitively, the leftward shift in the wage setting schedule following an adverse supply shock is exactly counterbalanced by a rightward shift in labour demand following the decrease in mark-ups.

Under the assumption of the Phillips curve (a = 1) Equation (12) reduces to:

$$U_t^* = \exp\frac{1}{b}[f + (1-l)\Delta z_t^{'}k]. \tag{13}$$

Setting $U_t = U_t^*$ in Equation (13) and inserting into Equation (3) yields, after some manipulations:

$$s_t^L = s_0^L + l \sum_{i=0}^{t-1} \Delta z_{t-i}^{'} k \ . \qquad (14)$$

This equation shows that supply shocks have permanent effects on labour's income share if l = 1, because unemployment is independent of supply shocks.

In the more realistic case where 0 < l < 1, supply shocks will not permanently influence labour's income share, but give rise to autocorrelation in labour's income share, and therefore renders it difficult to find mean reversion of income shares in small samples. This prediction is consistent with the observation that labour's income share often deviates from its long run equilibrium over prolonged periods. An implication of this result is that empirical tests using small samples will tend to reject the hypothesis of mean reversion in labour's income share if l is close to one, even if the Phillips is the correct specification.

Under the assumption of the wage curve (a = 0), we get the joint solution to Equations (3) and (12) under the assumption that $U_t = U_t^*$ as follows:

$$s_t^L = l z_t^{'} k \ . \qquad (15)$$

Hence, labour's income share follows the time profile of supply shocks if l > 0. Mean reversion in labour's income share is therefore conditional on mean reversion of the wage push factors.[5] This result is important because it shows that estimates of Equation (3) under the assumption that k = 0 in the error-correction term will not yield mean reversion of the error correction term under the joint hypothesis of the wage curve and imperfect competition in the goods market. Hence, the Blanchard-Katz test will reject the wage curve hypothesis when the wage curve hypothesis is correct.

6 Empirical evidence on the wage curve versus the Phillips curve

The results in the previous sections show that the statistical significance of the coefficient of the error-correction term cannot be used to discriminate between the wage curve and the Phillips curve. Furthermore, it was shown that it is difficult to discriminate between the two hypotheses in small samples. This calls for tests that have high power and which do not rely on the statistical significance of the error-correction term. This section seeks to discriminate between the two hypotheses of wage formation using panel estimates of cointegration and error correction estimates for 18 OECD countries over the period from 1952 to 1999, which gives 864 effective observations.

6.1 Cointegration estimates

The following cointegration model is estimated for OECD manufacturing:

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\begin{aligned}
w_{it} = {} & a_{0i} + a_1 p_{it}^{va} + a_2 pr_{it} + a_{3i} u_{it} + a_4 tot_{it} + a_5 rr_{it} + a_6 t_{it}^{d} + a_7 pr_{it}^{x} + a_8 p_{it}^{x} \\
& + a_9 TD_t^{'} w + t_t^{'} v \; + n_{it}
\end{aligned} \tag{16}
$$

where the subscript i signifies country i. Here, pr is the log of manufacturing labour productivity, $(y - l)$, p^{va} is the log of the manufacturing value added price deflator, tot is the log of the ratio of the economy-wide and the manufacturing value added price-deflators, t^d is the log of the ratio of direct taxes and nominal GNP, pr^x is the log of the ratio of labour productivity in manufacturing and the whole economy, rr is the real interest rate measured in decimal points, and is calculated as the interest rate on a long-term government bond minus contemporaneous consumer price inflation, p^x is the log of the ratio of consumer prices and the economy-wide value added price-deflator, wi is the log of indirect labour costs as a percentage of total labour costs, t is a time trend, which is assumed to capture the effects of omitted variables on wages, v is a stochastic disturbance term, and a_1-a_9, v , and w are fixed parameters. The coefficients of the value-added price deflator and labour productivity for manufacturing are restricted to one following the natural rate hypothesis.[6]

Equation (16) is a standard wage equation. The wage push variables represent the wage push factors in the most important models of unemployment (see Bean, 1994, and Madsen, 1998, for a discussion). The variables pr^x and p^x are included to allow for spill-over effects from the whole economy wages to manufacturing wages due to discrepancies in productivity and price advances. The coefficients of pr^x and p^x are expected to be approximately the same but of opposite sign.[7] These variables are not wage push factors because they cannot be operative on a nationwide scale, otherwise the decline in the warranted wages since the mid 1970s, would have been too pronounced. Finally, the time-

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[5] Almost the same result is reached in the intermediate case where $0 \le a < 1$ and is therefore not shown.

[6] The null hypothesis of unity coefficients of the log of labour productivity and the log of the value-added price-deflator could not be rejected at the 1-percentage level [$c^2(2) = 5.64$].

[7] The following other important wage push factors were included in the estimates over a shorter estimation period (1961 to 1999) because they are not available from an earlier date for most countries: the unemployment benefit replacement ratio, the duration of unemployment benefits, the proportion of population of age between 20 and 24 years, and various mismatch variables (see Madsen, 1998, for construction of the variables). The estimated coefficients were of wrong sign except the estimated coefficient unemployment benefit replacement ratio. The estimated coefficient of the unemployment replacement ratio was obtained from another regression where only every second year data from 1961 to 1995 was used because internationally comparable data for unemployment benefits are only available from the OECD over this period. Estimating Equation (16) over the period from 1961 to 1995, omitting equal years, yields an estimated coefficient of the unemployment benefit replacement ratio of 0.0025(0.26) if the replacement ratio is measured in logs and of 0.00067(2.12) if the replacement ratio is measured in levels, where the numbers in parentheses are t-statistics. Hence, the influence of unemployment benefits on wages is at best marginal.

dummies capture the effects of omitted variables that follow the same time-profile across countries.

Estimates of Equation (16) can be used to discriminate between the two models of wage formation. The wage equation predicts that unemployment and wage push factors form a positive cointegrating relationship and that the coefficient of unemployment is negative. The Phillips curve predicts that wages are unrelated to wage push factors because supply shocks have only temporary effects on wages; and that wages are unrelated to unemployment.

Estimation method

Equation (16) was first estimated for each individual country. However, the coefficient estimates varied substantially across nations, were often of the wrong sign, and were mostly statistically insignificant.[8] Furthermore, none of the estimates were cointegrated at conventional significance levels, which is likely to reflect the low power of cointegration tests in small samples as well as omitted variables. The problem of low power of the tests in small samples is further aggravated by the fact that supply shocks have persistent effects on wages under imperfect competition in the goods market under the Phillips curve hypothesis.

To gain efficiency and enhance the power of the tests, Equation (16) is estimated using pooled cross section and time-series analysis. Major advances in the asymptotic theory of panels with integrated variable made by Phillips and Moon (1999, 2001) has made it possible to evaluate the consistency of panel data estimates and the distributional properties of estimators. Based on Phillips and Moon (1999), Phillips and Moon (2001) show the very strong result that panel OLS estimates of long-run relationships yield consistent coefficient estimates, even if the variables are not cointegrated. Phillips and Moon (1999) furthermore demonstrate that under weak regularity conditions, pooled OLS estimates of the coefficients are $\sqrt{n}$-consistent for the long-run average coefficients and have a limiting normal distribution. Banerjee (1999) argues that cointegration estimation in panels eliminates several problems that are associated with individual cointegration estimates, particularly the problem of a small sample bias.

How then, can one discriminate between potential cross-country differences in labour market behaviour from pooled estimates? Since unemployment plays a key role in the estimates to discriminate between the two competing hypotheses of wage behaviour, the coefficient of unemployment is allowed to vary across countries. The time trends are also allowed to vary across countries and they are restricted to be the same for country groups with similar coefficients following a stepwise procedure.[9] However, allowing all other coefficients to vary across countries amounts to estimating single country models and the power of the tests will not be enhanced by pooling.[10,11]

Data

Equation (16) is estimated for 18 OECD countries over the period from 1952 to 1999. The country sample is listed in the notes to Table 1. Manufacturing data are used because economy-wide data give a misleading picture of the path of factor shares, and hence the path of wages, productivity and value-added prices. This is mainly because economy-wide factor shares are heavily influenced by changes in sectoral compositions and changes in the proportion of self-employment. Income from self-employment is categorised as property income in national accounts. Since the proportion of self-employment varies substantially among sectors, it follows that the economy-wide S^L is influenced by changes in sectoral compositions. Furthermore, the fraction of self-employed declined substantially prior to the 1970s and hence artificially increased labour's income share (Chan-Lee and Sutch, 1985). Since the manufacturing sector has a low fraction of self-employed, it has not been significantly affected by the declining self-employment (Chan-Lee and Sutch, 1985).

Another problem associated with economy-wide income shares is that the increasing share of governmental services, such as investment in infrastructure and provision of

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[8] The test statistics were transformed to follow a standard normal distribution using the method of Stock and Watson (1993) and Phillips and Loretan (1991), and first-differences of the regressors were included in the estimates to allow for dynamic adjustment. The results are available from the author.

[9] Equation (16) was first estimated allowing the coefficients of the time trend to vary across all countries. The two lowest coefficient estimates were then tested for equality and restricted to be the same if the null hypothesis could not be rejected at the 1% level, otherwise coefficient equality was tested for the two countries with the second and the third lowest coefficients and merged if they were the same, and so forth.

[10] One could alternatively have started from a completely unrestricted model and then stepwise have restricted the coefficients to be the same across nations to the extent that the statistical tests allow them to be restricted to be the same. However, this procedure gives the problem of how to restrict the high proportion of implausible coefficient estimates. For example, how should coefficient estimates of wage push variables that are negative, or have elasticities that are in excess of one, be treated? Clearly, there is no objective way of dealing with these issues except by restricting the coefficients to be the same across countries since it yields coefficients that are unbiased even if countries have different coefficients. Hall *et al* (1999) demonstrate that if regressors for each individual (country) in the panel are driven by common stochastic trends, and that the variables for each individual are cointegrated, then consistent estimates are obtained in panels that impose coefficient homogeneity, even if the true model has heterogeneous coefficients.

[11] To investigate the sensitivity of the estimates to less restrictive assumptions, the coefficients were estimated under the assumption that they are the same in the countries that have experienced the smallest increase in the rate of unemployment over the period from 1970 to 1999. Under the maintained hypothesis of Blanchard and Katz (1999), among others, that unemployment is higher in some countries than others because supply shocks have had more persistent effects on unemployment in these countries, it follows that the influence of supply shocks on wages is proportional to the influence on unemployment. The following countries have had a distinctly smaller increase in unemployment than other OECD countries in the sample, and are therefore restricted to have the same coefficients: Canada, USA, Japan, Austria, Ireland, Netherlands, Norway and the UK. However, the null hypothesis that the coefficient estimates for this country group differ from the other country group cannot be rejected at the 1%. Hence, the coefficients were restricted to be the same, except the coefficients of unemployment and the time trends.

schooling, has artificially contributed to an increase in labour's income share because these governmental services do not earn operating surplus. Furthermore, there is also a strong argument in favour of excluding returns to real estate and financial services since they are imputed in national accounts, and profits of financial institutions include the seigniorage earned by the central bank. These considerations suggest that labour's income shares in manufacturing are better suited for long-run analysis than economy-wide income shares.

Estimation results

The results of estimating Equation (16) are presented in Table 1. The estimated coefficients of the "fixed effect" dummies are not shown because they are dominated by scaling effects. The null hypothesis of no cointegration is rejected at the 1% level.[12] Hence, the *t*-statistics approximately follow the standard normal distribution. Some of the estimated coefficients of the time-dummies were insignificant at the 1-percentage level and the associated time-dummies were consequently deleted to simplify the presentation in Table 1. The estimated coefficients of p^x and pr^x are 0.49 and -0.44, respectively, and are statistically highly significant. This suggests that sectoral wage spillover effects play an important role for wage determination, but at the same time indicates that wages are also influenced by idiosyncratic productivity and price advances.

Table 1. Parameter estimates of Equations (17) and (18).

w_t				Δw_t [1]			
t_t^d	0.19(12.6)	TD_{1968}	0.05(4.50)	Δw_{t-1}	0.19(6.20)	u_t^{Spa}	-0.78(3.35)
tot_t	0.40(8.78)	TD_{1969}	0.04(3.30)	Δp_t^{va}	0.45(9.68)	u_t^{Swe}	-0.79(5.50)
rr_t	0.04(0.59)	TD_{1970}	0.05(4.50)	Δp_{t-1}^{va}	0.21(4.40)	u_t^{UK}	-0.53(5.24)
pr_t^x	-0.44(23.8)	TD_{1971}	0.06(5.95)	Δpr_t	0.40(14.0)	Δu_t	0.21(2.75)
wi_t	-0.01(2.50)	TD_{1972}	0.07(6.28)	Δpr_{t-1}	0.09(2.89)	TD_{1961}	0.02(3.45)
p_t^x	0.49(19.3)	TD_{1973}	0.08(6.56)	Δp_t^x	0.31(6.81)	TD_{1972}	0.02(3.02)
u_t^{Can}	-0.02(1.21)	TD_{1974}	0.09(7.37)	Δp_{t-1}^x	0.33(6.98)	TD_{1973}	0.03(4.93)
u_t^{USA}	0.11(4.42)	TD_{1975}	0.11(9.09)	Δt_t^d	0.11(8.77)	TD_{1974}	0.06(9.84)
u_t^{Jap}	0.20(8.70)	TD_{1976}	0.11(8.86)	Δt_{t-1}^d	0.02(1.22)	TD_{1975}	0.03(5.63)

[12] Maddala and Wu (1999) recommend the p_I test based on Fisher (1932) to test for cointegration and unit roots in panels. The test is based on the *p*-values of the test-statistics for cointegration for each cross-sectional unit and is given by:

$$p_I = -2\sum_{i=1}^{N}\ln(p_i)$$

which is distributed as a chi-squared variable with 2*N* degrees of freedom under the assumption of cross-sectional independence. p_i is the *p*-value of the test statistic for individual *i*. The *p*-values are based on Dickey-Fuller tests.

u_t^{Aud}	0.02(1.71)	TD_{1977}	0.10(8.71)	Δtot_t	-0.13(3.01)	*Const*	-0.01(3.61)
u_t^{NZ}	0.01(1.64)	TD_{1978}	0.10(8.60)	Δtot_{t-1}	0.19(4.24)		
u_t^{Aut}	0.08(5.78)	TD_{1979}	0.11(8.96)	Δpr_t^x	-0.21(7.42)		
u_t^{Bel}	0.03(2.70)	TD_{1980}	0.09(8.05)	Δpr_{t-1}^x	-0.22(7.67)	N	828
u_t^{Den}	-0.02(2.34)	TD_{1981}	0.06(5.09)	Δwi_t	0.01(2.43)	R^2(mom)	0.94
u_t^{Fin}	0.02(2.17)	TD_{1982}	0.04(3.73)	Δwi_{t-1}	0.00(0.03)	DW(M)	2.03
u_t^{Fra}	0.06(6.50)	TD_{1983}	0.04(3.73)	$\hat{v}_{t-1}$	-0.16(8.82)	F(648,162)	0.22
u_t^{Ger}	0.08(11.7)	TD_{1984}	0.05(4.26)	u_t^{Can}	-0.74(3.84)	*Chow*(39,750)	0.89
u_t^{Ire}	0.02(1.16)	TD_{1985}	0.05(4.37)	u_t^{USA}	-0.62(3.98)		
u_t^{Itl}	0.09(4.04)	TD_{1986}	0.04(3.69)	u_t^{Jap}	-0.66(4.46)		
u_t^{Net}	-0.01(0.62)	TD_{1987}	0.04(3.33)	u_t^{Aud}	-0.71(4.48)		
u_t^{Nor}	-0.01(0.72)	TD_{1988}	0.06(4.82)	u_t^{NZ}	-0.45(3.29)		
u_t^{Spa}	-0.16(17.1)	TD_{1989}	0.07(6.20)	u_t^{Aut}	-0.62(4.33)		
u_t^{Swe}	0.02(1.59)	TD_{1990}	0.06(5.26)	u_t^{Bel}	-0.70(4.68)		
u_t^{UK}	0.02(4.01)	TD_{1991}	0.04(4.43)	u_t^{Den}	-0.93(5.69)		
TD_{1961}	0.03(2.65)	TT_A	-0.02(27.1)	u_t^{Fin}	-0.81(5.03)		
TD_{1962}	0.05(4.48)	TT_B	-0.01(19.8)	u_t^{Fra}	-0.83(6.06)		
TD_{1963}	0.06(5.23)	TT_C	-0.002(6.4)	u_t^{Ger}	-0.68(4.99)		
TD_{1964}	0.06(5.02)			u_t^{Ire}	-0.78(2.95)		
TD_{1965}	0.06(4.78)	$\bar{R}^2$	1.00	u_t^{Itl}	-0.92(3.79)		
TD_{1966}	0.06(5.21)	N	864	u_t^{Net}	-0.85(6.13)		
TD_{1967}	0.06(4.89)	$c^2(36)$	98.3	u_t^{Nor}	-0.84(4.62)		

Notes: Absolute t-statistics are given in parentheses. R^2(mom) = Buse's R-squared. N = number of observations. DW(M) = modified Durbin-Watson test for first order serial correlation in fixed effect panel data models (see Bhargava *et al*, 1982). *Chow*(i,j) = F test for coefficient constancy with breaking point in 1975/1976, and is distributed as F(i,j) under the null hypothesis of structural stability. F(i,j) = F-test for cross-country coefficient constancy, and is distributed as F(i,j) under the null hypothesis of coefficient constancy. $c^2(36)$ is Fisher's p_l test for cointegration, and is distributed as chi-squared with 36 degrees of freedom under the null hypothesis of no cointegration (see Maddala and Wu, 1999). TT_A = time trend for Ireland, TT_B = time trend for the USA, Japan, New Zealand, Belgium, Finland and France, TT_C = time trend for Canada, Australia, Austria, Denmark, Germany and Sweden. The following instruments are used for Δp_t^{va} : Δp_{t-1}^{va} , Δp_{t-1}^{x} , Dir_t, Δir_{t-1}, $\Delta m1_{t-1}$, Δp_t^c , and Δp_{t-1}^c , where Δir is the nominal interest rate on a long-term government bond, $m1$ is the log of $M1$, and Δp^c is commodity prices. Estimation period: 1952-1999 (Equation (17)) and 1954-1994 (Equation (18)). The data are collected from different national and international data sources. A detailed list of data sources is available from the author. The country sample consist of the countries as follows: Canada, USA, Japan, Australia, New Zealand, Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, the Netherlands, Norway, Spain, Sweden, and the UK.

1: The estimated coefficient of unemployment is multiplied by 100.

The estimates do not give support to the wage curve specification for two reasons. First, the estimated coefficient of unemployment is only significantly negative for Spain and significantly positive for eight countries, at the 1% level, and therefore points towards a spurious relationship for Spain.[13] Second, the only statistically significant wage push variables are direct taxes and terms-of-trade suggesting that these are the only supply variables that have permanent effects on real wages and hence unemployment in the OECD countries. Since the *tot* variable has remained almost constant in OECD countries on average in the sample period, and therefore has not shown a positive trend, the only potentially important wage push variable is direct taxes. However, simulations of the model suggest that the increase in direct taxes from 1970 to 1999 has only contributed to 1.5-percent increase in wages over the same period, and has therefore not been quantitatively important. Weighing this result against the insignificance of other wage push variables, particularly the unemployment benefit replacement ratio and the duration of unemployment benefits, the evidence for a wage curve is weak.

The estimated coefficients of the time-dummies are highly significant and have a time profile we would expect from the taxonomy in the literature, namely an increase in labour's income share of 11% from 1960 to a peak in the second half of the 1970s, and a declines back to its 1960 level by 1992 (Madsen, 1998). The time-profile of the time-dummies is likely to reflect the omission of a variable that has had temporary effects on wages for three reasons. First, if the present high unemployment in several OECD countries is due to excessive wages, then the time-profile of the time-dummies would *not* have been bell-shaped, but should have been increasing parallel with the increasing unemployment. Second, the wage curve predicts that wage push factors and unemployment form a positive cointegration relationship. However, since the estimated coefficients of unemployment are mostly positive, the wage push variables cannot simultaneously form a positive cointegration relationship with the wage push variables and be positively related to wages.

Third, there is substantial evidence suggesting that the increasing union activity in the late 1960s and the beginning of the 1970s was partly responsible for the wage explosion in the same period (Bruno and Sachs, 1985). By contrast, the unions have gradually weakened since the beginning of the 1980s (Bean, 1994), and this has probably contributed to the decline in labour's income share over the same period. Assuming a constant natural rate of unemployment, the wage reducing effects of higher unemployment, following the Phillips curve may have contributed to declining income share of labour. If labour's income share has been predominantly driven by the strengths of unions and unemployment, and not by exogenous forces, it follows that wage push factors, which have permanent wage effects,

[13] These results are not a result of the pooling of the data. Single country estimates gave almost the same results.

cannot have been omitted variables that have permanent effect on wages. The evidence in the next sub section supports this result.

6.2 Error-correction estimates

The dynamic counterpart of Equation (3) is estimated by allowing for sluggish adjustment of wages to price and productivity shocks. The following error-correction model is estimated, where lags are omitted for simplicity:

$$\begin{aligned}\Delta w_{it} &= b_i + b_1 \Delta p_{it}^{va} + b_2 \Delta pr_{it} + b_3 \Delta p_{it}^{x} + b_4 \Delta t_{it}^{d} + b_5 \Delta tot_{it} \\ &\quad + b_6 \Delta pr_{it}^{x} + b_7 \Delta wi_{it} + b_8 \Delta u_{it} + u_{it}^{'} t - (1-a)\hat{v}_{i,t-1} + TD_t^{'} V + e_{it} \qquad (17)\end{aligned}$$

where $n_{t\text{-}1}$ is the error correction term, which is the lagged residual from the estimates of Equation (16). The real interest rate is not included because it interferes with the dynamic adjustment of wages to changes in the prices via the inflation component in the real interest rate. Including the real interest rate results in an estimated coefficient of the real interest rate that is significantly negative and a very slow adjustment of wages to changes in prices. Instruments are used for p_{it}^{va} and these are listed in the notes to Table 1. One period lags of regressors and the regressant, except u and Δu, are included in the estimates of the wage equation. Note that the coefficients of u are allowed to vary across countries. The generalised instrumental variable method, where the covariance matrix is weighted by the correlation of the disturbance terms, is used.[14]

The wage equation predicts that wage growth is negatively related to the *change* in the log of the rate of unemployment, whereas the Phillips curve predicts that wage growth is negatively related to the log of the *level* of the unemployment rate. Note that the coefficient of the error-correction term cannot be used to discriminate between the wage curve and the Phillips curve hypotheses, as shown in the sections above.

The results of estimating Equation (17) are shown in the right hand side of Table 1. The F-test for pooling is not significant at any conventional significance level [$F(648, 162) = 0.26$], suggesting that wage behaviour is extraordinarily similar across countries. The coefficients of unemployment are consistently significant and negative for all countries, suggesting that the Phillips curve applies to all OECD countries in this study. The estimated coefficient of the log of unemployment is, on average, –0.0074, which implies that the

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[14] More specifically the following variance-covariance structure is assumed: $E\{e_{it}^2\} = s_i^2$, $i = 1, 2, \ldots N$, and $E\{e_{it}, e_{jt}\} = s_{ij}$, $i \neq j$, where s_i^2 is the variance of the disturbance terms for country $i = 1, 2, \ldots N$, s_{ij} is the

approximately 8% unemployment in the OECD countries on average over the past 5 years, annually lowers the growth in total labour costs by 1.5%. Thus, unemployment will relatively quickly eliminate the wage effects of supply shocks.

The estimated coefficients of unemployment are remarkably similar across the OECD countries. This is an important result because it suggests very similar wage behaviour and the same speed of adjustment towards equilibrium due to shocks, in the OECD countries. This result is a big challenge to the large and influential union related literature that stresses cross-country variations in the labour market institutions and structures of the unemployment benefit system as leading candidates to explain cross-country differences in unemployment.

The estimated coefficient of Δu_t is significantly positive, which goes against the predictions of the wage curve hypothesis of a negative coefficient of Δu_t. If allowed to vary across countries, the estimated coefficient of Δu_t is consistently positive, even for Spain, which had a significantly negative coefficient of the log of unemployment in the cointegration estimates. Coupled with the finding of significantly negative estimates of the coefficients of u_t, the evidence is highly favourable to the Phillips curve hypothesis.

The estimates show that wages are relatively slow to adjust to innovations in prices, productivity, wage push factors, and the wedge between manufacturing and economy-wide value added prices and labour productivity. Wages adjust less than 50% within a year and less then 75% within two years to innovations in manufacturing productivity and the value added price-deflator. The wage adjustment to supply shocks is much slower, which is consistent with the finding in Section 5 that supply shocks are slow to die out under the joint assumptions of imperfect competition and the Phillips curve.

Finally, the estimated coefficients of the time-dummies are consistent with the cointegration estimates, and highlight the wage explosion in the first half of the 1970s, which was probably due to the strong union activity as stressed by Bruno and Sachs (1985). The estimated coefficients show that wages increased by 14% over this period due to factors that were common across countries, but not explained by the wage push variables included in the estimates. The wage disinflation from 1979 to 1991, as identified in the cointegration estimates, has probably been too gradual to significantly show up in the estimates of the error correction model. Some coefficients were significant at the 5% level, but not at the 1% level, and therefore restricted to zero.

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covariance of the disturbance terms across countries i and j, and e is the disturbance term. s_i^2 and s_{ij} are estimated using the feasible generalized least squares method described in Greene (2000, Ch. 15).

7 Concluding remarks

This paper has challenged the finding of Blanchflower and Oswald that the wage curve may be close to an empirical law of economics and that the Phillips curve is inherently wrong (1994, p 361). It was argued that the Blanchflower and Oswald model is biased in favour of the wage curve, has too few effective observations to give reliable results, and that the results are to sensitive too estimator and model specification. It was furthermore shown that the model suggested by Blanchard and Katz (1997) cannot be used to discriminate between the Phillips curve and the wage curve because both hypotheses have the same error-correction term predictions and are sensitive to the time-series properties of the wage push variables and the degree of imperfect competition in the goods market.

Extending the model of Blanchard and Katz (1997, 1999), models evidence from 18 OECD countries over half a century gives strong support for the Phillips curve and suggests that there is no macro evidence of the wage curve in the OECD countries. First, the cointegration estimates revealed that the log of wages is predominantly positively related to the log of the rate of unemployment, which stands in strong contrast to the predictions of the wage curve. Furthermore, supply shocks were not found to have persistent effects on wages as predicted by the wage curve. Second, the error correction estimates showed that wage growth was consistently and significantly related to the level of unemployment for the OECD countries as predicted by the Phillips curve. Furthermore, the estimated coefficients of unemployment were remarkably similar across countries, suggesting that labour market institutions and unemployment benefit incentive structures are much less important for labour market flexibility than they are traditionally thought to be.

The findings of the paper have two important macroeconomic implications. First, that supply shocks have only temporary effects on unemployment, and second, that the labour market is best represented by the competitive model of the labour market. The latter is a great challenge to conventional new Keynesian theories of unemployment and suggests that goods market imperfections or long persistence of demand shocks may be the key to explaining the different unemployment experiences in the OECD countries.

APPENDIX

This appendix shows that both the Phillips curve and the wage curve are consistent with mean reverting behaviour of labour's share in total income under the assumption of perfect competition in the goods market. Equation (2) can be reparameterised as follows:

$$\Delta s_t^L = f - bu_t + \Delta z_t^{'} k + (a - 1)(s_{t-1}^L - z_{t-1}^{'} k) \qquad \text{(A1)}$$

If a = 0 (wage curve), then Equation (A1) becomes:

$$s_t^L = f - b\,u_t + z_t^{'} k \ . \qquad \text{(A2)}$$

Since the wage curve theory implies that unemployment and the *z*-variables form a cointegration relationship, the wage curve is consistent with a constant share of income going to labour, and therefore consistent with mean reversion in labour's income share.

If a = 1 (Phillips curve), then labour's income share follows the dynamic path as follows:

$$s_t^L = s_0^L + \sum_{i=0}^{t-1} (f - b\,u_{t-i} + \Delta z_{t-i}^{'} k) \, . \qquad \text{(A3)}$$

where s_0^L is labour's initial income share. Since $U_t = U_t^* = \exp[(f + \Delta z_t^{'} k)/b]$ under the assumption of the Phillips curve, labour's share will only deviate from its initial value due to demand shocks.

If 0 < a < 1, the solution to Equation (A1) is given by:

$$s_t^L = \frac{f}{1-a} + \sum_{i=0}^{t-1} a^i [\Delta z_{t-i}^{'} k - (1-a) z_{t-1-i}^{'} k - b\Delta u_{t-i} - b(1-a) u_{t-1-i}] \, . \qquad \text{(A4)}$$

From this equation it follows that labour's share will converge to a constant mean, $f/(1 - a)$, iff z_t and u_t are cointegrated, as predicted by the wage curve.

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w_t				Δw_t			
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tot_t	0.40(8.78)	TD_{1969}	0.04(3.30)	Δp_t^{va}	0.45(9.68)	u_t^{Swe}	-0.79(5.50)
rr_t	0.04(0.59)	TD_{1970}	0.05(4.50)	Δp_{t-1}^{va}	0.21(4.40)	u_t^{UK}	-0.53(5.24)
pr_t^x	-0.44(23.8)	TD_{1971}	0.06(5.95)	Δpr_t	0.40(14.0)	Δu_t	0.21(2.75)
wi_t	-0.01(2.50)	TD_{1972}	0.07(6.28)	Δpr_{t-1}	0.09(2.89)	TD_{1961}	0.02(3.45)
p_t^x	0.49(19.3)	TD_{1973}	0.08(6.56)	Δp_t^x	0.31(6.81)	TD_{1972}	0.02(3.02)
u_t^{Can}	-0.02(1.21)	TD_{1974}	0.09(7.37)	Δp_{t-1}^x	0.33(6.98)	TD_{1973}	0.03(4.93)
u_t^{USA}	0.11(4.42)	TD_{1975}	0.11(9.09)	Δt_t^d	0.11(8.77)	TD_{1974}	0.06(9.84)
u_t^{Jap}	0.20(8.70)	TD_{1976}	0.11(8.86)	Δt_{t-1}^d	0.02(1.22)	TD_{1975}	0.03(5.63)
u_t^{Aud}	0.02(1.71)	TD_{1977}	0.10(8.71)	Δtot_t	-0.13(3.01)	*Const*	-0.01(3.61)
u_t^{NZ}	0.01(1.64)	TD_{1978}	0.10(8.60)	Δtot_{t-1}	0.19(4.24)		
u_t^{Aut}	0.08(5.78)	TD_{1979}	0.11(8.96)	Δpr_t^x	-0.21(7.42)		
u_t^{Bel}	0.03(2.70)	TD_{1980}	0.09(8.05)	Δpr_{t-1}^x	-0.22(7.67)	N	828
u_t^{Den}	-0.02(2.34)	TD_{1981}	0.06(5.09)	Δwi_t	0.01(2.43)	R^2(mom)	0.94
u_t^{Fin}	0.02(2.17)	TD_{1982}	0.04(3.73)	Δwi_{t-1}	0.00(0.03)	DW(M)	2.03
u_t^{Fra}	0.06(6.50)	TD_{1983}	0.04(3.73)	$\hat{v}_{t-1}$	-0.16(8.82)	F(648,162)	0.22
u_t^{Ger}	0.08(11.7)	TD_{1984}	0.05(4.26)	u_t^{Can}	-0.74(3.84)	*Chow*(39,750)	0.89
u_t^{Ire}	0.02(1.16)	TD_{1985}	0.05(4.37)	u_t^{USA}	-0.62(3.98)		
u_t^{Itl}	0.09(4.04)	TD_{1986}	0.04(3.69)	u_t^{Jap}	-0.66(4.46)		
u_t^{Net}	-0.01(0.62)	TD_{1987}	0.04(3.33)	u_t^{Aud}	-0.71(4.48)		
u_t^{Nor}	-0.01(0.72)	TD_{1988}	0.06(4.82)	u_t^{NZ}	-0.45(3.29)		
u_t^{Spa}	-0.16(17.1)	TD_{1989}	0.07(6.20)	u_t^{Aut}	-0.62(4.33)		
u_t^{Swe}	0.02(1.59)	TD_{1990}	0.06(5.26)	u_t^{Bel}	-0.70(4.68)		
u_t^{UK}	0.02(4.01)	TD_{1991}	0.04(4.43)	u_t^{Den}	-0.93(5.69)		
TD_{1961}	0.03(2.65)	TT_A	-0.02(27.1)	u_t^{Fin}	-0.81(5.03)		
TD_{1962}	0.05(4.48)	TT_B	-0.01(19.8)	u_t^{Fra}	-0.83(6.06)		
TD_{1963}	0.06(5.23)	TT_C	-0.002(6.4)	u_t^{Ger}	-0.68(4.99)		
TD_{1964}	0.06(5.02)			u_t^{Ire}	-0.78(2.95)		
TD_{1965}	0.06(4.78)	$\bar{R}^2$	1.00	u_t^{Itl}	-0.92(3.79)		
TD_{1966}	0.06(5.21)	N	864	u_t^{Net}	-0.85(6.13)		
TD_{1967}	0.06(4.89)	$c^2(36)$	98.3	u_t^{Nor}	-0.84(4.62)		

Notes: Absolute t-statistics are given in parentheses. R^2(mom) = Buse's R-squared. N = number of observations. DW(M) = modified Durbin-Watson test for first order serial correlation in fixed effect panel data models (see Bhargava *et al*, 1982). $Chow(i,j)$ = F-test for coefficient constancy with breaking point in 1975/1976, and is distributed as $F(i,j)$ under the null hypothesis of structural stability. $F(i,j)$ = F-test for cross-country coefficient constancy, and is distributed as $F(i,j)$ under the null hypothesis of coefficient constancy. $c^2(36)$ is Fisher's p_l test for cointegration, and is distributed as chi-squared with 36 degrees of freedom under the null hypothesis of no cointegration (see Maddala and Wu, 1999). TT_A = time trend for Ireland, TT_B = time trend for the USA, Japan, New Zealand, Belgium, Finland and France, TT_C = time trend for Canada, Australia, Austria, Denmark, Germany and Sweden. The following instruments are used for Δp_t^{va}: Δp_{t-1}^{va}, Δp_{t-1}^{x}, Dir_t, Δir_{t-1}, $\Delta m1_{t-1}$, Δp_t^c, and Δp_{t-1}^c, where Δir is the nominal interest rate on a long-term government bond, $m1$ is the log of $M1$, and Δp^c is commodity prices. Estimation period: 1952-1999 (Equation (16)) and 1954-1994 (Equation (17)). The data are collected from various national and international data sources. A detailed list of data sources is available from the author. The country sample consist of the countries as follows: Canada, USA, Japan, Australia, New Zealand, Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, the Netherlands, Norway, Spain, Sweden, and the UK. The estimated coefficients of unemployment are multiplied by 100.