SWEDISH ECONOMIC POLICY REVIEW 2 (1995) 15-58

# Unemployment
# – Is Sweden Still Different?

Anders Forslund*

## Summary

■ Will Sweden, unlike the rest of Western Europe, be able to avoid persistently high unemployment? The upshot of the analysis here is that some optimism seems warranted. First, I have surveyed evidence of similarities and differences between Sweden today and Western Europe as a whole, including Sweden, in the 1980s. Such comparisons do not lead to any clear-cut conclusions. Second, using a highly aggregated macroeconomic framework, I also present new evidence on the development over time of the Swedish equilibrium unemployment rate. The main result of the analysis is that equilibrium unemployment has risen in the first years of the 1990s. The rise is significant (between 1 and 4 percentage points between 1990 and 1993), but not of the same order of magnitude as the "normal" European rise in equilibrium unemployment. It is also below the increase (6.6 percentage points between 1990 and 1993) in actual Swedish unemployment. ■

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# Unemployment – Is Sweden Still Different?

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The Swedish unemployment record was a matter of envy and admiration from an international audience for a long time. In the early 1980s, when most European countries experienced unemployment levels rising to the neighbourhood of 10 per cent, the Swedish unemployment rate peaked at less than 3 per cent. Towards the end of the 1980s, the Swedish unemployment rate was below 2 per cent, whereas the situation in the rest of Europe had not changed much for the better. The real, as yet unsolved, puzzle thus seems to be how Sweden could maintain these low rates, not that the Swedish unemployment rate has recently risen to more "normal" European levels with open unemployment around 8 per cent and "total unemployment" (including participation in labour market programmes) at 12–13 percent.

Although the puzzle of historically low Swedish unemployment rates has not been solved, this does not mean that there is a lack of explanations. In my view, the hypotheses fall into two broad categories. According to the first category, expansionary stabilisation policies, especially devaluations, are stressed (Calmfors, 1993a). The main message, perhaps slightly caricatured, is that the best method of avoiding high and persistent unemployment is to prevent it from getting high in the first place. According to the second category of explanations, Swedish institutions, especially active labour market policies, are what has made Sweden different (Layard *et al.*, 1991). Using observations only from the 1980s, it is hard to discriminate between the two hypotheses since Sweden undoubt-

* *The author gratefully acknowledges constructive comments from Susanne Ackum Agell, Lars Calmfors, Bertil Holmlund and Ragnar Nymoen as well as careful linguistic comments by Julie Sundqvist. The usual caveat applies.*

edly relied on both active labour market policies and expansionary demand policies.

The uncertainty surrounding the explanations of the low Swedish unemployment in the 1980s means that it is very difficult to judge the severity of the currently high unemployment. The crucial question is whether equilibrium unemployment has risen *pari passu* with actual unemployment. If differences in institutions, especially with respect to active labour market policies, were important in the past, perhaps we need not fear that the equilibrium rate has risen so much. But if the main difference from Western Europe in the 1980s was demand developments, there is less reason for optimism.

The main aim of this article is to shed light on the question of what has happened to the equilibrium rate of unemployment in Sweden in recent years. This question is important for judging not only possible future unemployment developments, but also the severity of budget deficit problems (to what extent is the deficit structural?) and the risks for wage inflation in the current upturn.

*First,* I survey some qualitative evidence in the form of comparisons between the present Swedish experience and overall Western European experiences including earlier Swedish developments. Such evidence can at most serve as a preliminary guide to the analysis of whether Swedish equilibrium unemployment has gone up. A quantitative assessment cannot be produced unless statistical methods are used to derive some estimate of the equilibrium unemployment rate. This is the *second* (and main) undertaking of this paper.

There are several possible strategies for estimating the equilibrium unemployment rate. The simplest approach is to estimate price (or wage) equations, where the change in the (wage) inflation rate is a function of the unemployment rate. The equilibrium unemployment rate can subsequently be computed as the level at which the change in inflation rate equals zero (the NAIRU). Another approach is to examine the time series properties of the unemployment rate in order to analyse whether there does exist a stable long-run equilibrium unemployment rate and, if so, how long it takes to return to this rate after various shocks (the degree of unemployment persistence).

The main focus in this paper is to estimate a small structural macroeconomic model that yields a solution for the unemployment rate. Suitable equilibrium conditions can be imposed on the model to give a solution not only for actual unemployment, but also for equilibrium unemployment. The advantage of this approach is that it supplies more than unspecified shocks as explanations for changes (if any) in the equilibrium unemployment rate: we get explanations in terms of estimated parameters and the development of the exogenous variables in the model. Thus, I hope not only to provide evidence on the question of whether Sweden is still different, but also to contribute to our understanding of which factors influence the Swedish equilibrium unemployment rate. This analysis builds on the work by Richard Layard, Steven Nickell and Richard Jackman (see Layard *et al.*, 1991).

The paper is structured as follows. Section 1 gives a brief background by surveying some evidence on Swedish unemployment experiences. Section 2 contains the theoretical framework for the empirical analysis. Section 3 deals briefly with data and estimation issues. Section 4 concerns the results, and Section 5 concludes.

## 1. Unemployment: the Swedish experience

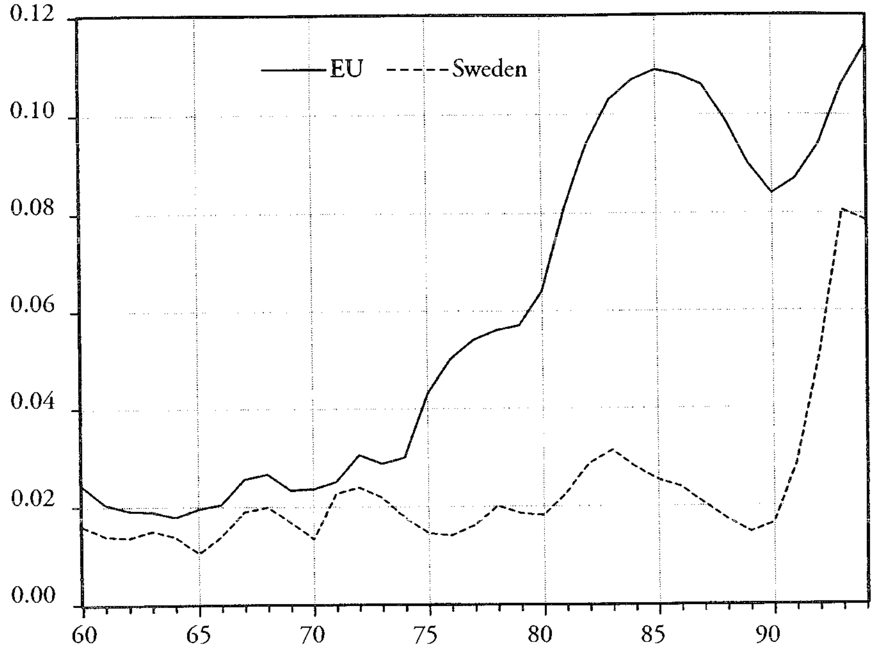
The discussion of the Swedish unemployment record is based on three main components: (i) some aspects of the Swedish unemployment experience and institutions are compared to their Western European counterparts; (ii) the present rise in Swedish unemployment is compared with the (much smaller) increase during the recession of the early 1980s to identify similarities and differences; and (iii) some simple time series evidence on the Swedish unemployment rate is presented.

### 1.1. Unemployment in Sweden and in the EU

A convenient starting point for a comparative discussion of the Swedish labour market is given in Figure 1, where annual Swedish and EU unemployment rates between 1960 and 1994 are plotted.

The Swedish unemployment rate is consistently lower than the average EU rate, with a small gap between them in the 1960s, widening in two steps in the wake of the two oil price shocks in the mid and late 1970s, accompanied by rapid increases in unemployment in Western Europe but not in Sweden. The big difference between Sweden and Western Europe is thus that Sweden somehow managed to avoid the two Western European unemployment hikes in the second half of the 1970s and the first half of the 1980s.

**Figure 1. Open unemployment rates in Sweden and the EU 1960–1994**




*Note*: The EU countries included are Germany, France, Italy, United Kingdom, Belgium, Ireland, Netherlands, Portugal and Spain.

*Sources*: For Swedish unemployment, see Appendix A. EU unemployment 1974–1993: OECD Economic Outlook; 1994: OECD Main Economic Indicators. For the period 1960–1973, the series was obtained by linking unemployment figures from the CEP-OECD data base to the figures from the OECD Economic Outlook.

The diagram also illustrates clearly that once EU unemployment has risen, it has no tendency to return to previous levels. It is also impossible to reject the null hypothesis of a unit root in the series using augmented Dickey–Fuller tests (regardless of whether the test is carried out conditional on linear or quadratic deterministic trends in the series).[1] Although unit root tests on short data series are by no means conclusive, the test results at least indicate a high degree of persistence in the Western Europe-

[1] A unit root test is basically a test for whether a time series has a constant mean (or a mean that changes deterministically) or whether its mean changes randomly. In the latter case the series displays a unit root and has no stable equilibrium. A series with a unit root also has infinitely long "memory" in the sense that any disturbance to it will have permanent effects.

an unemployment rate. It thus seems safe to conclude that, according to the evidence, once unemployment has been allowed to reach high levels, the Western European experience is that a return to lower levels is at best a time-consuming process. At worst there is no tendency towards a return to a stable equilibrium rate: unemployment exhibits hysteresis in the terminology of Blanchard and Summers (1986).

Are there any differences between Sweden and Western Europe which suggest that Sweden can more easily avoid persistently high unemployment? Calmfors (1994) points to a number of such positive factors: (i) absence of recent negative supply-side shocks; (ii) a very substantial real exchange rate depreciation with no counterpart in continental Europe in the 1980s; (iii) short periods of passive unemployment benefits and an emphasis on active labour market policies (ALMPs); and (iv) as yet only a limited increase in long-term unemployment.

However, Calmfors also notes a number of respects in which present Swedish conditions are similar to or even worse than the European situation in the early 1980s: (i) a high replacement ratio for the unemployed; (ii) a guarantee of placement in labour market programmes, effectively creating an indefinite period of income support for the unemployed; (iii) employment protection legislation, creating hiring and firing costs on average European levels; (iv) a compressed wage distribution in Sweden which may exacerbate unemployment problems to the extent that there is an ongoing shift in labour demand from unskilled to skilled labour; (v) high real interest rates; and (vi) demand shocks in the early 1990s without correspondence in other OECD countries (with the exception of Finland) and fiscal deficits which necessitate a long period of fiscal restraint.

Additional evidence in terms of regional employment dynamics is presented in Forslund and Krueger (1994). Their main finding is that with respect to both employment and unemployment, the patterns for Sweden resemble those of other European countries, whereas employment dynamics in the US is different from those in both Sweden and Western Europe. This suggests that ALMPs have not made a significant difference for labour market adjustments at the regional level in Sweden.

If we make a distinction between shocks and propagation mechanisms, it is tempting to draw the conclusion that there are obvious similarities between the current Swedish situation and that facing the rest of Europe in the early 1980s with respect to shocks. The initial employment shocks are certainly different, but there is nothing to indicate that the

Swedish shocks should be regarded as less severe than their earlier Western European counterparts.

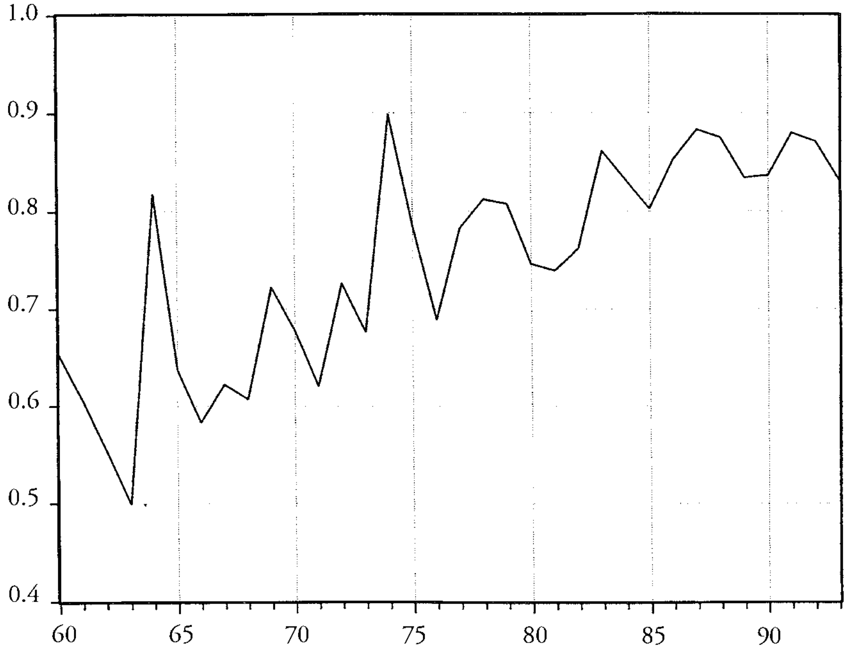
The situation is more complex regarding propagation mechanisms. On one hand, ALMPs and limited benefit duration, the gain in competitiveness and the limited rise in long-term unemployment could make a positive difference. On the other hand, similarities with respect to replacement ratios, high real interest rates, employment protection legislation and *de facto* indefinite income support for the jobless (through the combination of labour market programmes and unemployment insurance) is more worrisome. Moreover, the fiscal deficit effectively puts a brake on any attempts to "shock" unemployment back through fiscal expansion.

## 1.2. Swedish unemployment in the 1990s and earlier experiences

The fact that it is hard to tell whether Sweden is "different enough" from the rest of Europe to prevent us from using the OECD experiences in the 1980s as a predictor for Swedish labour market performance in the 1990s forces us to look for other evidence. One possibility is to make comparisons with earlier Swedish experiences: if Sweden of today is "similar enough" to Sweden in the early 1980s, when propagation mechanisms seem to have dampened the original shocks, there might be some hope for the future. I briefly review two types of evidence by (i) identifying a number of institutional changes since the early 1980s with potential bearing on the functioning of Swedish labour markets, and (ii) comparing recent labour market performance with the developments when unemployment rose in the early 1980s.

On the institutional side one important change is the introduction in 1983 of a *guarantee of placement in relief work* (later extended to other kinds of labour market programmes) in connection with the expiration of unemployment benefits. First, this can be expected to have weakened the incentives for the unemployed to search for jobs actively. Although the results in Carling *et al.* (1994) indicate that the exit rates to employment for recipients of unemployment insurance (UI) increase around the time of benefit exhaustion, the effect seems to be weaker than in corresponding studies for the US. Second, to the extent that the guarantee means that the unemployed are (self-)selected into programmes as a means of qualifying for a renewed UI period, the programmes are likely to be less

**Figure 2. Benefit replacement ratio for income earners in the private sector in Sweden 1960–1993**




*Sources and definition*: See Appendix A.

efficient in bringing the unemployed back to regular employment. There are indications that this may have become a problem. Regnér (1993) found negative effects on the incomes of participants in labour-market-training programmes in the early 1990s, whereas the opposite result was obtained by Axelsson and Löfgren (1991) for the early 1980s.[2]

The *benefit replacement ratio* (the ratio of the unemployment benefit to the wage) is another aspect of UI often found to matter for unemployment performance (Layard *et al.*, 1991). In Figure 2, the maximum replacement ratio for an average income earner in the private sector is plotted. Comparing the early 1980s to the early 1990s, we see that the replacement ratio is marginally higher in the latter period. As of July 1, 1993, the maximum replacement ratio was cut to 80 per cent, so the difference is likely to be rather small today. Furthermore, the estimates in

[2] Participation in retraining now qualifies for a new period with unemployment benefits. This was not the case in the period studied by Axelsson and Löfgren.

Carling *et al.* (1994) indicate that the benefit level has rather small effects on exits from unemployment to employment. This squares well with the difficulties of establishing that unemployment benefits affect wages in studies of aggregate wage formation. The combination of small changes in the replacement ratio and limited expected effects of a given change leads me to the conclusion that we should not expect changes in benefit levels to make a large difference for unemployment performance between the 1980s and the 1990s.

The *degree of centralisation* in the wage bargaining system has received much attention in discussions about labour market performance (Calmfors and Driffill, 1988; Alogoskoufis and Manning, 1988; Layard *et al.*, 1991; Soskice, 1991; Calmfors, 1993b). The empirical studies seem to indicate that higher degrees of centralisation and co-ordination in bargaining are conducive to real-wage moderation and low unemployment (Layard *et al.*, 1991, Soskice, 1991, Zetterberg, 1993). The change from centralised bargaining at the national level to bargaining at the sectoral level that has taken place in Sweden over the last decade might therefore be expected to have raised the equilibrium level of unemployment.[3]

The most obvious difference in labour market performance between the early 1980s and the early 1990s concerns the *level of unemployment* itself. Between 1980 and 1983, open unemployment rose by 1.4 percentage points, whereas the rate went up by 6.6 percentage points between 1990 and 1993. To the extent that unemployment is path dependent (if, for instance, long-term unemployment is positively correlated with the level of unemployment and the long-term unemployed gradually lose productive capacity), the sheer size of the rise in unemployment may signal future problems.

There are, however, also differences in the *structure of unemployment* with potential bearing on future developments in the labour market. First, unemployment in the early 1990s is more evenly distributed in a number of dimensions such as region, education, age and sex[4]. Although not conclusive, this evidence at least indicates that unemployment is not

[3] This view is not undisputed. Calmfors (1993a) notes that centralised bargaining in Sweden has actually involved bargaining at three different levels, and insofar as nominal rigidities are present at different levels of bargaining, such a system is more likely to work well in an inflationary environment than at present low inflation rates.

[4] The most striking feature of the sex distribution of unemployment is that the male unemployment rate was actually higher than the female rate at least until 1993, reflecting a sharp decline in industry and construction.

exclusively concentrated to less educated persons in the "wrong" regions; the unemployed should be "employable"[5]. Second, there is evidence that the rise in unemployment has been accompanied by an increase in long-term unemployment (with possible detrimental effects as hinted at above): although changes in methods of measurement make comparisons hazardous, the *duration of ongoing spells* of unemployment has gone up[6] (Edin and Holmlund, 1994). However, both the absolute size of the rise (7.1 weeks between 1991 and 1993) and the level in 1993 (24.4 weeks) are low from a European perspective. All in all, although the measured duration of unemployment spells has gone up (and this rise may be understating the true change due to an increased "circular flow" between open unemployment and labour market programmes[7]), there are no dramatic changes in the structure of unemployment as compared to the early 1990s.

## 1.3. Time series evidence on Swedish labour market performance

The *Beveridge curve*, which relates unemployment to vacancies, is often used to characterise labour markets: more vacancies give rise to a lower unemployment rate because the unemployed find jobs more easily.[8] Thus, shifts in observed Beveridge curves can be used as indicators of changes in matching efficiency in labour markets. For most European countries, the typical finding is an outward shift of the Beveridge curve during the 1980s (see e.g. OECD, 1992). For Sweden, no such shift has been observed, but this might be related to the fact that the Swedish unemployment rate never increased significantly until the last few years. Does the picture change if we include the period through 1994? Figure 3 depicts the Swedish Beveridge curve for 1960–1994.

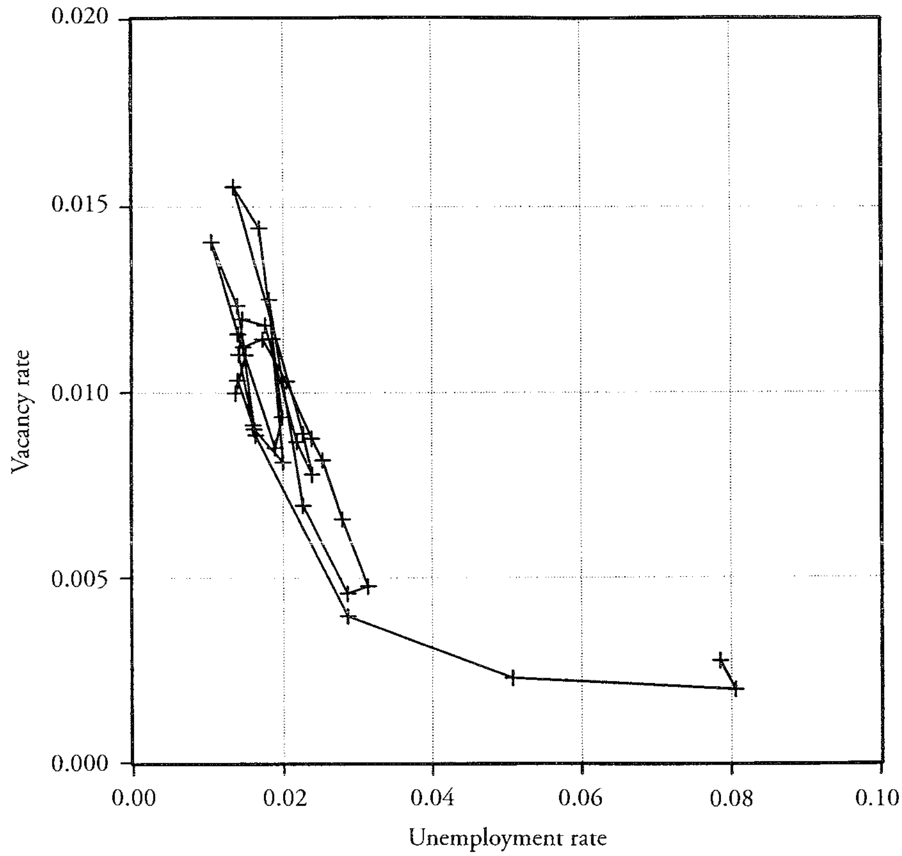
[5] I have used wage equations estimated on data from the Level of Living Survey of 1981 (see Ahlroth *et al.*, 1994) and data on characteristics of the unemployed from labour force surveys to impute wages for the unemployed. These computations indicate that the market value of the skills of the unemployed relative to those of the employed was higher in the early 1990s than in the early 1980s.

[6] This holds despite some changes in the rules for eligibility for labour market programmes.

[7] For evidence on this, see Ackum Agell *et al.* (1995).

[8] The status of the Beveridge curve in the description of labour market equilibrium is discussed in detail in Pissarides (1990).

**Figure 3. The Swedish Beveridge curve 1960–1994**




*Note*: The vacancy rate is the annual average of the end-of-month stocks of vacancies as a fraction of the labour force.

*Source*: National Labour Market Board. For source and definition of the unemployment rate data, see Appendix A.

Most data points are located very closely together in the neighbourhood of 2 per cent unemployment, and the observations for 1992–1994 in the southeast part of the plot are outliers. From the standpoint of the stability of the relationship, however, it is noteworthy that the recent high unemployment rates are matched by very low vacancy rates, meaning that the picture alone does not warrant any firm conclusion about possible shifts of the curve. This is also borne out by more formal analysis: regressing unemployment on lagged unemployment, vacancies and a quadratic trend does not produce any significant parameters on the trend terms. Thus, including the last few years does not provide evidence of a shift in the Swedish Beveridge curve. One possible explanation for this finding is that the otherwise unemployed have to a large extent been placed in labour market programmes, so that the Beveridge curve is actually not a good indicator of matching efficiency as pointed out in Calmfors (1993a). To test this, I regressed the sum of openly unemployed and programme participants ("total unemployment") on lagged total unemployment, vacancies and a quadratic time trend. This exercise did not produce any evidence of a significant shift either.[9] My reading of the evidence on the Swedish Beveridge curve is hence that it seems remarkably stable, showing no significant signs of a deterioration in matching efficiency. This interpretation, however, is partly contradicted by related evidence on matching behaviour in Edin and Holmlund (1991). They estimate the flow of hirings as a function of unemployment and vacancies, and find tendencies towards a lower rate of hirings in a given labour market situation.

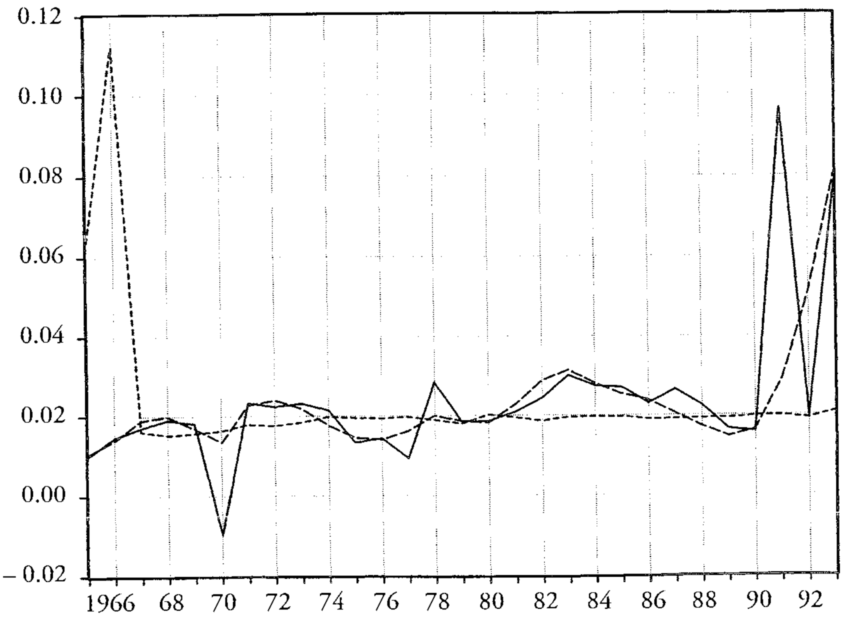
Another set of related evidence pertains to *labour mobility* in different dimensions. A low rate of labour mobility could be taken as a sign of a petrified labour market, in which adjustments to shocks are sluggish. The evidence presented in Edin and Holmlund (1994) indicates that external job mobility (change of employer) shows a trendwise decline, whereas the opposite is true for internal job mobility (change of position with a given employer). Both series show rapid cyclical declines in the early 1990s. Geographical mobility shows no trendwise change over the period since 1900 and is, if anything, slightly higher in the early 1990s than in the early 1980s.

Another type of evidence derives from recursive estimates of equilibrium unemployment rates based on Phillips-curve type models. These come in several forms, two of which are displayed in Figure 4 along with the actual unemployment rate. The NAIRU estimates were derived by regressing the change in the inflation rate (measured as the second difference in the log of the consumer price index[10]) on unemployment recur-

[9] This corroborates the findings in Forslund and Krueger (1994) on regional data. If the regression is run on the natural logarithms of the total unemployment and vacancy rates, the linear trend term is positive while the second-order term is negative. The absolute size of the shift is negligible, and the estimated parameters actually imply an inward shift during the last few years.

[10] Inflation was measured in terms of the CPI because use of the more natural producer price deflator (see Calmfors and Forslund, 1993) gives rise to some unreasonably big swings in the NAIRU series in the second half of the 1970s.

**Figure 4. NAIRU, NAWRU and actual unemployment rate 1965–1993**




sively and solving for the rate of unemployment at which the estimated equations predict constant inflation.[11] The NAWRU (non-accelerating wage rate of unemployment) was derived on the assumption that wages accelerate (decelerate) whenever unemployment is below (above) this rate of unemployment[12],

$$\Delta^2 w = \alpha(\text{NAWRU} - U), \tag{1}$$

where $w$ is the natural logarithm of the wage rate, $\alpha>0$ and $U$ is the rate of unemployment. Assuming that the NAWRU is constant between two

[11] The equations run were $\Delta^2 p_c = a + bu$, where $p_c$ is the CPI and $u$ is log open unemployment, giving NAIRU as $\exp(-a/b)$. Recursive regression means that the regression is run on the shortest possible sample and then observations are added, one at the time. This gives the time series for $a$ and $b$ used in the computations.
[12] Elmeskov (1994) gives a more detailed account of the estimation procedures as well as estimates for a number of countries.

consecutive years, $\alpha$ can be found by taking the first difference of equation (1) as

$$\alpha = -\Delta^3 w / \Delta U. \qquad (2)$$

This expression for $\alpha$ can then be substituted back into equation (1) to give a (possibly) time-varying estimate of the NAWRU as[13]

$$\text{NAWRU} = U - (\Delta U / \Delta^3 w)\, \Delta^2 w. \qquad (3)$$

Apart from occasional "blips", both the NAIRU and (especially) the NAWRU are close to the actual unemployment rate until 1990. Thereafter, the NAIRU stays basically constant, while the NAWRU oscillates wildly and lands at the same level as the open unemployment rate in 1993.[14] The optimist thus has an opportunity to believe in the NAIRU estimates, while the pessimist can verify his beliefs by looking at the NAWRU estimates. The problem for both is that neither measure has any solid theoretical underpinning, so there is no easy way to make a rational choice between them.

We can look at the time-series properties of the unemployment rate itself. Here the crucial issue is whether or not the series displays a unit- or near-unit root (cf. the discussion in footnote 1). If we find a unit root, then there is no such thing as a stable equilibrium rate: the mean value of unemployment changes randomly. If the root is close to unity[15], then there is a stable equilibrium (a stable mean value), but for all practical purposes, the adjustment to that equilibrium is so slow that we cannot distinguish this case from the no-equilibrium case. Such unit root testing can be performed either in a univariate or a multivariate setting.

Looking first at univariate tests, I have performed such on several different sub-samples of the (logged) available maximum series, which ex-

[13] I find it somewhat peculiar to assume a constant NAWRU in order to derive a time-dependent estimate of it. The procedure is, however, applied by the OECD, and I give the results mainly to illustrate the consequence of what seems to be a method in fairly wide use.

[14] The results regarding the NAIRU are consistent with the findings in Holmlund (1993).

[15] A root close to unity means that the process for the unemployment rate can be written $u_t = \alpha + \beta u_{t-1} + \varepsilon_t$, where $\varepsilon_t$ is a stationary disturbance term and $\beta$ is close to (but smaller than) unity. As time goes on, such a process tends to a stable equilibrium value, but with $\beta$ close to unity, the approach is very slow.

tends from 1911 to 1993.[16] Using the full sample, an augmented Dickey–Fuller test is not even close to rejecting a unit root[17]. Dropping observations from World War I and from the 1990s makes no difference. Taking different postwar sub-samples, the general picture is that the later the sample starts, the higher the probability of rejecting a unit root. Dropping the observations from the 1990s also tends to make a rejection of a unit root more likely. My reading of these results is that there is reason to believe at least in a great deal of persistence in the Swedish unemployment rate, and that this conclusion holds even if we restrict our interest to the postwar history (which could be motivated by a belief, warranted or not, that World War II marks a structural break in the workings of the Swedish labour market).[18]

Turning to *results of multivariate testing*, Jacobsson *et al.* (1994) estimated common trends models for Denmark, Norway and Sweden. This can be seen as a multivariate generalisation of univariate unit root testing. Under some set of identifying restrictions and tests for the number of common trends among a set of variables, the time paths of the variables of interest are induced by shocks that might or might not be stationary. If shocks can be described as stationary, the variable will have a stable equilibrium value, otherwise it will not. The main problem with this statistically sophisticated approach is that it gives only vague clues as to what the shocks represent.

In this setting, using quarterly data ranging from the mid 1960s to 1990, Jacobsson *et al.* find "...that there is hysteresis, not only in the Danish labour market, which has shown 'European' tendencies, but also in Norway and Sweden, where unemployment rates have been low and stable historically"(p. 25). Interestingly enough, thus, even using a short data series from the period before the rise in Swedish unemployment, it is

[16] The series was obtained by linking the labour force survey statistics for open unemployment (1960–1993) to UI fund statistics on unemployed members.

[17] A Dickey–Fuller (1979) test is performed by rewriting $u_t = \alpha + \beta u_{t-1} + \varepsilon_t$ as $\Delta u_t = \alpha + (1-\beta) u_{t-1} + \varepsilon_t$ and testing $(1-\beta) < 0$ against the null hypothesis $(1-\beta) = 0$. The test statistic for this test has a non-standard distribution tabulated by e.g. Fuller (1976). This test is valid, given essentially that $\varepsilon_t$ is a white noise process. If it is not, one procedure is to add lags of $\Delta u_t$ until $\varepsilon_t$ becomes white noise and then perform the test, which is the augmented Dickey–Fuller test.

[18] These results are somewhat at odds with the results for Sweden reported in Barro (1988), who found that Sweden is among the countries showing the least unemployment persistence. His point estimates of $\beta$ in $u_t = \alpha + \beta u_{t-1}$ are 0.39 for the period 1920–1938 and 0.53 for the period 1948–1986.

possible to trace hysteresis in the Swedish unemployment rate. This analysis supports the view that the favourable Swedish unemployment experience is primarily the result of historically small "shocks" rather than of "propagation mechanisms" that dampen these shocks.[19]

## 2. An open-economy model of wage- and price-setting

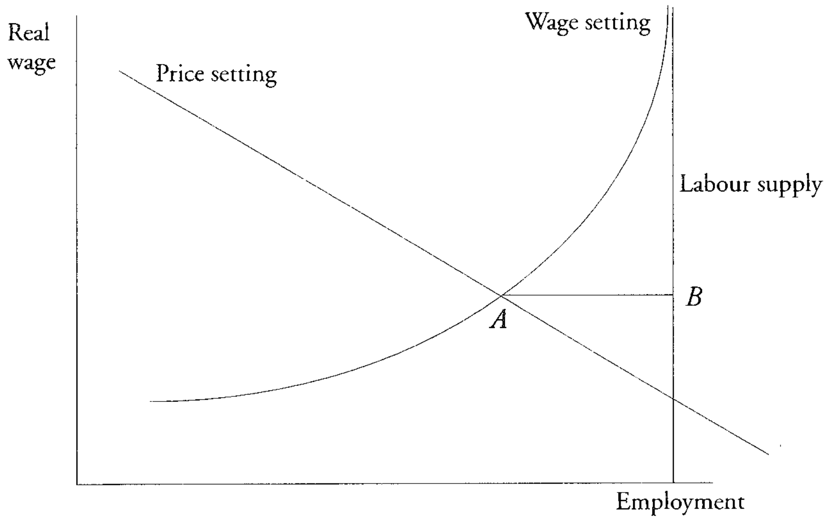
### 2.1. An overview

I now present the barebones of the model used as a basis for my empirical investigation. The framework is the well-known model of Layard and Nickell (1986), where both labour and product markets are assumed to be characterised by imperfect competition.[20] Two of the basic building blocs of the model are the price- and wage-setting schedules depicted in Figure 5. The upward-sloping *wage-setting schedule* shows that wage-setters tend to agree on higher real wages the better the employment prospects are. The downward-sloping *price-setting schedule* (which under perfect competition is an ordinary labour-demand schedule) reflects either pro-cyclical price mark-ups on wages or diminishing returns to labour. The vertical *individual labour-supply schedule* follows from the simplifying assumption that individuals supply labour inelastically. *Equilibrium* in this set-up obtains when the level of employment is such as to make wage-setting and price-setting decisions consistent, i.e., at the intersection of the wage and price-setting schedules. Equilibrium unemployment is then given by the horizontal distance *AB* between this intersection and the labour supply curve.

As a description of equilibrium, consistency between wage- and price-setting decisions is, however, incomplete. (Un)employment will be affected by, for instance, a fiscal expansion if it affects the positions of either of the two schedules. However, for such an expansion to be viable in the longer run, it should not be associated with an uncontrolled accumulation of government or foreign debt. In an intertemporal model, the definition of equilibrium thus involves conditions on government or foreign

[19] This conclusion is also reached by Assarsson and Jansson (1995) using an unobserved components model for unemployment in Sweden and Denmark.
[20] Wyplosz (1994) is an excellent example of how the model can be used to discuss unemployment problems.

**Figure 5. The labour market model**




debt. The counterparts in a static framework, as the one used in this study, are conditions on net government lending or the current account. In the model below, I impose current account balance as an equilibrium requirement.[21]

## 2.2. The empirical model

I begin by discussing the *price-setting schedule.* The general idea is that the planned price mark-up on wage costs is affected by two distinct kinds of influences. *First,* there are upward trends in technological efficiency and the capital-labour ratio causing labour productivity to rise. This tends to increase the demand for labour at given real wages. We thus expect long-run increases in labour productivity to lower the long-run price mark-up on wage costs. *Second,* changes in demand and international competitiveness might influence price-setting under imperfect competition insofar as the marginal productivity of labour depends on the amount of labour used (and thus on the level of output) or the price elasticity of demand perceived by the individual firm is affected.

The realised, as opposed to the planned mark-up, will also reflect expectational errors with respect to inflation: firms set prices given expecta-

[21] The model in Minford (1994) is closed in the same way.

tions about the general price level. As the demand curve facing the individual firm is assumed to be downward sloping, it is costly for the firm to set too high a price (it will lose customers to competitors). Thus, higher than expected prices give "too low" a realised mark-up. I make the simplifying assumption that inflation follows a random walk, and thus proxy inflation surprises (expectational errors) by the change in the inflation rate.[22]

A price-setting schedule reflecting the considerations above can be written

$$p - w = \beta_0 + \beta_1 q + \beta_2(y - \bar{y}) + \beta_3 c + \beta_4 \tau + \beta_5 \Delta^2 p, \qquad (4)$$

where $p$ is the producer price, $w$ the nominal wage rate, $q$ average labour productivity, $y-\bar{y}$ deviation of demand (and output) from a "normal" level, $c \equiv e + p^* - p$ is the real exchange rate ($e$ is the nominal exchange rate and $p^*$ the foreign price level; asterisks denote foreign variables), and $\tau$ is a measure of indirect taxes and subsidies to control for the fact that the price is the producer price, not the factor price deflator. The measure of this tax wedge is constructed as the ratio between value added in the business sector at producer prices and at factor cost. $\Delta$ is the first-difference operator, and $\Delta^2 p$ is thus the change in the rate of change in producer prices (the change in the inflation rate). All variables are expressed as natural logarithms (which I denote by lower-case letters).

As is clear from the discussion above, the expected sign of $\beta_1$ is negative. If the coefficient equals exactly minus unity, this implies constant factor shares of value added. This restriction is tested below.

The expected sign of $\beta_2$ is more uncertain. On one hand, in the presence of fixed or quasi-fixed factors of production, decreasing returns to labour in the short run would tend to make the mark-up an increasing function of demand. On the other hand, the empirical evidence, surveyed by, for example, Layard *et al.* (1991), seems to indicate that the mark-up is rather insensitive to changes in demand. This can be so for a number of reasons. Returns to labour might not be drastically decreasing.

[22] Inflation follows a random walk if $\Delta p_t = \Delta p_{t-1} + \varepsilon_t$, where $\varepsilon_t$ is white noise. Then $\Delta^2 p_t \equiv \Delta p_t - \Delta p_{t-1} = \varepsilon_t$ is white noise. The random walk hypothesis is not rejected by a unit root test. The resulting estimates will, however, be open to the Lucas critique in the sense that the description (and estimates based on it) will only be valid so long as inflation actually follows a random walk.

The price elasticity of demand might vary pro-cyclically. Price changes can be costly to implement and may therefore be avoided if changes in demand are viewed as temporary. Firms may employ more forward-looking pricing strategies than to choose the prices that maximise short-run profits.

The expected sign of $\beta_3$ is positive: higher foreign prices leave more room for domestic firms to raise their prices by making demand less elastic. The effect of the tax wedge (measured by $\beta_4$) is expected to be positive: the higher (net) indirect taxes are, the bigger the mark-up. Finally, $\beta_5$, the effect of surprise inflation, is expected to be negative. If inflation turns out to be higher than expected, firms will find that they have set their prices too low.

A dynamic version of equation (4) was estimated in the empirical analysis. This was necessary both because the static formulation neglects adjustment lags (which for a number of reasons are likely) and because it is impossible to reject the hypothesis that several variables included are non-stationary (have unit roots).

The conceptual model giving rise to the *wage-setting schedule* used is one of bargaining at the firm or industry level. In such a framework all variables that matter to firms and unions in the bargaining situation should in principle enter the wage equation. These variables can in turn be factored into the determinants of the values of "outside options" for both parties involved and the determinants of the mark-up on these outside options.

The value of the outside option for the trade union is determined by the probabilities of a laid-off worker ending up in different states outside the firm and the incomes in these states. I distinguish between three such states: regular employment elsewhere in the economy, $N$, open unemployment, $U$, and participation in a labour market programme, $R$.[23] The probability of ending up in regular employment is assumed to depend positively on the aggregate employment rate, $(N/L)$, where $L$ is the labour force. In addition, I assume that the probability of re-employment for a laid-off worker is smaller – given aggregate employment – the more unemployment has recently gone up. The reason is that the larger the in-

[23] This way of modelling the labour market is discussed in Calmfors and Forslund (1990, 1991). Note that aggregate employment is the sum of private employment, $N_p$, and public employment, $N_g$. The latter is assumed to be treated as exogenously given by wage setters.

crease in unemployment, the smaller the proportion of long-term unemployed, who are likely to search less and thus to compete less efficiently for the available jobs. Furthermore, the conditional probability that a non-employed worker will end up in a labour market programme is assumed to depend positively on the fraction of the "non-employed" put into labour market programmes, $\Gamma \equiv R/(R+U)$.

The unemployed receive unemployment benefits, $B$, and participants in programmes get compensation $W^r$. I assume that the government decides on replacement ratios (rather than on compensation levels) both in UI and in labour market programmes, so that the replacement rates $r_u \equiv b-w$ (for unemployment benefits) and $r_p \equiv w^r-w$ (for compensation in programmes) are treated as given by wage-setters. I did not have access to any weighted measure of compensation in programmes, but the level of compensation in most programmes has been above the level of unemployment compensation for most of the estimation period.

The union (bargaining) model suggests that both the variables determining the probabilities of ending up in the various states ($N/L$ and $\Gamma$) and the replacement rates ($r_u$ and $r_p$) should affect the wage bargain (Calmfors and Forslund, 1990, 1991; Holmlund and Lindén, 1993; Forslund, 1994; Calmfors and Lang, 1995). An additional explanatory variable is the wedge between the real consumption wage, which affects the utility of an employed worker, and the real product wage, which determines employment. If $W_p$ is the real product wage, the real consumption wage is $W_c \equiv W_p(1-T)P/(1+S)P_c$, where $T$ is the average income tax rate, $S$ is the payroll tax rate, $P$ is the producer price and $P_c$ is the consumer price. The wedge is defined as $\bar{\theta} \equiv (1+S)/(1-T)(P/P_c)$. Assuming consumer prices to equal a weighted average of domestic and foreign (import) prices and adding a mark-up reflecting value added taxes, $VAT$, the consumer price is given by $P_c = (1+VAT)P^{\lambda}(P^*)^{(1-\lambda)}$, where $\lambda$ is the weight of domestic goods in the consumer price index. Hence, the wedge is $\bar{\theta} = (1+S)(1+VAT)(EP^*/P)^{(1-\lambda)}/(1-T)$ and reflects taxes and the real exchange rate. I define the tax part of the wedge as $\theta \equiv (1+S)(1+VAT)/(1-T)$.

The size of the wage mark-up on the value of the outside option depends mainly on three factors: the size of the surplus to be divided in the wage negotiations (which in turn depends on labour productivity), bargaining power and the parties' valuation of changes in the wage rate. One potentially important factor influencing the union's valuation of a wage increase is the degree of income tax progressivity. The higher the progressivity, the smaller the benefit of a given wage increase and, thus, the lower the wage demands and the lower the predicted real wage (Lockwood and Manning, 1993; Holmlund and Kolm, 1994). In the empirical formulations I used the elasticity of post-tax income with respect to pre-tax income, $\pi \equiv (1-M)/(1-T)$, where $M$ is the marginal tax rate, as the measure of income tax progressivity and included this variable in the wage equation to be estimated.

With given relative bargaining power of the two parties, a likely outcome is that the product real wage will be unit elastic with respect to productivity in the long run (if the "cake" to be divided through bargaining grows due to productivity growth and if bargaining power is unchanged, the wage rate will tend to grow at the same rate as productivity, or, in other words, be unit elastic with respect to productivity). Labour productivity has grown for centuries. As this has taken place without secular trends in unemployment, it is a common conjecture that labour markets work so that price- and wage-setting are affected symmetrically by labour productivity. If this is the case, price- and wage-setting schedules will shift equiproportionately, and real wages, not unemployment will be affected. Constraints amounting to equal elasticities (possibly equal to unity) with respect to labour productivity in wage setting and price setting were imposed on the empirical estimates (and tested).

Moreover, just as price setters, wage setters may make mistakes when predicting the rate of inflation, which can make the realised real wage rate differ from the desired rate. This is taken care of by entering the change in the inflation rate as a measure of expectational errors in the wage equation as well.

The above considerations lead to the following wage-setting schedule (using a log-linear form):

$$w-p = \alpha_0 + \alpha_1 q + \alpha_2 c + \alpha_3 (n-l) + \alpha_4 \Delta (u-l) + \alpha_5 \theta + \alpha_6 r_u + \alpha_7 r_p + \alpha_8 \gamma + \alpha_9 \pi + \alpha_{10} \Delta^2 p. \qquad (5)$$

The expected effect of productivity, $q$, is positive, and I also tested the restriction $\alpha_1 = 1$, i.e., the hypothesis that the real wage in the long run grows at the same rate as productivity, resulting in a constant wage share of value added. The sign of the effect of the real exchange rate, $c$, reflects two counteracting forces. On one hand, it is part of the wedge between product and consumption real wages, and this effect can be expected to push wages up: if the consumer price index rises because of higher import prices, this lowers the consumption real wage. To the extent that wage earners can compensate themselves for this, the real product wage tends to increase. On the other hand, a rise in the real exchange rate may increase the mark-up of prices on wages (see the discussion of the price-setting schedule). This is tantamount to an inward shift of the labour-demand schedule and exerts downward pressure on the real product wage. This effect will be more pronounced the more important insiders are for the wage formation process.

An increase in the aggregate employment rate, $(n-l)$, improves the value of the outside option for the employed (regular employment elsewhere becomes more probable), and the expected sign of $\alpha_3$ is hence unambiguously positive. The change in the unemployment rate, $\Delta(u-l)$, is supposed to be associated with harder competition for jobs among the unemployed, and the expected sign of $\alpha_4$ is negative. The tax-wedge, $\theta$, can be expected to raise wage pressures for the reasons given above. Higher replacement ratios, $r_u$ and $r_p$, mean better outside options for union members (higher incomes for those out of work), and consequently higher real wages.[24] A larger probability of placement in a labour market programme, $\gamma$, means better prospects for the jobless, and hence higher wages. A greater elasticity of post-tax income with respect to pre-tax income, $\pi$, makes it more worthwhile to press for higher wages, so the expected sign of $\alpha_9$ is positive. Finally, underpredicted prices cause too low realised real wages and hence $\alpha_{10}$ is expected to be negative.

To close the model, I need two additional relationships. *First,* the product-market demand-pressure variable, $(y-\bar{y})$, appearing in the price-setting relation is related to the labour market situation by an "Okun's-law" type of equation. *Second,* as discussed above, equilibrium in the model involves current account balance. Consequently, an equation for the current account is also specified.

[24] I did not have access to any good summary measure of compensation in labour market programmes (so no such measure could be included in the estimated equation). For some programmes (e.g. relief work) compensation is tied to contractual market wages. In others (e.g. training programmes) compensation is at least as high as unemployment benefits. Thus, on average it seems warranted to assume that compensation in programmes exceeds unemployment benefits. If this is so, then an increase in participation in programmes (a rise in $\gamma$) is likely to be wage raising. In addition, results in Korpi (1994) indicate a positive effect on the psychological well-being among those in programmes compared to the unemployed. This further strengthens the presumption of a wage-raising effect of increased participation in programmes.

The measure of product demand used pertains to the private sector. The *Okun's-law* equation estimated therefore relates deviations from capacity output in the private sector to deviations of private employment ($N_p$) from the labour force less public employment ($L-N_g$):

$$y-\bar{y} = \varphi_0 + \varphi_1(n_p-\log(L-N_g)). \tag{6}$$

The general idea motivating the *current account* equation is that domestic (as well as foreign) demand depends on output (incomes) and real interest rates, and that the distribution of demand between domestically and internationally produced goods is determined by the real exchange rate. In addition to this, the relative price of oil, $p_{\text{oil}}$, is introduced to capture changes in prices of imported raw materials. To avoid modelling the (obviously) endogenous domestic real interest rate, I made use of the fact that the Swedish current account balance is the negative of foreign net exports. These depend on an international real interest rate. After some experimentation, the real interest rate chosen was the short-run German real interest rate, $i^*$.

$$ca = \sigma_0 + \sigma_1 c + \sigma_2(y-\bar{y}) + \sigma_3(y-\bar{y})^* + \sigma_4 i^* + \sigma_5 p_{\text{oil}}, \tag{7}$$

where *ca* is the current account relative to trend nominal GDP.

A rise in the real exchange rate, $c \equiv e+p^*-p$, means a better competitive position for domestic firms, so $\sigma_1$ is expected to be positive. At a given real exchange rate, higher domestic income means more domestic demand; higher foreign income means more foreign demand, so the expected signs of $\sigma_2$ and $\sigma_3$ are negative and positive, respectively. The international real interest rate is introduced to reflect that the current account is the balance between foreign savings and investment, which may both be influenced by a real interest rate. A higher foreign rate implies more saving and less investment abroad, and thus it should be expected to have a negative effect on the Swedish current account balance. The relative price of oil is in the equation as a proxy for the price of imported raw materials. The expected sign is negative.

The complete model as written down in equations (4)–(7) has a short-run solution for the real wage rate, (un)employment, the change in the inflation rate (the expectational error) and the current account, conditional on all exogenous variables, the real exchange rate and the level of demand.[25] The latter two variables are assumed to be determined by an open-economy IS–LM type sub-system (which is not modelled explicitly) in the short run. This system consists of an ordinary LM-curve, an IS-curve augmented with the real exchange rate and an equation for uncovered interest rate parity. Together, these equations determine the real exchange rate and output (demand) as functions of fiscal and monetary policies, the foreign real interest rate, expected inflation and the long-run equilibrium real exchange rate.[26]

In long-run equilibrium, we require a constant inflation rate and current account balance. The system then determines equilibrium values of the real wage rate, unemployment, aggregate demand and the real exchange rate. It should be noted that according to our model, fiscal and monetary policies thus do not influence the long-run position of the economy. The interpretation is that they have to adjust so as to be consistent with the equilibrium conditions imposed on the model.

In order to derive estimates of the equilibrium unemployment rate, the following strategy is followed below. *First*, dynamic counterparts to the structural equations (4)–(7) are estimated. *Second*, the long-run static equations are derived from the estimated dynamic equations. *Third*, the solution to this system when the change in the inflation rate and the current account are set to zero gives the estimated equilibrium unemployment rate.

## 3. Data and estimation

The data used are annual observations for the period 1960–1993.[27] The level of aggregation is the total economy. The main data source is Swedish National Accounts Statistics. Most of the variables used are standard series, but some are not. A description of the data is given in Appendix A.

Analysts of empirical relations between macro variables always have a delicate choice to make as to the sampling frequency of the data used.

[25] Actually, there is a fifth equation in the model, the definition of the labour force. This equation is used to eliminate private employment from equation (6).
[26] The latter three variables enter the model through the interest rate parity condition. For the details of a sub-system like the one discussed in the text, see Layard *et al.* (1991, Ch. 8). The IS–LM system also naturally suggests instruments to use in the empirical estimations.
[27] Many variables are available for a longer time period, but the wage rate for the private sector and all data pertaining to employment cannot be traced back further than to 1960.

*Ceteris paribus,* both long time periods and many observations are desirable properties of time series. On one hand, quarterly data give many observations, but a shorter time span than annual data. This means that a greater proportion of quarterly observations fall in recent time periods. This might be a desirable property if one believes that underlying relationships have changed over time and if one is interested in projections for the near future. On the other hand, the shortness of the observation period when using quarterly data means that relationships between variables in lower frequency bands ("long-run" relations) are hard to identify. My main reason for having chosen annual observations, though, was a strong suspicion of measurement errors in many quarterly time series.

As is well known by now, there are problems involved in identifying the wage equation in models of the type used in this study.[28] This derives from the fact that in the general wage-bargaining model all factors that shift the price-setting (or labour-demand) schedule ought also to affect the wage-setting schedule. Thus, there is a problem of imposing restrictions on the wage equation (that is, motivating why certain variables should be omitted from this equation). Hence, there is always a risk that parameters in estimated wage equations are actually linear combinations of parameters of both wage-setting and price-setting equations. If the main interest lies in identifying the wage equation, this is a serious problem. The focus of this paper, however, is on unemployment. For this purpose the crucial question is whether the model has a solution for unemployment or not, and this model indeed has one. It can also be shown that the solution is invariant to adding multiples of the price-setting equation to the wage-setting schedule. For the purpose of analysing the determinants of unemployment, as opposed to the question concerning wage setting, identification is thus not crucial. As a check for robustness, however, I also estimated a reduced form for the unemployment rate. These estimates are presented as a complement to the estimates of the structural equations.

As has been noted above, the estimated model is dynamic. The dynamics are admittedly of an *ad hoc* nature. The general strategy was to incorporate prior information from univariate unit root tests into the theoretical model. Basically, this procedure suggests where to expect relationships between variables in levels and first-difference forms, respectively. This is a necessary step if one wants to avoid problems of inference relat-

[28] See e.g. Bean (1994), Layard *et al.* (1991) or Manning (1993).

ed to unbalanced and spurious regression models.[29] The reduced-form estimates derive from a specification search, where the starting point is to include all exogenous variables suggested by the theoretical model. The search amounts to eliminating all insignificant variables as long as tests do not indicate specification problems (serially correlated residuals, parameter instability, etc.).

The structural model was estimated using two-stage least squares methods. Although the principal problems of identification should be kept in mind, formal identification is achieved. The instruments used include some foreign demand and price variables, some domestic policy variables and measures of technical efficiency and the capital-labour ratio.[30] Also in the structural model, insignificant variables were eliminated by similar procedures as in the reduced-form estimations.

## 4. Results

The estimated dynamic model is reproduced in Appendix B. The corresponding long-run equations are given below:

$$p-w = 0.20 + 4.11\Delta^2 p + 1.38c - q - 11.03\tau + 2.13(y-\bar{y}) \qquad (4')$$

$$w-p = -0.40 - 0.05\Delta u - 0.65\Delta^2 p + 1.44(n-l) + 0.13\gamma - 0.50c + q + 0.13\theta \qquad (5')$$

$$y-\bar{y} = 0.04 + 0.59(n_p - \log(L-N_g)) \qquad (6')$$

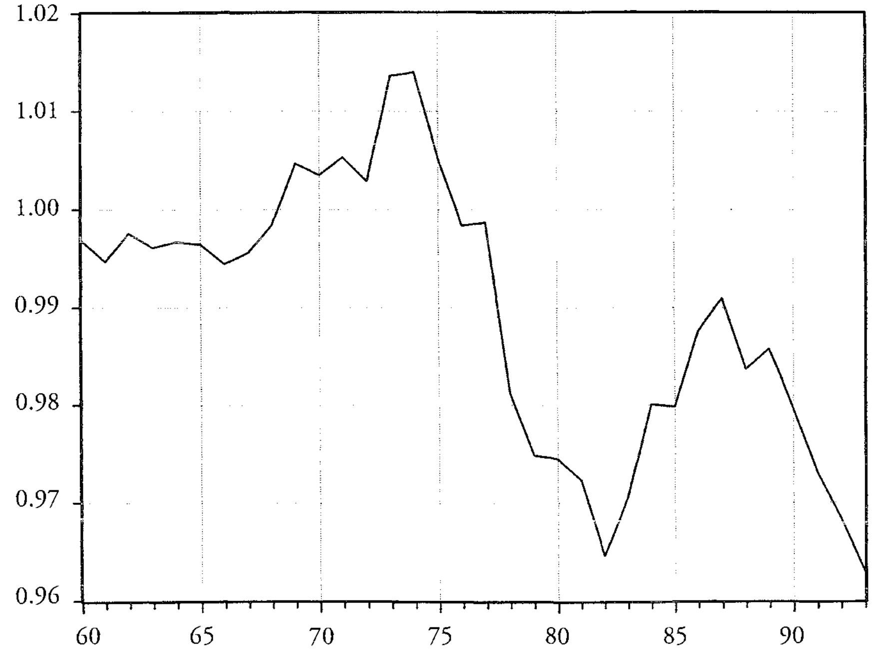
$$ca = 0.07 + 0.21c - 0.01p_{\text{oil}} - 0.002r^* - 0.16(y-\bar{y}) + 0.57(y-\bar{y})^* \qquad (7')$$

Most parameters have the expected signs; the notable exceptions both pertain to the price-setting equation. The mark-up should according to theory depend negatively on inflation surprises, whereas the estimate in-

[29] An equation is said to be unbalanced when the orders of integration on the left- and right-hand sides do not coincide. A stationary variable is integrated of order zero, a non-stationary variable that can be made stationary by taking first differences is integrated of order one and so on. A regression using variables which are integrated of the first order on both the left-hand and right-hand sides is said to be spurious. Test statistics do not have standard distributions in either case, see e.g. Banerjee *et al.* (1993).

[30] A comprehensive list of instruments is included along with the printout of the estimated model in Appendix B.

**Figure 6. Tax wedge between value added at producer prices and at factor cost 1960–1993**




*Source*: See Appendix A.

dicates a rather strong effect in the opposite direction. This casts some doubts on whether the term actually captures what it is supposed to. The sign on the (net) tax wedge between value added at producer prices and at factor cost is also surprising. Taken at face value, the estimate implies that a one percent decrease in this wedge (higher government subsidies to the business sector) causes the long-run mark-up of producer prices on wage costs to increase by 11 per cent. It is not clear what this estimate captures. Some lead can be obtained from Figure 6, which plots the series for the wedge. It should be noted that the wedge falls (subsidies increase) both between 1974 and 1982 and from the end of the 1980s and onwards, whereas there is an increase (lower subsidies) during most of the 1980s. This development does suggest that the variable may actually capture cyclical demand over and above the output gap variable: a higher price-wage mark-up at a given output gap (employment rate) is synonymous with lower output and employment at a given real wage, i.e., a leftward shift of the labour demand schedule. Another possible explanation is that the wedge variable actually acts as a proxy for the degree of competitive pressure. If this interpretation is correct, higher subsidies (a lower wedge) would signal less of competitive pressure and raise the mark-up. Especially the development between 1974 and 1982 (a rapid rise in net subsidies) lends some support to this interpretation. It is well known that industrial policies over this period involved subsidising weakly performing firms. One might also hypothesise that the wedge variable captures the accommodative stance of fiscal policies, i.e., the degree to which policies are expected to counteract the adverse employment effects of supply-side wage shocks.

The long-run coefficients on productivity in price-setting and the wage-setting equations are restricted to minus and plus unity, respectively. These restrictions are not rejected at conventional significance levels. The implications are that unemployment is independent of labour productivity and that the wage share of value added is constant in the long run.

In the wage-setting equation, neither the replacement ratio in the UI system nor the measure of tax progressivity produced significant coefficients. This might reflect an absence of effects, but it is in my opinion equally likely that the reason is related to the general problem of identifying separate effects of many (often collinear) variables in short time series.

The model cannot be explicitly solved for the unemployment rate. I circumvented this problem in the following way. First, equation (6') was used to eliminate the demand pressure variable from equations (4') and (7'). Then, equation (7') was used to eliminate the real exchange rate from equations (4') and (5'). In the next step, equations (4') and (5') were solved for terms in $(n-l)$ and $(n_p-\log(L-N_g))$ as functions of all the exogenous variables, the current account and the change in inflation. In equilibrium, the latter two variables are set to zero. The equilibrium rate of change in open unemployment plus participation in programmes ("total unemployment") as a fraction of the labour force can subsequently be solved for in terms of rates of changes in all the exogenous variables (including public employment) using a linear approximation. Finally, the rate of change in total unemployment is distributed between changes in open unemployment and participation in programmes using the rate of change in $\gamma$, the share of the jobless placed in programmes. As the computations involve linearisations, they are less accurate, the greater the rates of change. Errors of approximation also accumulate over time. Thus, as the focus of interest is on the last few years, I concentrate on the development since 1990. The computed rate of change in equilibrium

**Table 1. Annual rates of change in equilibrium unemployment and its determinants 1991–1993**

| Effect of change in | 1991 | 1992 | 1993 |
|---|---|---|---|
| $\Delta u$ | –0.10 | 0.00 | 0.01 |
| $\gamma$ | 0.00 | –0.01 | 0.05 |
| $p_{oil}$ | –0.05 | –0.01 | 0.01 |
| $i^*$ | 0.01 | 0.00 | –0.03 |
| $\tau$ | 0.35 | 0.14 | 0.11 |
| $n_g$ and $l$ | –0.01 | 0.02 | 0.07 |
| $\theta$ | –0.06 | –0.01 | 0.00 |
| $(y-\bar{y})^*$ | 0.22 | 0.08 | 0.09 |
| Total change | 0.34 | 0.14 | 0.26 |
| Total change less effects from foreign demand | 0.12 | 0.06 | 0.17 |

unemployment since 1990 and its proximate determinants are reproduced in Table 1.

A number of interesting features stand out in the table. *First*, labour market programmes ($\gamma$) and public employment ($n_g$) make modest contributions to the total change in the first two years. However, the net effect in 1993 amounts to a 12 per cent increase in equilibrium unemployment. For labour market programmes, the net effect reflects the fact that the direct effect of lifting people out of open unemployment dominates the indirect wage raising effect. Public employment also raises wage pressure through its effect on the employment rate in equation (5'). This effect pushes the equilibrium rate upwards. In addition, there is a channel working through "Okun's law" (equation (6')): higher public employment means higher demand pressure. Higher demand pressure increases the price mark-up on wage costs according to the price-setting schedule (4'). Given the estimates, this latter effect dominates the former in equilibrium (partly because the current account is rather inelastic with respect to domestic demand according to equation (7')), so the fall in the public employment share in 1992 and 1993 raises equilibrium unemployment.

*Second*, although the tax wedge ($\theta$) between the consumption real wage and the product real wage has a non-zero long-run effect, even the rather rapid fall in the wedge in 1991 did not lower the equilibrium rate by more than 6 per cent (assuming equilibrium unemployment in 1990 to equal 3.5 per cent, this amounts to 0.2 percentage points).

*Third*, the rapid rise in the unemployment rate itself ($\Delta u$) put a brake on the rise in the equilibrium rate in 1991, but has since then not had much effect.

*Fourth*, the influence of foreign variables derives mainly from the changes in foreign cyclical demand $(y-\bar{y})^*$. As it is not unreasonable to define equilibrium given "normal" demand conditions in the rest of the world, the bottom line of Table 1 gives the changes in the equilibrium unemployment rate at constant cyclical foreign demand. As is clear from the table, this affects the results substantially. Both in 1991 and 1992, the drop in foreign cyclical demand accounts for more than half of the net increase in the equilibrium rate. Also in 1993 the increase is reduced significantly if the effect from foreign demand is set to zero. The obvious implication is that much of the increase in Swedish (equilibrium) unemployment is driven by the international recession of the early 1990s.

*Fifth*, the most troublesome feature of the results is the strong estimated impact of the tax wedge ($\tau$) between value added at producer prices and at factor cost. As discussed above, it cannot be ruled out that this effect captures cyclical demand, although here it is regarded influencing equilibrium unemployment.

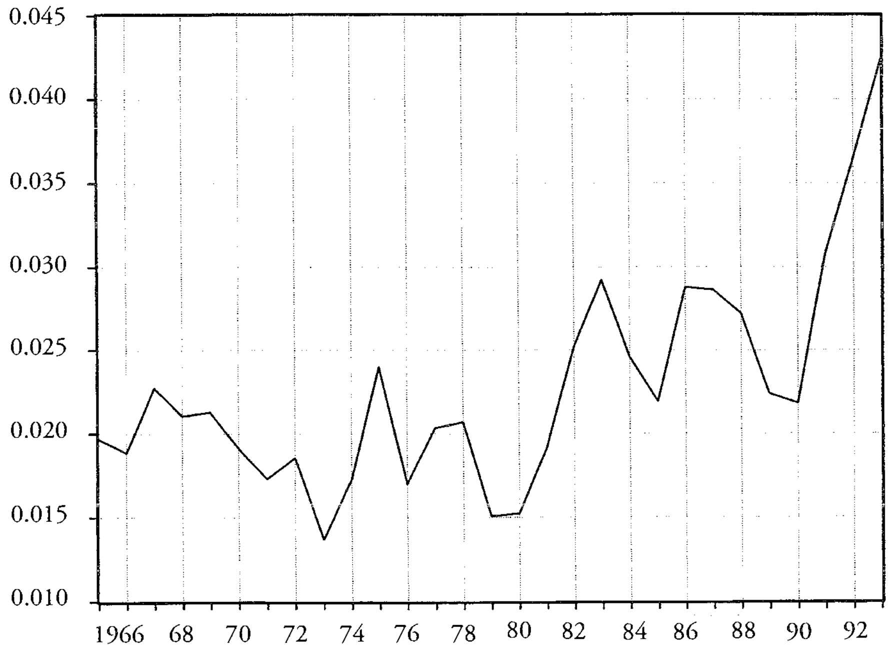
Given the estimated change in the equilibrium unemployment rate reproduced in Table 1 and a "guesstimate" of the rate in 1990, it is easy to compute estimates of subsequent levels of the equilibrium rate of unemployment. Four such estimates are reproduced in Table 2. The estimates differ in the level chosen for 1990 and in whether or not changes in foreign cyclical demand are accounted for. The guesses about the level in

**Table 2. Estimates of the equilibrium unemployment rate 1990–1993**

| | 1990 | 1991 | 1992 | 1993 |
|---|---|---|---|---|
| Level at actual foreign demand | 0.030 | 0.042 | 0.048 | 0.063 |
| Level at unchanged foreign demand | 0.030 | 0.034 | 0.036 | 0.043 |
| Level at actual foreign demand | 0.035 | 0.049 | 0.057 | 0.073 |
| Level at unchanged foreign demand | 0.035 | 0.039 | 0.042 | 0.050 |

*Note*: The levels in 1990 are "guesstimates", the levels in subsequent years are computed using the rates of change from Table 1 at actual and unchanged foreign cyclical demand, respectively.

**Figure 7. Equilibrium unemployment rate from reduced-form unemployment equation 1965–1993**




1990 are, of course, somewhat arbitrary. The two alternatives used are 3 and 3.5 per cent, which are both above the NAIRU and NAWRU estimates in Figure 4 as well as the rate predicted by the reduced form (see Figure 7).

The general impression given by the table is that the results indicate a significant rise in the equilibrium rate of unemployment taking place in the early 1990s if we assume that the tax wedge between value added at producer prices and factor cost reflects "structural parameters" rather than cyclical demand. However, given the guesstimates for 1990, even in this case the equilibrium level does not reach the actual level in 1993.

The equilibrium unemployment rate derived from the reduced-form unemployment equation is displayed In Figure 7. The reduced form was found by regressing unemployment on exogenous variables and (some) instruments from the structural model with one important proviso: in order to make the distinction between unemployment and equilibrium unemployment, the current account and the change in inflation rate were included among the regressors.[31] The latter, however, was never close to significant, and has consequently been dropped. The current account was only significant differenced, implying an absence of a long-run relationship with unemployment. The estimated parsimonious reduced-form equation is:

$$\Delta(u-l) = -1.01 + 6.44\Delta ca + 0.58 r_{u(-2)} - 0.47(u-l)_{(-2)} - 0.13 p_{\text{oil}(-1)} -$$

$$0.81\Delta\gamma - 5.56(y-\bar{y})^{*} \tag{8}$$

Equation (8) deserves a few comments. (i) As accommodation through labour market policies enters the equation differenced, the long-run effect of labour market policies on open unemployment equals zero. In the short run, programmes do, however, lower both open and total unemployment. (ii) The signs are, with the exception of the sign of the oil-price effect, as expected. One possible conjecture is that much of the contractionary effect of an oil-price increase actually works through the foreign demand variable. The negative estimated effect of lagged oil prices might then capture policy reactions triggered by the development of world oil prices. (iii) The rather strong effect of the benefit replacement ratio is notable, as that variable was found to be insignificant in the estimated wage-setting equation. This highlights the problems of identifying separate effects of several collinear variables. (iv) The point estimate of the short-run effect of changes in the current account implies that an increase in the current account amounting to one percentage point of GDP increases unemployment by about 6.5 per cent (which at 5 per cent open unemployment translates into about 0.3 percentage points).

The equilibrium rate given in equation (9) and Figure 7 is the unemployment rate predicted by the reduced form with all lags eliminated and the current account and the change in the fraction of the jobless in programmes ($\Delta\gamma$) set to zero:

$$u-l = -2.17 - 0.28\, p_{\text{oil}} + 1.24 r_u - 11.93\,(y-\bar{y})^{*} \tag{9}$$

[31] The principle was to include instruments suggested by the IS–LM model as determinants of the real exchange rate and aggregate demand. The complete regression printout is reproduced in Appendix B.

The general picture conveyed by the figure for the period until 1990 is very similar to the NAIRU and NAWRU estimates presented in Section 2: the series displays only small fluctuations in the neighbourhood of 1.5–3 per cent. The development from 1990 and onwards, on the other hand, qualitatively resembles the results from the structural model. This resemblance notwithstanding, it is important to note that the direct estimates of the reduced form point to a different set of variables driving the results. On one hand, the tax wedge between producer prices and factor cost which, according to the structural model, was the most important factor behind the recent rise in equilibrium unemployment, is not even close to significance in the reduced form. On the other hand, the replacement ratio in the UI system, which was never found significant in the wage-setting equation, exerts the expected upward pressure on the equilibrium unemployment rate according to the reduced form. It is also noteworthy that the reduced form allocates much of the driving force to foreign demand. Setting foreign cyclical demand equal to zero, the model still predicts an increase in the equilibrium rate between 1990 and 1993, but now only from 1.8 to 3.0 per cent. As this is driven entirely by falling oil prices, one might feel somewhat sceptical about the result.

## 5. Concluding comments

Is Sweden, unlike the rest of Western Europe, likely to be able to avoid persistently high unemployment? The upshot of the evidence presented here, although admittedly shaky, is that some optimism seems warranted. I have surveyed evidence of similarities and differences between Sweden today and the rest of Western Europe, including Sweden, in the 1980s. Such comparisons do not lead to any clear-cut conclusions, so there is definitely a market for more empirical analysis.

Using a highly aggregated macroeconomic framework, this study presents new evidence on the development over time of the Swedish equilibrium unemployment rate. This rate is defined as the unemployment rate that the estimated version of the model predicts conditional on current account balance and the absence of inflationary surprises. One interpretation of the results is that equilibrium unemployment has risen in the first years of the 1990s. If we take the results at face value, the rise could be of the order of magnitude of 1–4 percentage points between 1990 and 1993, which appears far smaller than the "normal" European rise in equilibrium unemployment. It is also below the increase (6.6 percentage points between 1990 and 1993) of actual Swedish unemployment. To the extent that my structural variables – notably the wedge between value added at producer prices and at factor cost – actually reflect demand variables, my estimates of the rise in equilibrium unemployment may, however, be biased upwards. In all, this suggests that there may be good reasons for optimism regarding future unemployment developments.

The results should naturally be treated with caution. I therefore conclude by mentioning some obvious limitations of the analysis.

(i) Even though the estimated equations generally seem to characterise the data fairly well, the model is very small. Although this has the benefit of making the model transparent, it is certainly accomplished at the cost of neglecting potentially important relationships.

(ii) Although both the structural and reduced-form estimates produce reasonably similar estimates of how equilibrium unemployment has developed in recent years, they do not agree on the causes. This points to a general problem of model selection when time series are short and many variables singled out by theory are highly correlated.

(iii) As pointed out on several occasions, there is great uncertainty as to exactly what the wedge between value added at producer prices and at factor cost really captures. More fundamentally, it is crucial for the interpretation of the results whether the variable is a proxy for cyclical or structural influences. If the former is the case, then the increase in equilibrium unemployment is very modest. This raises the following problem: if, as the survey evidence reviewed in Section 1 indicates, only minor structural changes in the Swedish labour market have taken place, then models of the kind used in this study perhaps by necessity will allocate most of the observed changes in unemployment to cyclical factors. Perhaps another type of modelling is needed to catch hysteresis phenomena arising from the interaction between unchanged institutions and a large demand shock such as the one recently experienced.

(iv) There is considerable uncertainty as to whether estimates based mainly on a long period with very low unemployment rates can be used to predict the future development starting in a situation with high unemployment.

# Appendix A. The data

The series on *deviations from full employment output* $(y-\bar{y})$ was derived using a Whittaker-Henderson filter, which basically decomposes the variations in a given time series into trend and cycle (see e.g. Danthine and Girardin, 1989). The *real exchange rate* $(c)$ *and foreign demand* $(y-\bar{y})^*$ variables both involve weighing together variables from a large number of countries. The weights used are in all cases from the IMF MERM model. The real exchange rate is defined in terms of consumer prices, and the foreign demand variable was derived in the same way as the deviations from Swedish full employment output. The *capital stock* and the *index of technical efficiency* (which are used to instrument productivity) are both updates of Bengt Hansson's estimates (Hansson, 1991, and Bergman and Hansson, 1992). The *labour market programmes* include relief work, retraining, youth programmes, recruitment subsidies and job introduction projects. The *labour force* was not taken from labour force surveys; instead it was generated as the sum of private employment, public employment, unemployment and participants in programmes (those not considered employed in the statistics). The *replacement ratio,* $(r_u)$, is defined as the ratio between the maximum before-tax unemployment compensation for a member of an insurance fund and the wage income on a weekly basis (average weekly hours times the average wage rate). The maximum, rather than the average, level of compensation was chosen to reflect the fact that the median union member in all probability faces this level. A similar line of reasoning lay behind the decision not to weigh together cash assistance benefits (CA) with the compensation level of the insurance fund member: the union member is always insured. Progressive taxes mean that there is a slight difference between before-tax and after-tax replacement ratios, but since the difference is slight, it has been neglected. The *wedge* $(\theta)$ between product and consumption real wages was computed using payroll tax rates for blue-collar workers and an estimated average income tax rate for a labour union member. *Net taxes, (txne),* were computed using the change in government debt and the nominal value of government consumption and investment: $Tx-Tr = C_g+I_g-\Delta GD$, where $Tx$ and $Tr$ are taxes and transfers, respectively; thus, $Tx-Tr$ net taxes, $C_g$ and $I_g$ are government consumption and investment, respectively, and GD is government debt.

## List of variables

| | |
|---|---|
| $c \equiv e + p_c^* - p_c$ | real exchange rate, defined as above. Sources: exchange rates, foreign consumer prices and weights: Bank of Sweden. Swedish consumer prices: Statistics Sweden. |
| $\tau$ | tax wedge between value added at producer prices and at factor cost, computed as ratio between value added at producer price and at factor cost. Source: Statistics Sweden, National Accounts Statistics. |
| $a$ | index of technical efficiency ("Solow residual") in the private sector. Sources: Bergman and Hansson (1992), own computations. |
| $k$ | capital stock in the private sector. Sources: Hansson (1991), own computations. |
| $\bar{n}_p$ | labour force less public employment. Sources: labour force: see list of variables; public employment: Statistics Sweden, National Accounts Statistics. |
| $g$ | real public consumption and investment. Source: Statistics Sweden, National Accounts Statistics. |
| $txne$ | net taxes as fraction of potential nominal GDP, defined as above. Sources: Riksgäldskontoret (public debt), Statistics Sweden (nominal public expenditures) |
| $(y-\bar{y})^*$ | foreign cyclical demand. Whittaker-Henderson filtered series of MERM-weighted foreign GDP. Sources: CEP-OECD data base (GDP 1955–1990), OECD Main economic indicators (GDP 1990–1993), Bank of Sweden (weights). |
| $i^*$ | short-run real German interest rate. Sources: Bank of Sweden (consumer prices, short-term interest rate 1963–1992), IFS (short-term interest rate 1955–1962, 1993). |
| $p_{\text{oil}}$ | relative price of oil (relative to $p^*$). Sources: IFS, Per Jansson. |

| | |
|---|---|
| $\theta$ | tax wedge between real product and consumption wages. Sources: NIER (payroll taxes), Statistics Sweden, National Accounts Statistics (VAT), own computations (average income tax rates). |
| $n_g$ | public employment. Source: Statistics Sweden, National Accounts Statistics. |
| $l$ | labour force. Sources: Statistics Sweden, National Accounts Statistics (private and public employment), Statistics Sweden, Labour force surveys (open unemployment), National Labour Market Board (participation in labour market programmes). |
| $r_u$ | benefit replacement ratio, defined as above. Sources: National Labour Market Board – the insurance unit, own computations. |
| $\gamma \equiv \ln(R/(R+U))$ | fraction of the jobless placed in labour market programmes. Sources: National Labour Market Board (participation in programmes), Statistics Sweden, Labour force surveys (open unemployment). |
| $\pi \equiv \ln(1-M)/(1-T)$ | elasticity of post-tax income with respect to pre-tax income. Sources: own computations (T), Locking (1994). |
| $p$ | implicit price deflator for value added in private sector at producer price. Source: Statistics Sweden, National Accounts Statistics. |
| $w$ | hourly wage cost in private sector. Sources: Statistics Sweden, National Accounts Statistics (wage sums, hours worked), NIER (payroll tax rates). |
| $q$ | real value added per hour worked in the private sector. Source: Statistics Sweden, National Accounts Statistics. |
| $(y-\bar{y})$ | domestic cyclical demand (Whittaker-Henderson filtered real private sector value added). Source: Statistics Sweden, National Accounts Statistics, own computations. |

| | |
|---|---|
| $n$ | employment (private plus public); includes both employed and self-employed in the private sector. Source: Statistics Sweden, National Accounts Statistics. |
| $n_p$ | employment in private sector. Source: Statistics Sweden, National Accounts Statistics. |
| $ca$ | current account balance relative to trend nominal GDP. Sources: Bank of Sweden (current account balance), Statistics Sweden, National Accounts Statistics (nominal GDP), own computations. |
| $u$ | number of openly unemployed. Source: Statistics Sweden, Labour force surveys. |
| $r$ | number of participants in labour market programmes, as defined above. Source: National Labour Market Board. |

## Appendix B. The estimated models

The estimated parameters are reproduced along with some test statistics of the model underlying the static equations used to compute the equilibrium unemployment rate as well as the reduced-form equation for unemployment.

### The structural model

The structural model was estimated using a two-stage least squares estimator and the price- and wage-setting equations were estimated under the assumption that the error terms follow first order autoregressive processes. The corresponding parameters of residual serial correlation are B(15) and C(13), respectively.

*Instruments*: const, $\tau_{(-2)}$, $\Delta p^*$, $\Delta p^*_{(-1)}$, $a_{(-1)}$, $a_{(-2)}$, $(k-\bar{n}_p)_{(-1)}$, $(k-\bar{n}_p)_{(-2)}$, $g_{(-1)}$, $txne_{(-1)}$, $e_{(-1)}$, $(y-\bar{y})^*_{(-1)}$, $(y-\bar{y})^*_{(-2)}$, $p_{\mathrm{oil}}$, $p_{\mathrm{oil}(-1)}$, $i^*$, $\theta_{(-2)}$, $(n_g-l)_{(-2)}$, $r_{u(-1)}$, $\gamma_{(-1)}$, $\pi_{(-2)}$.

*Estimation period*: 1962–1993.

## Table B.1 The estimated parameters of the structural model

| | Coefficient | Std. Error | T-Statistic | Prob. |
|---|---|---|---|---|
| B(1) | 0.036 | 0.024 | 1.509 | 0.135 |
| B(2) | –0.184 | 0.054 | –3.425 | 0.001 |
| B(3) | –2.035 | 0.296 | –6.867 | 0.000 |
| B(5) | 0.758 | 0.205 | 3.704 | 0.000 |
| B(6) | 0.392 | 0.152 | 2.581 | 0.011 |
| B(8) | 0.254 | 0.117 | 2.182 | 0.032 |
| B(9) | –1.446 | 0.422 | –3.427 | 0.001 |
| B(12) | –0.980 | 0.238 | –4.121 | 0.000 |
| B(13) | 1.199 | 0.364 | 3.293 | 0.001 |
| B(15) | –0.561 | 0.190 | –2.953 | 0.004 |
| C(1) | –0.706 | 0.163 | –4.335 | 0.000 |
| C(2) | –0.280 | 0.084 | –3.327 | 0.001 |
| C(4) | 1.018 | 0.371 | 2.744 | 0.007 |
| C(5) | 0.092 | 0.026 | 3.585 | 0.001 |
| C(6) | 0.779 | 0.135 | 5.786 | 0.000 |
| C(7) | –0.458 | 0.245 | –1.869 | 0.065 |
| C(8) | –0.032 | 0.014 | –2.189 | 0.031 |
| C(10) | –0.356 | 0.115 | –3.109 | 0.003 |
| C(11) | 0.090 | 0.036 | 2.473 | 0.015 |
| C(13) | –0.488 | 0.210 | –2.322 | 0.022 |
| F(1) | 0.006 | 0.008 | 0.816 | 0.417 |
| F(2) | 1.192 | 0.178 | 6.706 | 0.000 |
| F(3) | –0.368 | 0.169 | –2.169 | 0.033 |
| F(5) | 1.130 | 0.408 | 2.768 | 0.007 |
| F(6) | –2.067 | 0.903 | –2.289 | 0.024 |
| F(7) | 1.041 | 0.581 | 1.794 | 0.076 |
| I(1) | 0.066 | 0.017 | 3.934 | 0.000 |
| I(2) | 0.207 | 0.056 | 3.716 | 0.000 |
| I(3) | 0.573 | 0.162 | 3.531 | 0.001 |
| I(4) | –0.002 | 0.001 | –2.025 | 0.046 |
| I(5) | –0.013 | 0.004 | –3.369 | 0.001 |
| I(6) | –0.335 | 0.157 | –2.126 | 0.036 |
| I(7) | 0.472 | 0.236 | 2.002 | 0.048 |
| I(8) | –0.293 | 0.161 | –1.822 | 0.072 |

*Price-setting equation:*

$$\Delta(p-w) = B(1) + B(2)q_{(-1)} + B(3)\tau_{(-1)} + B(2)(p-w)_{(-1)} + B(5)\Delta^2 p + B(6)(y-\bar{y})_{(-1)} + B(8)c_{(-1)} + B(9)\Delta(y-\bar{y})_{(-1)} + B(12)\Delta q + B(13)\Delta q_{(-1)} + [AR(1) = B(15)]$$

Observations: 31, *R*-squared: 0.9977, Adjusted *R*-squared: 0.9968, S.E. of regression: 0.0156, Durbin–Watson stat: 2.38.

*Wage-setting equation:*

$$\Delta(w-p) = Dq + C(1)\{(w-p)_{(-1)}-q_{(-1)}\} + C(2) + C(4)(n-l)_{(-1)} + C(5)\gamma_{(-1)} + C(6)\Delta((w-p)_{(-1)}) - C(6)\Delta q_{(-1)} + C(7)\Delta^2 p + C(8)\Delta u + C(10)c_{(-2)} + C(11)\theta_{(-1)} + [AR(1) = C(13)]$$

Observations: 31, *R*-squared: 0.9971, Adjusred *R*-squared: 0.9958, S.E. of regression: 0.0177, Durbin–Watson srat: 2.08.

*"Okun's Law" equation:*

$$(y-\bar{y}) = F(1) + F(2)(y-\bar{y})_{(-1)} + F(3)(y-\bar{y})_{(-2)} + F(5)(n_p-\bar{n}_p) + F(6)(n_p-\bar{n}_p)_{(-1)} + F(7)(n_p-\bar{n}_p)_{(-2)}$$

Observations: 32, *R*-squared: 0.8517, Adjusted *R*-squared: 0.8232, S.E. of regression: 0.0140, Durbin–Watson stat: 2.07.

*Current account equation:*

$$ca = I(1) + I(2)c + I(3)(y-\bar{y})^* + I(4)i^* + I(5)p_{\text{oil}} + I(6)(y-\bar{y})_{(-1)} + I(7)(y-\bar{y})_{(-2)} + I(8)(y-\bar{y})_{(-3)}$$

Observations: 32, *R*-squared: 0.5978, Adjusted *R*-squared: 0.4805, S.E. of regression: 0.0104, Durbin–Watson stat: 1.72.

The static equations were derived assuming that all adjustments to implied steady states have taken place; the only exceptions were that I did not eliminate the change in the inflation rate and the unemployment rate.

## The reduced-form model

The *reduced form for unemployment* was also estimated by two-stages least squares. The results are given below.

*Dependent variable:* $\Delta(u-l)$
*Estimation period:* 1962–1993. Observations: 32.
*Instrument list:* const, $(y-\bar{y})^*_{(0\text{ to }-1)}$, $p_{\text{oil}(0\text{ to }-1)}$, $g_{(-2)}$, $r_{u(-2)}$, $i^*_{(0\text{ to }-2)}$, $\theta_{(-2)}$, $\Delta\gamma_{(-1)}$, $txne_{(-2)}$, $e_{(-1\text{ to }-2)}$, $p^*_{(0\text{ to }-2)}$, $\tau_{(-1\text{ to }-2)}$, $\pi_{(-2)}$.

## Table B2. The estimated parameters of the reduced form model

| Variable | Coefficient | Std. Error | $t$-Statistic | Prob. |
|---|---|---|---|---|
| const | –1.009 | 0.572 | –1.763 | 0.09 |
| $\Delta\,ca$ | 6.436 | 2.092 | 3.076 | 0.01 |
| $r_{u(-2)}$ | 0.576 | 0.256 | 2.254 | 0.03 |
| $u_{(-2)}$ | –0.466 | 0.108 | –4.309 | 0.00 |
| $p_{\text{oil}(-1)}$ | –0.130 | 0.056 | –2.329 | 0.03 |
| $\Delta\gamma$ | –0.807 | 0.275 | –2.935 | 0.01 |
| $(y-\bar{y})^*$ | –5.558 | 1.533 | –3.626 | 0.00 |

$R$-squared: 0.7495, Adjusted $R$-squared: 0.6894, S.E. of regression: 0.1342, Akaike info criterion: –3.826, Schwartz criterion:–3.505, $F$-statistic: 13.112, Durbin–Watson stat: 1.78, Prob($F$-statistic): 0.000001.

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