Evaluating the effectiveness of the French work-sharing reform*

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This revision: December 2003

Abstract

We analyze the macroeconomic impact of the French work-sharing reform of 2000 (a reduction of standard working hours in combination with wage subsidies). Using a vector error correction model (VECM) for several labor market variables as well as inflation and output we produce out-of-sample forecasts for 2000/2001. A comparison of these forecasts –which serve as a benchmark simulation without shocks– to the realized values (with shocks) suggests significant beneficial employment effects of the policy mix. Output, productivity, hourly labor costs, and inflation are only transitorily affected or not at all.

JEL: E24, E27, C32
Keywords: unemployment, work-sharing, France, VECM, forecasting

Introduction

At the turn of the millennium, the French unemployment rate decreased by 2.2%-points (1999-2001) after almost a decade with stubbornly high unemployment. The

*We wish to thank participants of the Berlin Econometrics Seminar Helenenau 2002 and the EALE 2003 meeting as well as Michael Funke, Bernard Salzmann, and an anonymous referee for helpful comments, and Nathalie Fourcade for help with the data. We remain responsible for any mistakes.

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rest of the Eurozone experienced a decline as well, but the amount of 1.2%-points was markedly lower (see table 1 below for this and other basic data). An analogous feature can be observed in employment figures: after growth rates below average in 1995-1998, France experienced above average growth rates from 1999 to 2002. The rest of the Eurozone shared the tendency of the movement, but not quite its extent.

This development may partly be attributed to the above Eurozone-average real GDP growth in France (3.1% p.a. for France against 2.5% p.a. for the rest of the Eurozone in 1997-2001). However, another event that had an important influence on the economic environment of France during this period was a "work-sharing" reform. A brief summary of the complex institutional setting could read as follows: On the one hand, the standard workweek was reduced from 39 to 35 hours, first on a voluntary basis coupled with incentive schemes conditional on employment creation (Robien act 1996, Aubry I act 1998), then on a compulsory basis (Aubry II act 2000). On the other hand the following alleviations were offered to firms: Greater managerial flexibility w.r.t. working time allocation (Robien, Aubry I), payroll tax cuts through lower social security contributions for employers (in all acts), and a negotiated mid-term wage-income growth restraint.

The explicit aim of this policy mix was to reduce unemployment levels. Thus it is interesting to analyze if the reform had any noticeable effect or if it was just ordinary economic growth that did the work. Also from the perspective of economic theory it is still an unsettled issue whether a reduction of standard hours (alone) will lead to more or less unemployment since in most models the outcome is in general ambiguous.

Earlier studies provided mixed results, see below, and a general consensus has not been reached. Recent empirical assessments of other work-sharing reforms have not been positive, see Hunt (1999) or Crépon and Kramarz (2002). But the respective German and (1982) French reforms included persistent hourly wage hikes as income compensation for workers and no alleviation for firms, such that a negative outcome is perhaps not too surprising. However, the institutional environment of the recent French reform was quite different which provides an interesting new perspective on the issue of work-sharing.

The existing literature in the context of the recent French reform falls into roughly three classes: First there were ex-ante simulations of several scenarios within macro models (see e.g. DARES-BDF-OFCE 1998), second there is descriptive evidence (e.g. DARES 2002), and finally micro data for selective samples of firms have been thoroughly analyzed (Bunel 2002, Passeron 2002). For those studies that allow drawing conclusions with respect to employment effects, the results are mostly optimistic.

Our aim is to complement the literature by shedding some light on the macroeconomic consequences of the recent French work-sharing reform. To this end we estimate an empirical system of labor market variables (as well as inflation and output) using a vector error correction model (VECM) up to 1999. As there are two competing possible break dates (namely 1999q4 and 2000q1, see section 2 for detailed results), we actually estimate two slightly different models. With these models we can produce forecasts for the subsequent period, acknowledging the fact that those forecasts are based on the previous institutional regime. A comparison to actual observations after the introduction of the 35h work-week allows us to pinpoint significant changes. For both models we get very similar results; we find that unemployment was reduced throughout the analyzed horizon at the cost of a transitory depression of total hours worked reflecting a short-run rise of hourly labor costs. Labor costs and inflation remain somewhat higher than their forecasts, but not significantly so. Output is forecast quite well in every period and in general seems unaffected by the break.

These results can be attributed to the work-sharing-*cum*-lower-costs reform because other forces that could have had an influence on the unemployment rate were absent: For example, the labor supply grew at normal rates (see table 1), if not faster than usual. The coverage of active labor market policy actually declined somewhat during our forecast period (Boulard and Lerais 2002), which makes our estimates even slightly conservative. National accounts data show that other stimuli such as an expansionary fiscal policy did not occur, and not even ECB officials or other pro-Euro economists claimed that the introduction of the Euro would fight unemployment; instead they have always called for structural labor market reforms. (Also recall that the Euro was adopted already in the beginning of 1999, not in 2000.) But even if these arguments are disputed, our real output forecast turns out to be surprisingly accurate, such that potentially stimulating business-cycle effects would be controlled for.

Given the various preceding stages of the work-sharing reform project, our approach of dating the structural break not earlier than 1999q4 may be criticized, although the average working time data in table 2 reveal that notable changes only occur around 1999q4/2000q1. Furthermore, our model is empirically stable until the end of 1999 which also suggests that the effects of the preliminary reforms are not quantitatively important. But even in a broader sense an earlier and "smoother" structural break would *not* invalidate our analysis, either: The observed decrease in unemployment in 1998-1999 is not counted as caused by the reform by choice of our breakpoint date. Therefore, our tests are actually somewhat conservative in the sense that the overall impact of the reform in terms of unemployment reduction was probably even bigger.

The rest of this contribution is structured as follows: In the next section we provide a brief survey of the relevant literature, we describe some institutional details of the reform, and we discuss estimates and implications of existing studies. Section 2 presents the data and explains our model specification along with the resulting forecasts. The final section summarizes the main findings.

[Tables 1 and 2 about here]

1 Theory, institutions, and existing evaluations

1.1 A glimpse at the recent work-sharing debate

Economists are usually very critical about work-sharing due to the "lump-of-output" fallacy, i.e. the assumption of a fixed output that is often implicitly made in public debates on the issue (Layard, Nickell, and Jackman 1991, Snower 1997). But most of the theoretical studies yield ambiguous predictions depending on what the precise model or parameter values are.

For example, in the reference model of Calmfors and Hoel (1988) even the most skeptical model variant predicts a definitely negative response of output and employment only if the overtime premium is constant. If a progressive premium system is allowed instead, the response becomes uncertain. In the same spirit, although more optimistic, Askenazy ((2000a) and (2000b)) builds a model where work-sharing induces organizational changes that increase productivity. Depending on the functional form of disutility of labor and the associated wage compensation, the effect on employment changes its sign; the longer-run effect also depends on how the government finances the wage subsidies. In some cases even a Pareto improvement is possible, which is contrary to the conclusions in Marimon and Zilibotti (2000), where the possible employment creation would make firms worse off. A possible Pareto improvement is also found in the matching models by Rocheteau (2002) and by Ortega (2003), where lower standard hours may offer a firm more flexibility to react to demand shocks.

On the empirical side, the results are also mixed. Crépon and Kramarz (2002) analyze the reduction of 40 to 39 hours in 1982 for France. At that time full wage income compensation without reorganizational possibilities were the norm. Not surprisingly, they find a negative effect on employment. In Hunt (1999), a similar work-sharing reform without wage moderation is analyzed for West Germany in 1984-1994. Depending on the data set and method used she finds mostly insignificant coefficients of both signs. On the other hand, after conditioning on wages Franz and König (1986) find a positive partial effect of normal hours reduction on employment for West Germany (p. S241), although it is not mainly a study about work-sharing.[1]

In contrast to former experiences, the reforms that recently took place in France implied working-time reduction against tax cuts and subsequent wage growth restraint. This policy mix therefore may have been more effective than earlier work-sharing attempts.

1.2 The institutional background in France

The shortening of the standard work-week was implemented in several stages. A first act to reduce working-time (named Robien) was passed in June 1996. This voluntary measure had only very little impact, covering only 0.3m employees between June 1996 and June 1998 (Passeron 2002).

The main reform project was implemented in two stages. First the Robien act was replaced by the Aubry I act in June 1998 which lowered the legally standard work-week to 35 hours. (Martine Aubry was the minister of labor between June 1997 and October 2000.) However, while it became effective for firms with more than 20 employees on January 1^{st}, 2000, smaller firms would not be affected until January 2002. Additionally an incentive scheme was introduced to promote a quicker *effective* work-week reduction. The most important exception in this legislation was the civil service (but not state-owned firms). There were between 14.7m (Gubian 2000) and 16.2m (DARES 2002) employees potentially affected by this reform, of which about one third worked in small firms, and 15 to 18% were part-time workers. (The overall number of employees in France in 2001 was about 22m people.)

The Aubry I incentive scheme provided reductions of social security contributions for firms that effectively reduced the work-week and guaranteed a certain level of employment. A useful source for details is Passeron (2002). The criteria for subsidies are given by an effective reduction of working time by at least 10% and an increase of the employment level by 6%. (Alternatively, if mass layoffs had been planned, it was sufficient for a firm to retain a corresponding number of the workers under threat of displacement.) The subsidies were degressive and paid over a span of five years; for the first year they amounted to a reduction of social security contributions of EUR 1372 per year and employee. Each year they were reduced by EUR 152. The starting date was relevant, because there was a penalty for initiating the work-week reduction later. In the end of 2001 there were 8.6m (=53% of all potentially affected) workers in firms that had undertaken a work-week reduction (DARES 2002). Looking only at workers in bigger firms (that were subject to the reform already in 2000) reduces the number only to 8m (=73% of affected workers back then). Among the firms with a work-week reduction up to the end of 2001, 58% participated in the Aubry I incentive scheme (this covers 28% of all workers that had their working time reduced).

The second stage (Aubry II, passed in December 1999/January 2000) confirmed the 39 to 35 hours transition and instated a system of structural aids. For bigger firms the Aubry I scheme was terminated, whereas for smaller firms the end date is 2002. (More precisely, firms in the Aubry I scheme were subsidized until the end of the five-year schedule, they could even partially cumulate those subsidies with new structural aids; but no new Aubry I subsidy plans could be started after 2000.) The structural aids depend on wages and vary from EUR 3278 per year and employee at the minimum wage (*SMIC*) to EUR 610 for employees that earn more than 1.8 times the SMIC. The average tax reduction per year and employee was EUR 1067 (Passeron 2002); it was not conditional on guaranteed employment levels anymore. Furthermore the method of working time accounting could be changed such that effective reductions of less than 10% ($35/39-1$) were possible to qualify for the structural aids. For a transition period of one year, the overtime premium for the first four hours was reduced from 25% to 10% (the next four hours remained subject to a 25% premium and hours beyond that must be paid at 150%). This working-time reduction was mostly accompanied by an initial wage income compensation (Pham 2002). For example, in 2000 98% of all employees covered by Aubry II enjoyed a full wage compensation; however, 1/3 had to accept a wage stagnancy and 14% a wage growth moderation for the following one to three years.

Finally it should be noted that in some sense the work-sharing experience is already history, because right after its election in early 2002 the new center-right government started to reverse the reform.

1.3 Previous evaluations

Apart from the already mentioned descriptive evidence in DARES (2002), the macroeconomic studies for France were based on ex-ante simulations e.g. within the macro models of the OFCE institute, the central bank (BdF), and the ministry of finance (MINEFI). For a survey see Conseil Supérieur de l'emploi, des revenus et des coûts (1998), DARES-BDF-OFCE (1998), or Commissariat Général du Plan (2001). Depending on the various assumptions the assessment of the employment effects range from optimistic (up to 700,000 additional jobs in the simulations of the OFCE and of the BdF) to more sceptical (between 200 and 300,000 according to MINEFI, even negative if a blockade between unions and employers is assumed).

The existing microeconomic studies (based on observed data) use samples of firms with specific characteristics, apparently due to data limitations. Passeron (2002) for example analyzes firms in the Aubry I scheme. He concludes that the employment gain induced by the work-week reduction is between 6 and 7.5%. (Effects where firms become eligible for subsidies without having been *induced* to meet the criteria are supposedly not included in those numbers.) The resulting productivity gain effect in his study is 4% which is a little more than what is often assumed on theoretical grounds. According to Passeron the government subsidies decreased total labor costs of the Aubry I firms by 4%. A similar conclusion can be found in Gubian (2000); when cumulated with the Aubry II aids this even amounts to 5.5%. In addition to that, Passeron calculates an "anti-seniority" effect, because newly hired workers are cheaper relative to their productivity. He estimates that effect as 1%. Wage moderation is seen as 0.8%, such that altogether the work-week reduction was approximately cost-neutral over the horizon until 2002. This study is not representative for aggregates, and even the sign of the bias is ambiguous: Firms that waited until 2000 to reduce their work-week receive less subsidies (ca. 2 to 3%, Gubian 2000), but also have less restructuring costs because the effective reduction may then be less than 10%. Another study for selected firms is Bunel (2002), who uses a special data set ("Passages") to compare all firms that reduced their *effective* weekly working time to 35 hours.

The existing macro evidence is for scenarios, and it is not clear what the selective micro studies imply for the aggregate level. An empirical macro analysis therefore seems useful, and we now turn to it.

2 Empirical methods and results

2.1 Variables and data

In the context of imperfect competition, prices and wages are set simultaneously by economic agents, see e.g. Layard, Nickell, and Jackman (1991), where the unemployment rate gives some feedback to the system. Therefore we require the following variables: wages, prices, an employment measure, real output, and the unemployment rate.

For wages there are basically three possibilities: 1. Total compensation including social security contributions paid by employers, which represents the total cost of a labor unit; 2. "gross" wages that include only social security contributions and taxes paid by the worker herself; 3. net wages without taxes or contributions. For price setting and the labor demand total compensation is clearly the appropriate variable. For wage setting the level of net wages could also be relevant, but we restrict ourselves to the total compensation measure, not least because net wages is a time series which is difficult to obtain.

Prices: For labor demand and price setting it is the GDP deflator which is important, while adding the consumer price index (CPI) would make sense for wage setting analysis. Here we chose to include only the GDP deflator. Note that both price indices display roughly the same development (not shown), such that the exclusion of the CPI is not problematic.

Employment can be the labor volume (in hours) or the number of employed people, where average working time links the two concepts. As the hours worked per person is a central variable for the present study we include both employment measures. For real output it is natural to choose real GDP, and the chosen unemployment rate is the one according to ILO definitions.

W.r.t. the selected variable set it might be argued that potentially important variables are missing. Apart from the price/tax wedge and import prices, there are many possible extensions. For example capital user costs may play a significant role as part of marginal costs, and labor demand could also depend on the sectoral composition of output. However, in this paper we do not follow up on these issues because our aim is to work with a manageable labor market model. By analyzing the described data set we have tried to follow a pragmatic middle-of-the-road approach.

The data are from the following sources: OECD, the French statistical office (*INSEE*), and the statistical department of the French labor ministry (*DARES*); all series are seasonally adjusted. The variable names, a survey of the sources and calculation methods are given in table 3. We present the analyzed time series in figure 1 with some anticipated transformations: $LWReal_t = LW_t - LPY_t$ will denote the hourly real wage, and $Infl_t = \Delta LPY_t(*400)$ is essentially the first difference of the log price level, i.e. the inflation rate.

Our general sample choice is determined by two facts: Data on hours worked are only available from 1980 on, and it turns out that there is instability in the unemployment rate equation already towards the end of 1999, see the system analysis below for more details. Thus the estimation period is set to 1980q1-1999q3, where the first observations will be used as starting values for the necessary lagged regressors.

[Table 3 about here]

[Figure 1 about here]

2.2 Preliminary univariate data analysis

Obviously the inflation rate displays a very persistent behavior, such that using the price level in the model would be difficult. But the inflation rate itself can well be included as an integrated series ($I(1)$) in the multivariate system. Instead of the nominal wage we therefore include the real wage “LWReal” ($\equiv LW - LPY$) in the system. The variable vector is thus given by

$$y_t = (UR_t, LEMP_t, LY_t, Infl_t, LWReal_t, LVol_t)' .$$

The results of standard unit root tests are shown in table 4. For all series the null hypothesis of a unit root *cannot* be rejected, with the exception of the labor volume. For the sample starting in 1981q3 the unit root is not rejectable; but from 1982q1 on with up to seven lags it would have to be rejected on the 5% level. This does not pose any problems *per se*, but another test in the system context would still be interesting (see below).

[Table 4 about here]

Before moving to the system analysis we perform a simple univariate forecast of unemployment after the reduction of standard hours. Here we apply the standard ARIMA model for the sample 1978q2-1999q3. We impose the unit root restriction and after eliminating insignificant terms arrive at an ARIMA(3,1,2) specification. The forecasts derived from this model beyond 1999q3 are displayed in figure 2 along with the forecast error confidence bands and the actually observed development. The development of UR is clearly overestimated; actually, this model treats the unemployment rate more or less as a random walk and thus simply sets the forecast close to the last observed value. This automatically raises the question whether the information contained in other variables enables us to produce better forecasts, or if the forecast failure is due to a policy-induced structural break.

[Figure 2 about here]

2.3 The multivariate forecasting model

Our statistical framework is the vector autoregressive model (VAR) with Gaussian innovations that is widely used in empirical macroeconomics.[2] We combine the $n = 6$ variables in the column vector y_t for $t = 1, ..., T$, where in our case T refers to 1999q3 or 1999q4, depending on the model variant, see below. Then the VAR has the following shape:

$$y_t = \sum_{k=1}^{K} \Phi_k y_{t-k} + \tau t + \mu + \epsilon_t \tag{1}$$

As deterministics a constant μ and a linear trend (serving as a proxy for technical progress etc., with coefficient τ) are allowed. After the appropriate reparametrization the following vector error correction model (VECM) is obtained:

$$\Delta y_t = \Pi y_{t-1} + \sum_{k=1}^{K-1} \Gamma_k \Delta y_{t-k} + \tau t + \mu + \epsilon_t \tag{2}$$

We adopt the conventional denomination for the matrix of the cointegration vectors, β, and the matrix of adjustment coefficients, α, both of dimension $n \times r$ for a given cointegration rank r, such that $\Pi = \alpha\beta'$. The linear trend can only appear in the cointegrating relations, i.e. we impose $\tau = \alpha\rho'$ (ρ freely varying).

First a choice about the number of lags in the VAR needs to be made. We use a maximum of six lags, as this already means estimating 36+2 parameters in each equation. The Schwartz and HQ information criteria suggest $K = 2$, only the (inconsistent) Akaike criterion chooses $K = 6$. However, because of remaining residual autocorrelation with two lags we are led to a choice of $K = 3$. A single outlier in 1984q1 (especially in the unemployment equation) distorts the otherwise Gaussian properties of the innovations, such that we include a restricted impulse dummy for that observation.[3] Then the residual diagnostics are fully satisfactory, see table 5.

[Table 5 about here]

The rank of the matrix Π is an important property of the system, although it is not as crucial for forecasts as it would be for other analyses. We apply the well-known Johansen procedure accounting for a restricted trend, see table 6. Johansen (1995) shows that this test procedure is unbiased if the rank tests are interpreted as a sequence, starting from rank zero and stopping at the first insignificant test statistic. At first sight the standard trace test statistics (in the third column) are all significant, and thus the conclusion would be a (trend) stationary VAR without any unit roots in the system. Apart from the fact that this would contradict the univariate unit root test evidence, it would be a highly unusual finding for macroeconomic time series.

However, the rank test is substantially oversized in small samples (i.e. it rejects too often although the null hypothesis is true), and thus this nominal result may be exaggerated. Fortunately, there exists a novel method to investigate this suspicion. Following the Bartlett correction principle, Johansen (2002) develops a correction factor w.r.t. the rank test statistic. The idea is to find the expected value of the test statistic for given models in small samples, compare that to the asymptotic value and derive the corresponding correcting factor. It is obvious that this factor depends on nuisance parameters that are asymptotically (and hence for the standard test setup) irrelevant. The factor is applied to the measured test statistic and thereby the bias of the test is decreased. Simulation studies show a beneficial effect.

In order to calculate the Bartlett corrections for the last four null hypotheses, we estimated the VAR four times with the respective rank under the null to obtain the necessary estimates of the parameter matrices. These were entered into the provided program; the lag length is still fixed at $K = 3$. The results of this procedure are provided in the last column of table 6, and the test conclusions change considerably: We have to stop at $H_0 : r = 4$ when interpreting the sequence of corrected trace tests, as this is the first non-rejected hypothesis. This choice implies that $n - r = 2$ independent stochastic trends drive the system which is a reasonable property of such a model. The results are not entirely straightforward because of the fact that – viewed in isolation– $H_0 : r = 5$ is also rejected. However, this is irrelevant for the appropriate testing strategy of the Johansen procedure as explained before. Unreported evidence about the estimated characteristic roots of the system also supports the choice of exactly two unit roots. Therefore we proceeded with an imposed cointegration rank of $r = 4$.

[Table 6 about here]

It is beyond the scope of this paper to provide a structural analysis of the French labor market. This aspect and the relatively high cointegration rank induced us to refrain from an economic interpretation of the cointegrating relationships. In this sense the specified components of the cointegration space are arbitrary and without economic content. However, no harmful normalizations were used, e.g. no zero parameters were normalized to unity.

We can now investigate a number of hypotheses within the cointegrated VAR: First we return to the trend stationarity of the (log) labor volume. We test this as the hypothesis that LVol alone is one of the components of the cointegration space. The LR test of this restriction cannot reject, with $\chi^2(2) = 3.76, p = 0.15$, which confirms the univariate evidence. A second interesting question is whether the linear trend is actually needed in the cointegrating relations. The corresponding exclusion is clearly rejected ($\chi^2(4) = 32.1, p = 0.00$), so the trend is essential for an adequate model. Finally, note that output does not adjust to any equilibrium deviations (tested as a zero row in the α matrix, $\chi^2(4) = 3.56, p = 0.47$), and is thus weakly exogenous for the long-run parameters. This partly explains why we find that output is mostly unaffected by the structural break, see below.

2.4 Stability of the model and choice of breakpoint

The final check of the model is about its stability. Especially for our purposes a stable specification in the estimation period is obviously a desirable feature because an unstable model would not yield meaningful forecasts.

Several recursively estimated test statistics indicate that parameter stability for the described sample up to 1999q3 clearly holds, see figures 3 through 5. This means that the quantitative impact of the previous reforms –Robien and Aubry I as described before– was small and could easily be subsumed under the error term.

[Figures 3, 4, and 5 about here]

However, based on the *a priori* information about the beginning of the reform we originally conjectured that any potential structural break should have happened in 2000q1. But when we add the observation 1999q4 to the estimation period, there is clear-cut evidence for instability at least in the unemployment equation (see figure 6), reflecting an unusual decline of the unemployment rate. As the coming reform was publicly known in the end of 1999, this suggests that announcement effects were already at work; the favorable product demand environment during that time probably helped, too.

[Figure 6 about here]

Given these findings, our strategy for the forecast test is to check both breakpoints 1999q4 and 2000q1 in the following way: For 1999q4, we simply estimate the model until 1999q3 as described before and start our system forecast immediately afterwards. For the 2000q1 break date variant, we extend the estimation period until 1999q4, but introduce an impulse dummy for the last observation to account for the significant stability failure documented in figure 6. (This impulse dummy is not restricted to the cointegration space.) Note that the second variant is quite conservative in the sense that *no* movements before 2000q1 are attributed to the work- sharing reform.

Before we apply the forecasting model, note also that several robustness analyses were done in the course of preparing this paper, varying both the number of lags, the cointegration rank, and applying several model reduction strategies. The forecasts on which our interest is centered were always very similar.

2.5 Forecasts and reality

Now we are in the position to answer the central question of this study, namely if the policy of reducing standard hours in combination with wage subsidies had a significant influence on the unemployment rate and other variables. To this end we compare the *observed* development of the vector y_t until 2001q2 (i.e. $y_{T+1}, y_{T+2}, \ldots, y_{T+h}$) and the corresponding dynamic forecasts $y^f_{T+1}, y^f_{T+2}, \ldots, y^f_{T+h}$. For the first variant we have $T = 1999q3$ and $h = 7$, whereas the second variant uses $T = 1999q4$ and $h = 6$ (with an impulse dummy in 1999q4). Our approach is related to the test of Box and Tiao (1976) which however is about the *joint* significance of the forecast errors; a more detailed interpretation is possible by considering the forecast errors separately. (The variance formulae for the entire sequence of forecast errors can also be found in Clements and Hendry (1998).)

In line with our strategy of considering both 1999q3 and 1999q4 as potential breakpoints, we present two sets of corresponding graphs in figures 7 and 8. There are some quite interesting results to be pointed out:

[Figures 7 and 8 about here]

- The forecast of the unemployment rate (UR) falls slightly, but its confidence bands drift apart quickly and become extremely large. *In spite of this* the observed development of unemployment is *significantly* lower than the forecast.

- Consequently, the employment development (LEMP) is consistently underestimated. At the end of the forecast horizon the discrepancy of the forecast w.r.t. reality constitutes about 0.5 million additional employed workers.

- The real output (LY) is forecast surprisingly well, so in this sense no extraordinary goods market developments were responsible for the fall of the unemployment rate.

- There is nothing important to be seen in the inflation path, especially no dramatic rise because of rising labor costs. However, the reliability of the forecast is very low, given the extremely wide confidence bands stretching far into the negative range.

- W.r.t. real hourly labor costs (LWReal) there is an unpredicted increase in the beginning despite the paid subsidies. (The maximal difference between reality and forecast is about 2%.) This is not surprising since an initial income compensation had been negotiated which represents a short-run hourly wage hike. However, this is gradually eliminated by subsequent wage growth restraint.

- Finally we observe a quite drastic slump of the labor volume (LVol) especially in 2000q1, although this is slowly offset in the following quarters. It seems as if the path of the labor volume relative to its forecast were partly a mirror image of the development of labor costs. One would probably not have expected that the labor volume after seven quarters is hardly distinguishable from its (forecast) normal path.

During the forecast horizon higher labor costs correspond to higher productivity ($LY - LVol$) and somewhat higher inflation, although not all these deviations are significant. Towards the end of the forecast horizon average hourly productivity is back on track, a fact which also holds for labor costs, at least in the 2000q1-breakpoint specification.

These developments make sense in a model where the increased individual productivity due to less working hours is subsequently offset by lower productivity of the formerly unemployed. It also implies that the labor cost subsidies were an important factor of the policy mix.

As was already mentioned in the introduction, other obvious effects that might have affected unemployment were not present during the forecast period. For example, there was no hidden expansion of active labor market policies; the number of the covered workers even dropped in 2000 (Boulard and Lerais 2002). Overall fiscal deficits were also reduced in comparison with the 1998-1999 period (source: INSEE national accounts data). There was no sign of other labor supply shifts happening (again, see the labor force developments in table 1), and the overall economic climate is captured quite well by our implicit empirical output model.

Hence we conclude that there exists relatively strong evidence that the French work-sharing-*cum*-labor-cost-subsidies reform was responsible for the lower unemployment rate at least in 2000-2001 and in this sense was successful.

3 Summary

As the effects of a reduction of standard hours are not predictable on purely theoretical grounds, it is the task of empirical studies to determine the efficacy of such policy options. The present paper provides evidence for the case of France, where a reduction of weekly standard hours in the beginning of 2000 was accompanied by subsidies of the social security contributions.

Detailed firm level data only exist for subgroups of firms that are not representative; also, labor demand is only part of the story behind unemployment developments. Therefore we used aggregate data to measure possible influences of the reform. We specified an empirical macroeconomic labor market model and examined the differences between observed data and the dynamic model forecasts in the horizon between the end of 1999 and mid 2001. Given that the economic environment is either captured within our model (most importantly demand conditions) or remained stable in France during our forecast period (most importantly active labor market policy), the effect can be attributed to the mentioned reforms.

Our analysis of the development of unemployment and other variables in France imply that the reduction of standard hours in combination with the offered wage subsidies was at least partly successful. This finding is significant in the sense that it holds after accounting for the forecast uncertainty. Although there was a short-term wage push and the labor input volume (in hours) displayed a sudden slump (implying higher productivity in the short run), wages and labor demand afterwards slowly recovered from this disturbance. Together with the shorter working hours, this meant that the employment level grew faster than its forecast. But above all, the unemployment rate fell more than would have been predicted on the basis of the old policy regime. Real output as well as the inflation rate seemed relatively unaffected.

All things considered, the optimism of several French authorities seems to have been well founded, and the mix of imposing restructuring costs on firms while at the same time offering cost alleviation in favor of unemployed workers was a good choice. But of course this relatively radical reform only helped roughly one fifth of all unemployed, and unemployment in France remained a mass phenomenon.

Notes

[1]This list of empirical studies is of course not complete, but conveys the status quo of the research, namely that evidence is diverse, and that a consensus has not been reached. Further references are given in Ortega (2003), Marimon and Zilibotti (2000), and Hunt (1999). The link between actual hours and legal hours was studied for the U.S. by Trejo (1991) and Costa (2000), arguing that work-sharing is irrelevant if firms and workers negotiate a constant compensation/workload-package. They find non-neutrality of lower working hours, whereas in Trejo (2001) irrelevance of work-sharing cannot be rejected.

[2]For a textbook treatment see Lütkepohl (1991), and for the theory of a VAR with cointegration Johansen (1995). The reported results were computed mainly with Pc-Give 10, see Doornik and Hendry (2001), with the exception of the Bartlett correction of the rank test which was performed with the program for RATS mentioned in Johansen (2002).

[3]"Restricted" means that the impulse dummy $i84q1_t(\{\ldots,0,0,1,0,0,\ldots\})$ is only allowed in the levels of the data and not in the differences. This is achieved by including $i84q1_t$ in the cointegration space and its difference $\Delta i84q1_t$ unrestrictedly in the VAR.

A Data Appendix

The DARES publishes the average working time of full-time employees in each quarter on the basis of the ACEMO survey carried out among employers. However, this data covers only plants with more than ten employees in the non-agricultural private sectors (hereafter competitive sector, excluding civil service, health services, etc. with about 6 to 7 million employees in the 1990's). This is equal to the average working-time of all employees under two assumptions: first full-time employees of small plants (less than 10 employees) work as long as those in medium and large companies, and second, the working time in the competitive sector is the same as in the rest of the economy. (As we use the log of the data, a weaker assumption is actually sufficient, namely that the ratio of the different working hours is constant.)

The INSEE provides the number of full-time equivalent employees, thus part-time effects are corrected for.

The volume of paid hours is thus the product of the average working time of all full-time employees and the number of full-time equivalent employees.

However, starting in 1998 the effects of the shortening of the work-week became noticeable, which up to 2002 concerned almost exclusively bigger firms. We therefore applied a correction which only transmits part of the working time changes (published by the DARES and concerning bigger plants) to the working time of all plants. Thus we modify our first assumption by holding the working time of small plants unchanged. (At the end of 2000 only less than 5% and at the end of 2001 less than 10% of all small plants ($<$20 employees) had reduced their working time, supporting our modified first assumption (Pham 2003).) As bigger firms employ about two thirds of all employees (source: Eurostat, News releases, Memo No 01/99, 10 March 1999) the correction beginning in 1998 is:

$$g[WorkingTime_{\text{all full-time employees}}] = 2/3 * g[WorkingTime_{\text{published by DARES}}]$$

with $g[.]$ denominating the quarterly growth rate.

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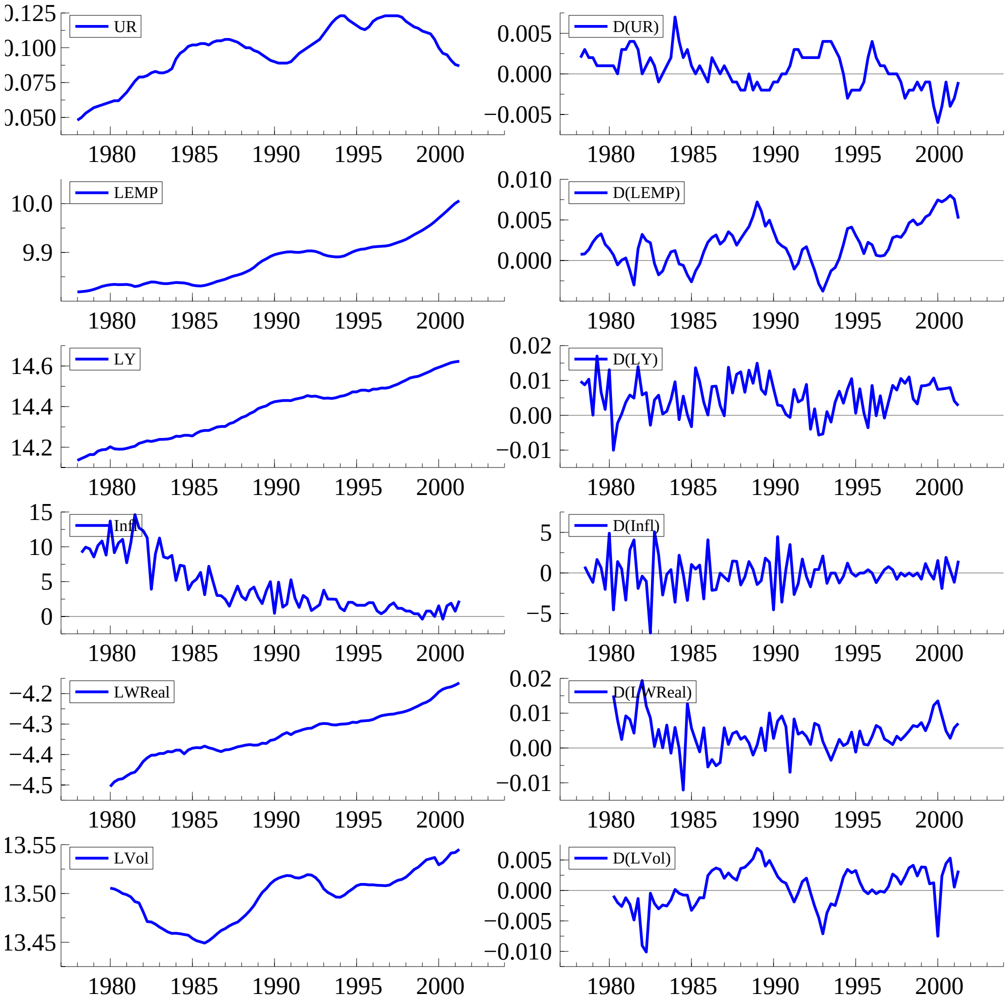
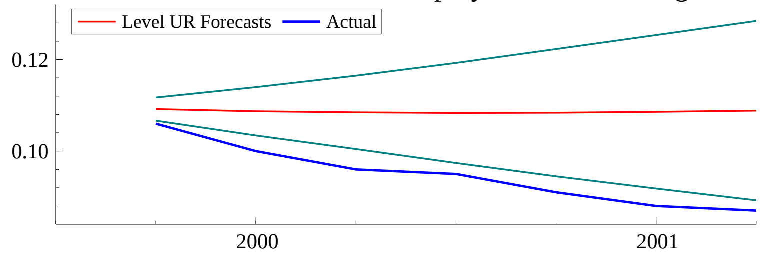
Figures

Figure 1: Time series graphs



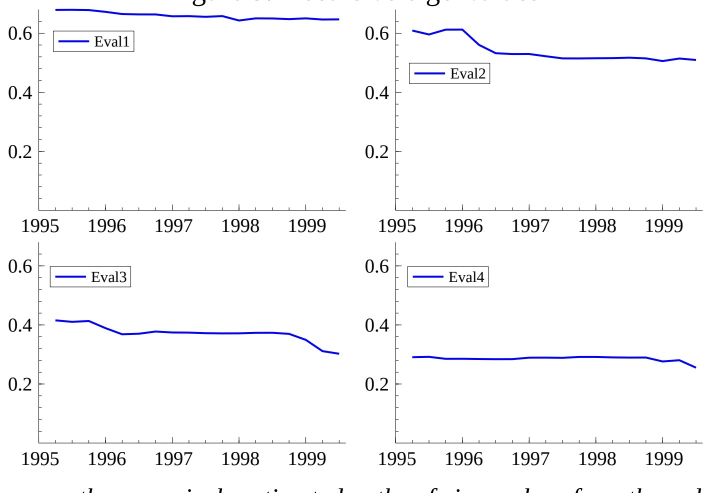
Remarks: *Left panel: the components of the vector* y_t*. Right panel: the time differences. The inflation rate is expressed in (approximate) annual growth rates.*

Figure 2: Univariate forecast of the unemployment rate starting in 1999q4



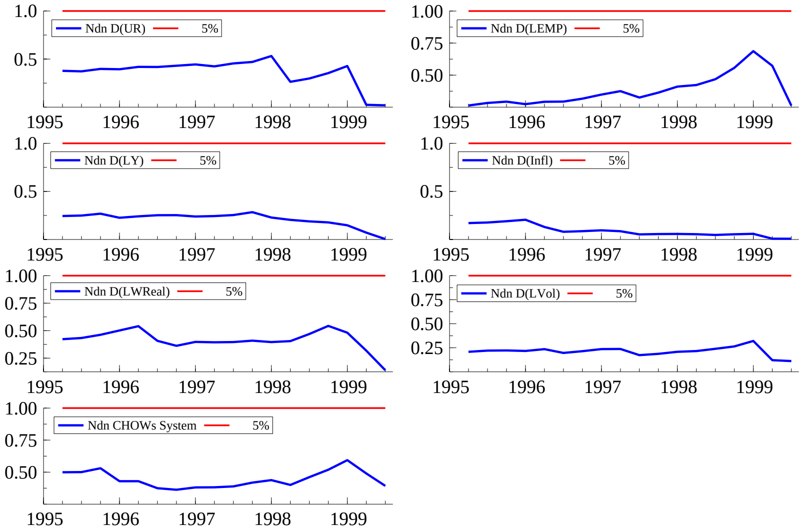
Remarks: *Forecasts of the unemployment rate (UR) in France 1999q4-2001q2. The (95%) forecast error confidence bands take into account the innovation variances as well as parameter uncertainty.*

Figure 3: Recursive eigenvalues



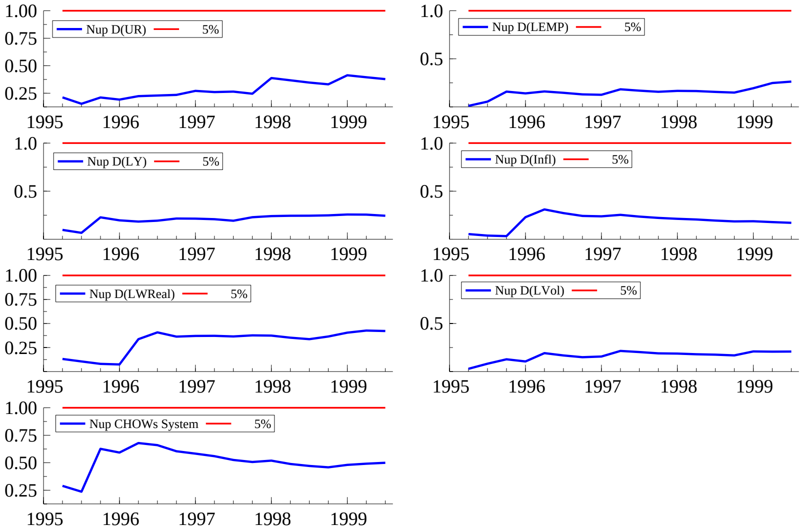
Remarks: *These are the recursively estimated paths of eigenvalues from the reduced rank regression under the restriction of cointegration rank $r = 4$; fluctuating estimates would hint at instabilities of the cointegration space.*

Figure 4: Chow breakpoint tests



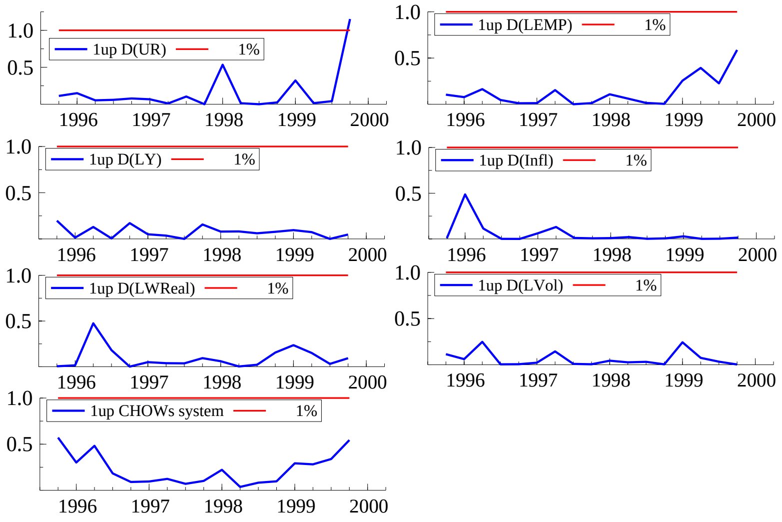
Remarks: *Test statistics based on forecasts from the respective dates on the x-axis until the end of the sample, and scaled relative to the 5% critical value which is normalized to unity. The underlying model is the one described in the text, with* $K = 3$ *lags and cointegration rank* $r = 4$.

Figure 5: Chow forecasts tests



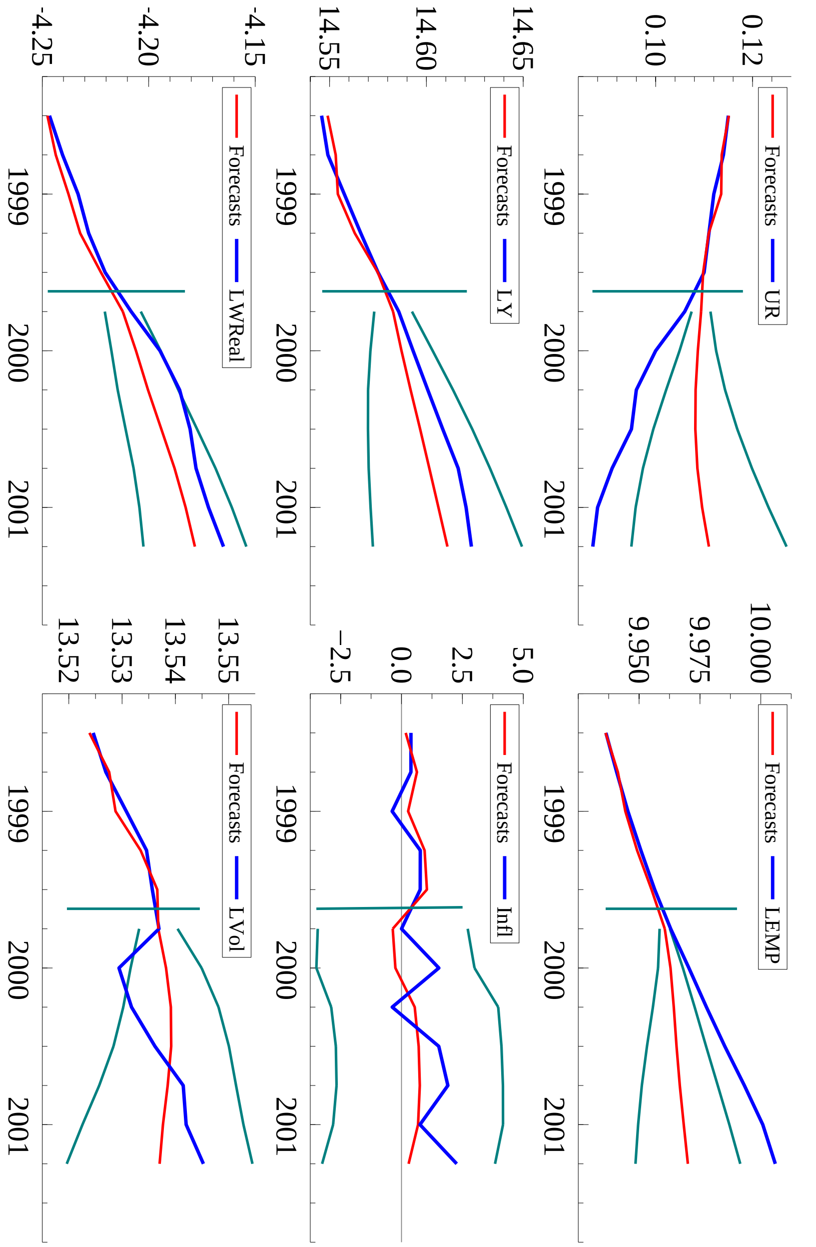
Remarks: *Test statistics based on forecasts from a fixed date on the left of the x-axis until the respective dates on the x-axis, and scaled relative to the 5% critical value which is normalized to unity.*

Figure 6:



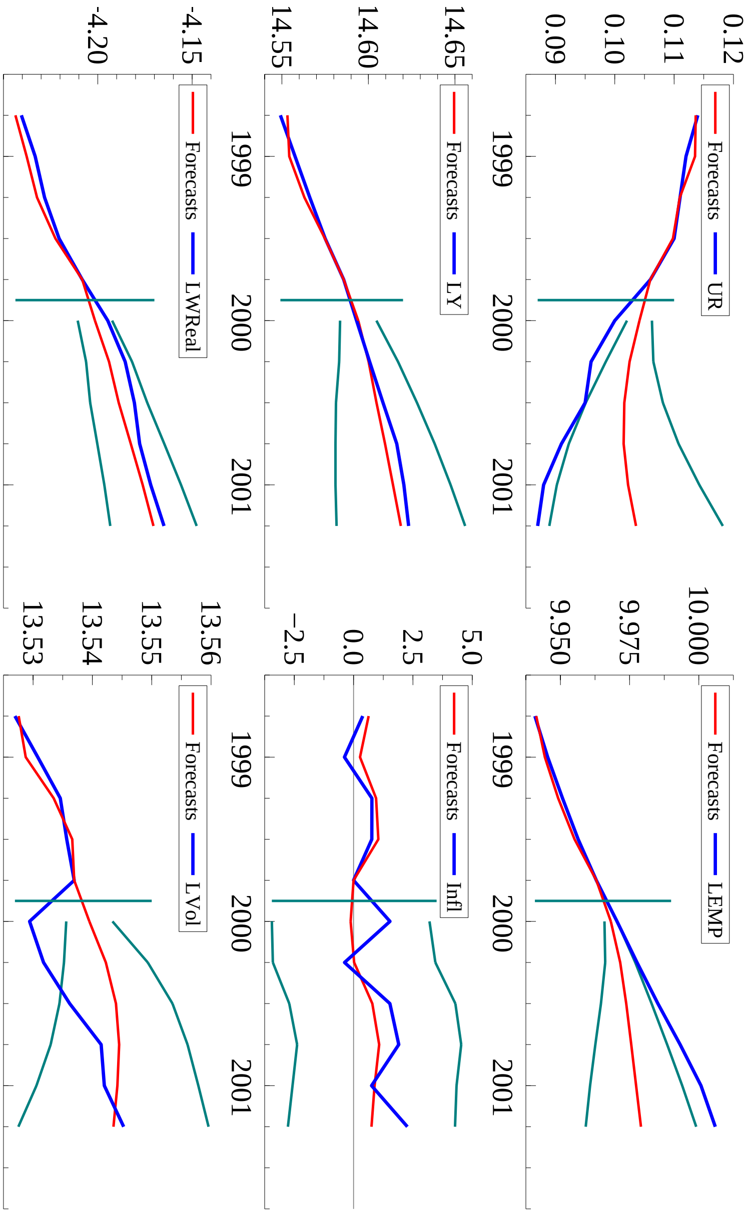
Remarks: *The Chow test statistics are based on recursive 1-step forecast errors that are scaled relative to the (1%) critical value which is normalized to unity.*

Figure 7: VECM forecasts, break date 1999q4



Remarks: *The bands denote the 95% forecast confidence intervals taking into account the innovation variance as well as parameter uncertainty.*

Figure 8: VECM forecasts, break date 2000q1



Remarks: *The bands denote the 95% forecast confidence intervals taking into account the innovation variance as well as parameter uncertainty. The estimation sample includes the observation 1999q4, but corrected with an impulse dummy due to the stability failure documented in figure 6.*

Tables

Table 1: Comparison of the macroeconomic context between France and the Eurozone

	EU	EU w/o Fr	France	EU	EU w/o Fr	France	EU	EU w/o Fr	France	EU	EU w/o Fr	France
	real GDP[1]			Employment[1]			Labor costs[1,2]			Population[1]		
1990			2.6			1.0						0.5
1991			1.0			0.2			2.8			0.5
1992	1.5	1.5	1.5	-0.9	-1.0	-0.7	7.3	7.7	5.9	0.5	0.5	0.5
1993	-0.8	-0.8	-0.9	-2.0	-2.2	-1.2	3.2	2.4	5.7	0.5	0.5	0.4
1994	2.4	2.4	2.1	-0.1	-0.1	0.2	2.5	2.5	2.2	0.3	0.3	0.4
1995	2.3	2.4	1.7	0.6	0.6	0.5	3.4	3.4	3.4	0.3	0.3	0.4
1996	1.4	1.5	1.1	0.6	0.6	0.5	2.8	2.9	2.3	0.3	0.3	0.3
1997	2.3	2.5	1.9	0.8	0.9	0.4	0.3	0.3	0.4	0.3	0.3	0.3
1998	2.9	2.7	3.4	1.9	1.9	1.7	0.8	0.5	1.8	0.2	0.2	0.4
1999	2.8	2.7	3.2	2.1	2.0	2.1	2.8	2.7	3.0	0.3	0.2	0.4
2000	3.5	3.4	3.8	2.3	2.3	2.5	2.7	2.8	2.2	0.4	0.3	0.5
2001	1.4	1.3	1.8	1.6	1.4	2.3	2.7	2.8	2.5	0.4	0.4	0.5
2002	0.8	0.7	1.2	0.4	0.3	0.6	2.5	2.5	2.6	0.4	0.3	0.5
	Inflation rate (CPI)[2]			Employees[1]			Productivity[1,5]			Population (15-64)[1]		
1990						1.5			1.6			
1991	4.3		3.4			0.8			0.8			0.0
1992	3.8		2.5	-1.2	-1.4	-0.1	2.5	2.6	2.2	0.5	0.5	0.2
1993	3.4		2.2	-1.7	-1.9	-0.8	1.2	1.4	0.4	0.4	0.5	0.2
1994	2.8		1.7	-0.4	-0.6	0.6	2.4	2.6	1.9	0.3	0.3	0.2
1995	2.6		1.8	0.6	0.6	0.9	1.6	1.8	1.2	0.2	0.2	0.2
1996	2.3	2.3	2.1	0.5	0.5	0.8	0.8	0.9	0.6	0.2	0.2	0.3
1997	1.7	1.7	1.3	0.8	0.9	0.5	1.5	1.5	1.5	0.2	0.2	0.3
1998	1.2	1.3	0.7	2.0	2.0	1.9	1.0	0.8	1.7	0.2	0.1	0.3
1999	1.1	1.2	0.6	2.1	2.1	2.1	0.7	0.6	1.1	0.1	0.1	0.3
2000	2.1	2.2	1.8	2.4	2.3	2.8	1.1	1.1	1.2	0.2	0.2	0.4
2001	2.4	2.5	1.8	1.6	1.4	2.3	-0.2	-0.1	-0.4	0.3	0.2	0.4
2002	2.3	2.3	1.9	0.6	0.5	1.0	0.4	0.4	0.5	0.3	0.3	0.5
	Real eff. exch. rate (CPI)[2]			Unemployment rate[2,3]			Unit labor costs[1,2]			Labor force[1,4]		
1990	8.7		3.2			8.6			-1.6			
1991	-2.8		-3.6			9.1			2.0			0.6
1992	3.7		1.3			10.0	4.8	5.1	3.6	0.0	-0.1	0.4
1993	-5.3		-2.2	10.1	9.9	11.3	2.0	1.0	5.4	0.0	-0.1	0.2
1994	-0.5		-0.6	10.8	10.6	11.8	0.1	-0.1	0.3	0.4	0.3	0.7
1995	6.2		1.4	10.6	10.4	11.3	1.8	1.6	2.2	0.4	0.4	0.3
1996	0.7		1.2	10.8	10.6	11.9	2.0	2.1	1.7	0.8	0.7	1.0
1997	-6.9		-2.9	10.8	10.6	11.8	-1.2	-1.2	-1.1	0.8	0.8	0.4
1998	2.7		0.4	10.2	10.0	11.4	-0.2	-0.3	0.1	1.3	1.4	1.0
1999	-3.8		-2.4	9.4	9.1	10.7	2.0	2.1	1.9	0.9	0.8	1.3
2000	-8.2		-5.8	8.5	8.3	9.3	1.5	1.7	1.0	1.1	1.1	1.0
2001	2.6		-0.5	8.0	7.9	8.5	2.9	2.9	3.0	0.9	0.8	1.0
2002	3.4		0.2	8.4	8.3	8.8	2.1	2.1	2.1	0.9	0.8	1.1

All numbers are growth rates (yoy) in %, except for the unemployment rate, which is in level and in %. "EU" is the Euro area.

[1] AMECO, own calculations

[2] EUROSTAT, own calculations

[3] harmonized

[4] labor force statistics

[5] measured as output per head

Source: AMECO, Eurostat, own calculations

Table 2: Different measures of working time

	1999q1	1999q2	1999q3	1999q4	2000q1	2000q2	2000q3	2000q4
ACEMO (2003)[1]	38.6	38.6	38.3	38.0	37.2	36.9	36.8	36.6
	-0.1	*-0.2*	*-0.6*	*-0.7*	*-2.2*	*-0.7*	*-0.4*	*-0.4*
DARES (2003)[2]	36.5	36.5	36.3	36.1	35.7	35.4	35.4	35.3
	-0.3	*0.0*	*-0.5*	*-0.6*	*-1.1*	*-0.7*	*-0.2*	*-0.2*
our data[3]	38.7	38.7	38.6	38.4	38.0	37.7	37.5	37.4
	-0.1	*-0.1*	*-0.3*	*-0.5*	*-1.0*	*-1.0*	*-0.4*	*-0.3*

The numbers in italics are quarterly growth rates in %.

[1] Working time published by the MES-DARES from the poll ACEMO. This data refers to firms with more than 10 employees and full-time employees. It stems from the DARES database (as of November 2003).

[2] Working-time calculated by the MES-DARES correcting the results of ACEMO for firms with less than 10 employees and part-time employees. It corrects additionally for a statistical break of the definition of the working-time in 2000 induced by AUBRY II. This figure was published in the DARES database (as of November 2003). Note that this series could not be used in our analysis because it only dates back to 1993.

[3] See the appendix for the exact calculation.

Table 3: Description of the data

Abbrev.	Meaning	Source / details
UR	unemployment rate	Standardized unemployment rate (ILO concept) from the OECD.
LEMP	number of employed workers (log of)	Source INSEE.
LY	real GPD (1995 prices, log of)	From OECD Main Economic Indicators (MEI).
LPY	GDP deflator (log of)	From OECD MEI. The first difference (times 400 to achieve approximate annual growth rates) is the inflation rate measure, $Infl_t \equiv 400 * \Delta LPY_t$.
LW	hourly wage (log of)	Total compensation taken from the quarterly national accounts of the OECD (QNA) and then divided by the labor volume, see below.
LVol	labor input volume of employed workers (hours, log of)	See the appendix.

Table 4: ADF unit root tests

variable	deterministics	lags (Δ)	sample	ADF stat
UR	const	4	1980q1-1999q3	-2.41
LEMP	trend, const	5	1980q1-1999q3	-2.55
LY	trend, const	7	1980q1-1999q3	-2.26
LY	trend, const	3	1980q1-1999q3	-2.36
Infl	const	5	1980q2-1999q3	-1.53
LWReal	trend, const	2	1982q1-1999q3	-2.67
LWReal	trend, const	1	1982q1-1999q3	-2.26
LVol	trend, const	5	1981q3-1999q3	-3.37
LVol	trend, const	7	1982q1-1999q3	-3.89^{*}
LVol	trend, const	5	1982q1-1999q3	-3.65^{*}

Remarks: *Tests are for the 5% and 1% level, an asterisk denotes significance at the 5% level. Lags refer to lagged differences. Several results per variable are reported if the lag specification is ambiguous.*

Table 5: Diagnostic tests for the levels VAR with three lags

ΔUR_t	$\Delta LEMP_t$	ΔLY_t	$\Delta Infl_t$	$\Delta LWReal_t$	$\Delta LVol_t$
absence of autocorrelation (1-5), F(5,47)					
.23	.82	.44	.28	.81	.34
normality, $\chi^2(2)$					
.42	.16	.63	.25	.41	.22
absence of ARCH (1-4), F(4,30)					
.49	.83	.61	.42	.54	.14

Remarks: *Sample 1981q2-1999q3. The levels VAR with $K = 3$ lags is unrestricted in terms of cointegration. Deterministics: linear trend, constant, an impulse dummy for 1984q1 and its difference (see text). Numbers are marginal significance levels (p-values).*

Table 6: Johansen cointegration test

eigenvalue	$H_0 : r \leq$	trace stat.	nominal p-value	Bartlett corrected trace stat.
	0	209**	0.000	
0.65	1	132**	0.000	
0.51	2	79.4**	0.001	63.0*
0.30	3	52.7**	0.003	42.5*
0.26	4	30.9**	0.009	24.8
0.20	5	14.1*	0.026	12.5*
0.17	6			

Remarks: *Sample 1981q2-1999q3, 3 lags (i.e. 2 lagged differences), restricted trend and restricted impulse dummy for 1984q1 (see text). Significance denoted by * (5%) and ** (1%). For the Bartlett correction see the text.*