# Did worksharing work in France? Evidence from a structural cointegrated VAR model

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**Abstract**

French employment increased significantly after a labor-market reform in 2000. This paper analyzes whether that development was driven by worksharing (the mandated reduction of the workweek length) as claimed by the government. We use a structural VAR model in error correction form (SVECM) to assess the impact of shocks to the workweek length. It turns out that downward workweek shocks actually had adverse employment effects. We conclude that other reform components were responsible for the employment success in France, namely reduced non-wage labor costs and possibly higher firm-level flexibility of temporarily adjusting the workweek.

**Keywords:** worksharing, structural vector error correction model, employment

**JEL codes:** C32 (multi-equ. time series methods), E24 (macro employment & wages), E65 (particular policy episodes), J23 (employment determination)

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# 1 Introduction

In many European countries unemployment has been high and persistent for decades, indicating the need for appropriate labor-market reforms. Worksharing is a reform that is often proposed in the political arena, meaning an intended redistribution of the total labor volume in the economy by cutting average work-hours per worker. One of the latest examples of this idea was the French reform package "Aubry II" (named after the minister of labor) which had been passed to become effective in 2000 and which was mainly promoted as a worksharing reform. For the period around the year 2000 figure 1 indeed shows the joint occurrence of a marked reduction of the workweek length and of employment growth rates that were high by historical standards,[1] and Logeay and Schreiber (2006) showed that the reform impact was significant. However, other important ingredients of the reform package were a reduction of non-wage labor costs (mainly through lower social security contributions for employers), and increased short-term discretion for firms with respect to the allocation of labor inputs over a calendar year. So while officially the reform consisted of the introduction of the 35-hour workweek, in reality it was the introduction of the 1600-hour working year with accompanying subsidies.[2] Obviously, it is important (e.g., for policy makers in other economies) to identify which part of the reform was actually responsible for its success.

Therefore the aim of this contribution is to assess whether worksharing really worked in France, or if its *prima-facie* success should rather be credited on other components of the reform package, like reduced social security contributions or the increased flexibility of labor allocation that was granted to firms. To answer this question we analyze the data before the reform in a structural vector error correction model (SVECM, i.e., a structural vector-autoregressive model with an explicit treatment of the cointegration properties of the data), and we derive the effect of a shock to the length of the workweek on employment

[1] The other downward spike in the workweek growth series was due to an earlier worksharing reform in 1982.

[2] For a more detailed institutional description see Logeay and Schreiber (2006) and references therein, e.g., Pham (2002).

by using impulse response analysis. We then find that structural shocks which reduced the workweek typically lowered employment instead of raising it. After a discussion of the potential pitfalls of using shock-based evidence for this particular policy evaluation, we still conclude that the reduction of the workweek length cannot be reasonably regarded as the cause of the employment boom after 2000. Instead we point to the reduced non-wage labor costs and the increased flexibility of labor allocation over the year as having lowered the shadow cost of labor and thus as the driving force behind the success of the French labor market reform.

This study provides empirical evidence against worksharing as a policy option on its own, which is interesting because many earlier results were ambiguous. In the theoretical literature see for example the seminal contribution by Calmfors and Hoel (1988), or more recently Marimón and Zilibotti (2000) where the possible employment creation would lower firms' profits. However, in the matching models by Rocheteau (2002) and by Ortega (2003) a possible Pareto improvement through lower standard hours is found, because it may offer a firm more flexibility to react to demand shocks. Empirical research has mainly backed economists' scepticism about worksharing, see for example the inconclusive results of Hunt (1999) or the negative findings by Crépon and Kramarz (2002). But it should be noted that the reforms analyzed by them increased wage costs markedly to achieve roughly unchanged income for incumbent workers, whereas the French reform package of 2000 operated without such labor cost hikes, which makes it especially interesting.

We proceed by discussing the (S)VECM framework and the reduced-form estimates in the next section. Then, as a preliminary exercise and drawing on Logeay and Schreiber (2006) we test and verify the effectiveness of the reform formally in section 3. In section 4 we discuss the identification assumptions of our baseline model and present the identified effects of workweek shocks, along with some alternative specifications. We consider it helpful to include a relatively detailed methodological discussion in section 5, and section 6 concludes.

Figure 1: French employment and length of the workweek


**Note:** Quarterly growth rates in decimals. See the data appendix for precise definitions of the series.

## 2 The SVECM specification

Let us first recapitulate the structural vector error correction model (SVECM) as the statistical framework for our analysis; a useful survey can be found for example in Breitung et al. (2004). The system of equations can be written as

$$\Delta \mathbf{y}_t = \alpha \begin{bmatrix} \beta' & \beta^{*'} \end{bmatrix} \begin{bmatrix} \mathbf{y}_{t-1} \\ i(\tau)_{t-1} \end{bmatrix} + \sum_{k=1}^{K-1} \Gamma_k \Delta \mathbf{y}_{t-k} + \delta \Delta i(\tau)_t + \mu + B\mathbf{u}_t, \tag{1}$$

where $\mathbf{y}_t$ is the $n$-dimensional vector of $I(1)$ variables,[3] and $i(\tau)_t$ denotes an impulse dummy $(\{...,0,1,0,0,...\})$ which will turn out to be necessary for date $\tau = 82q1$, and which is restricted to the cointegration relations with coefficients $\beta^*$. The corresponding differenced dummy $(\{...,0,1,-1,0,...\})$ is added unrestrictedly with coefficients $\delta$; note

[3] We have conducted standard unit root pretests on all variables with the result that $I(1)$-ness cannot be rejected in favor of (trend) stationarity. (And we could reject $I(2)$-ness in favor of $I(1)$.) Given that these results seem unsurprising for the variables described below, we chose not to report the details to save space.

that it does not induce mean shifts or similar breaks in the system. The vector $\mu$ is a constant, and $\mathbf{u}_t$ is an $n$-dimensional vector of white noise structural shocks with a contemporaneous identity covariance matrix, $E(\mathbf{u}_t\mathbf{u}_t') = I_n$. These structural innovations are mapped into the reduced-form residuals $\mathbf{e}_t = B\mathbf{u}_t$ by the non-singular $n \times n$-matrix $B$, leading to reduced-form residuals with covariance matrix $E(\mathbf{e}_t\mathbf{e}_t') = \Omega$. Thus $B$ must satisfy $BB' = \Omega$, which imposes $n(n+1)/2$ restrictions on $B$, and the remainder of $n(n-1)/2$ restrictions must be chosen by the researcher to identify the $n^2$ elements of $B$.

The $n \times r$-matrices $\alpha$ and $\beta$ with $r \leq n$, but with full column rank, reflect the possible cointegration between the variables. While $\beta$ holds the cointegration coefficients in its columns in the sense that $\beta'\mathbf{y_t}$ are the stationary deviations from equilibrium, $\alpha$ contains the loading coefficients which determine the adjustment of the endogenous variables induced by those equilibrium deviations. Given that the constant vector $\mu$ is unrestricted, it will serve as an intercept in the cointegrating relations as well as cumulate to linear deterministic trends in the individual variables.

The long-run impact of the reduced-form innovations $\mathbf{e}_t$ is given by (see, e.g., Johansen 1995)

$$C = \beta_\perp \left( \alpha_\perp' \left( I - \sum_{k=1}^{K-1} \Gamma_k \right) \beta_\perp \right)^{-1} \alpha_\perp', \tag{2}$$

where the orthogonal complements $\beta_\perp, \alpha_\perp$ are full column-rank matrices of dimension $n \times (n-r)$ and satisfy $\alpha_\perp' \alpha = 0$, $\beta_\perp' \beta = 0$. The long-run impact of the structural shocks $\mathbf{u}_t$ directly follows as $CB$. These long-run impact matrices $C$ and $CB$ have rank $n-r$, and for the interpretation of the structural shocks it is natural and common practice to assume that there are $n-r$ structural shocks with permanent effects and $r$ purely transitory shocks (see, e.g., King et al. 1991), otherwise the long-run behavior of the system would not really be identified. This means that $CB$ has $r$ columns of zeros.

In $\mathbf{y}_t$ we include the variables that are most relevant for our analysis, namely log employment ('EMP', full-time equivalent jobs), log output ('Y'), log hourly labor costs ('W'), and finally the log of the average length of the workweek ('workweek'). Output and labor costs are in real terms, deflated by the GDP deflator; see the data appendix for

Table 1: Diagnostic tests

| equation | no autocorr.(1-5) | normality | no ARCH(1-4) | homoscedasticity |
|---|---|---|---|---|
| EMP | 1.2237 [0.3100] | 9.3747 [0.0092] | 0.23442 [0.9178] | 1.3469 [0.2035] |
| Y | 0.57636 [0.7178] | 0.47319 [0.7893] | 0.98260 [0.4248] | 0.68277 [0.8421] |
| W | 0.79453 [0.5581] | 15.112 [0.0005] | 0.84737 [0.6658] | 0.84737 [0.6658] |
| workweek | 1.7002 [0.1493] | 5.6281 [0.0600] | 1.0208 [0.4706] | 1.0208 [0.4706] |

further details. Note that measures such as total hours worked and productivity per person or per hour are linear combinations of 'EMP', 'Y', and 'workweek', and thus implicitly included. For our baseline specification we refrain from adding other variables such as inflation or unemployment because the number of necessary identifying restrictions would become uncomfortably large. However, we will also present some sensitivity analysis related to these variables. But the baseline dimension of the system is $n = 4$, and the variables are ordered as follows:

$$\mathbf{y}_t' = (EMP_t, Y_t, W_t, workweek_t) \tag{3}$$

The available sample prior to the reform is 1980q1-1999q4, and in order to achieve satisfactory residual properties we must model the earlier workweek reform in (the first quarter of) 1982 by including a corresponding dummy variable $i(82q1)_t$. Diagnostic testing also demands that we choose $K = 3$ for the lag length, which is slightly higher than the recommendation of two lags by the usual information criteria. The resulting diagnostics are given in table 1 and show that the model is an adequate description of reality; the documented deviations from the normal distribution of the residuals are not problematic for our purposes.

The next crucial property of the model is the cointegration rank or the number of common stochastic trends, respectively. We apply the standard Johansen procedure as well as the less known test by Saikkonen and Lütkepohl (2000), which adjusts the deterministic terms under the respective null hypothesis. The results in table 2 clearly show that the rank is $r = 2$ at the 5% level, and thus the number of stochastic trends is also two.

Table 2: Cointegration tests

| $r=$ | Johansen p-value | Saikkonen and Lütkepohl (2000) p-value |
|---|---|---|
| 0 | 0 | 0.002 |
| 1 | 0.006 | 0.002 |
| 2 | 0.279 | 0.124 |

***Notes:*** The model setup is given in (1), i.e., the linear trend is excluded from the cointegration space, but allowed in the data. Note that in this model $r=4$ would imply an exclusion of the linear trend from the data, which is not sensible; therefore a test of $H_0: r=3$ vs. $H_1: r=4$ is not meaningful. The lag length is $K=3$ and the effective sample is thus 1980q4-1999q4, T=77.

We identify the cointegration (a.k.a. long-run equilibrium) coefficients $\beta$ in the following way to yield a labor cost setting relation and a demand relation for total hours worked (standard errors in parentheses):

$$W_t = \underset{(0.035)}{0.510}Q_t - \underset{(0.135)}{1.47}\, workweek_t - \underset{(0.027)}{0.230}i(82q1)_t \tag{4}$$

$$VOL_t = \underset{(0.044)}{0.546}Y_t - \underset{(0.074)}{0.677}W_t - \underset{(0.025)}{0.243}i(82q1)_t, \tag{5}$$

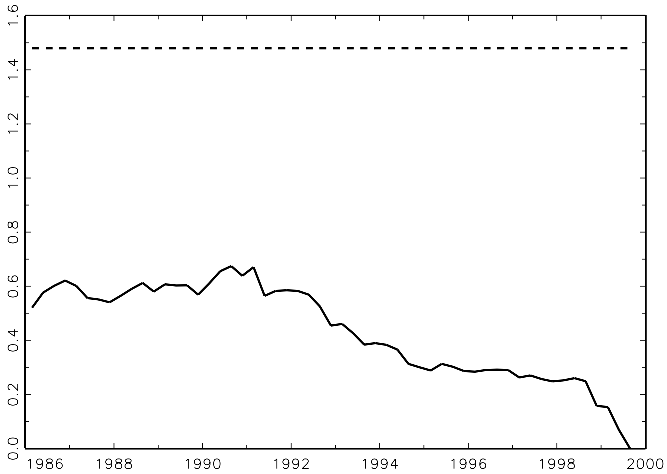
where $Q \equiv Y - EMP$ represents log productivity per capita and $VOL \equiv EMP + workweek$ is log total hours worked. The estimated coefficients have expected signs and plausible magnitudes; for example the elasticity of labor costs with respect to productivity is roughly one half, and the own-price elasticity of demand for hours worked is about minus two thirds. Note that the above restriction scheme is just-identifying for $\beta$ and thus does not affect other estimates such as $\widehat{\alpha}$.

After sequentially eliminating insignificant elements from $\widehat{\alpha}$, the resulting estimate of the loading coefficients corresponding to the identification scheme in (4) and (5) is given by

$$\widehat{\alpha}' = \begin{bmatrix} 0 & 0 & \underset{(0.139)}{-0.694} & \underset{(0.006)}{-0.071} \\ \underset{(0.007)}{-0.017} & 0 & \underset{(0.132)}{0.746} & 0 \end{bmatrix},$$

which produces an overall restriction likelihood ratio (LR) test result of 1.41 with p-value

Figure 2: Stability test




**Note:** Null hypothesis: all eigenvalues are stable. Recursive test statistics relative to the 5% critical value (dashed line).

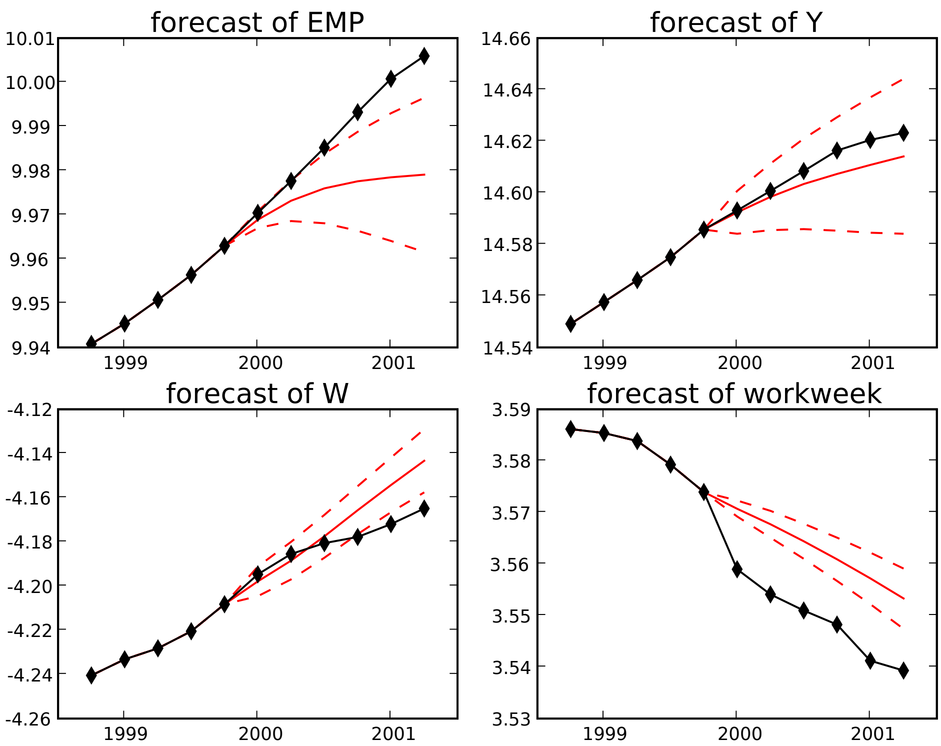
= 0.8418 ($\chi^2_4$). Hence we see that output $Y$ is found to be weakly exogenous and therefore its innovations have permanent effects. Figure 2 shows that the estimated model is stable over the given sample.

# 3 Effectiveness of the reform

In this section we contrast the forecast of employment with its actual development to assess whether the 2000 labor market reform had any significant effect. We test this by using the estimated system in reduced form to forecast the variables, calculate forecast confidence bands, and check whether the observed developments are covered by the forecast uncertainty.[4] Figure 3 displays this information. It can be seen that output is forecast surprisingly well, and of course that the actual workweek path is significantly lower than the forecast. More importantly, however, employment growth is significantly higher than expected, while real labor costs fall short of their forecast from the end of 2000 onwards.

[4]See Logeay and Schreiber (2006) for small-sample cointegration and extensive stability analysis of an empirical model that also includes the unemployment and inflation rates. The inflation forecast turned out not to be very informative, and the unemployment development essentially mirrors that of employment. Therefore the smaller 4-variable model in the present paper captures all important aspects of reality.

Figure 3: Forecast graph




**Notes:** Out-of-sample dynamic forecasts calculated from the reduced-form cointegrated VAR/VECM, based on the sample up to 1999q4. 95% confidence bands.

Given the quantified forecast uncertainty the discrepancy cannot be dismissed as a random fluctuation, and we conclude that the further analysis in this paper –which aims to identify the cause of this discrepancy– is the logical and necessary next step.

In the rest of the paper we focus on the effects of the workweek reduction because other potential explanations for the increase of employment do not apply. For example, the data used in Bassanini and Duval (2006) indicate that unemployment benefit replacement rates have increased quite a bit in France around 2000, instead of falling. Furthermore, national accounts show that expansionary fiscal policy did not occur, and the coverage of active labor market policy actually declined somewhat during the forecast period (Boulard and Lerais 2002). Finally, we repeat that our model appears to capture general demand conditions quite well in terms of the output forecast, such that normal business cycle fluctuations also cannot account for the extraordinary employment developments.

# 4 The effect of workweek shocks

## 4.1 Baseline identification

To identify the effect of a shock to the workweek length we need to impose sufficient restrictions on the model. First of all we assume that there are only two shocks that have permanent effects. This is a standard assumption at least since King et al. (1991), because it allows to interpret the permanent shocks as the sources of the stochastic trends in the system. Since the second variable $Y$ is weakly exogenous, it is clear that the second shock must be permanent. We therefore order the shocks such that the first two are permanent and the last two are transitory, which means that we restrict the last two columns of $CB$ to be zero:

$$CB = \begin{bmatrix} \bullet & \bullet & 0 & 0 \\ \bullet & \bullet & 0 & 0 \\ \bullet & \bullet & 0 & 0 \\ \bullet & \bullet & 0 & 0 \end{bmatrix}, \tag{6}$$

where $\bullet$ denotes an unrestricted element. However, because of the reduced rank of $CB$, this scheme provides only four linearly independent restrictions, and since $4(4-1)/2 = 6$ restrictions are needed to identify all shocks, we still need to restrict two elements of $B$ to be zero. In our case these two additional restrictions must be evenly distributed among the permanent and the transitory shocks, such that one restriction belongs in the first two columns of $B$, and the other in the last two columns.

We loosely label the two permanent shocks as a labor force shock and a technology shock, where the labor force shock might comprise a wide range of fluctuations in deeper variables such as preferences, education, fertility, and others. Therefore it seems reasonable that this first permanent shock could affect everything and we leave its impact unrestricted, leading to an unrestricted first column of $B$. The second permanent shock stemming from outside the labor market should have a negligible immediate effect on employment due to sluggish labor market adjustment, and possibly may not affect hourly labor costs directly, either. This last restriction would already be over-identifying the permanent shocks and is thus testable. The second column of $B$ is therefore specified as $[0, \bullet, 0, \bullet]'$.

The transitory shocks may be interpreted as equilibrium deviations in terms of the cointegration relations. In our case this means that we have a wage setting shock and a shock to the relation for total hours worked. The wage setting shock should not have a noticeable immediate effect on $EMP$, again because of labor market sluggishness, implying that the third column of $B$ is $[0, \bullet, \bullet, \bullet]'$. This assumption already identifies the transitory shocks, and thus the remaining total-hours shock may be allowed to affect everything. Altogether this yields the following pattern for the $B$-matrix:

$$B = \begin{bmatrix} \bullet & 0 & 0 & \bullet \\ \bullet & \bullet & \bullet & \bullet \\ \bullet & 0 & \bullet & \bullet \\ \bullet & \bullet & \bullet & \bullet \end{bmatrix} \tag{7}$$

Altogether, this SVECM model with short- and long-run restrictions produces a clearly acceptable (over-) identification test result: 0.3834, p-value = 0.5358 ($\chi^2_1$).

As a workweek shock constitutes a deviation from the total-hours relation, it should be noted that in our baseline model it does not have permanent effects. This property may at first seem inappropriate given the political aim to reduce unemployment permanently, but we argue that it is perfectly reasonable: First of all, the estimated responses to workweek shocks are actually quite long-lived as will become clear below, being significantly different from zero for about three to five years. Compared to the forecast horizon used in figure 3 this is quite long already. Also, controversial reforms such as the workweek reduction are often not perceived as permanent by the general public. Indeed, the party in power changed after the following elections and partially reversed the worksharing reform. In general, in a political world without binding commitment it is rational for agents not to treat every reform package as infinitely lived. Finally, the workweek shows a long-run downward trend in industrial economies, especially in Europe, driven by other factors such as income effects. This implies that an initially binding workweek restriction will become redundant (for the majority of workers) after some time and therefore will have only transitory effects even if the restriction itself is permanent.

We now turn to the responses to a shock on hours worked in figure 4, which represents the main empirical results of the present paper. The employment response in the first panel has the "wrong" sign, i.e., a reduction of the workweek would tend to lower employment instead of raising it! The second panel shows the response of output which has a shape that is similar to the employment response, as expected. In the third panel we see that positive (negative) workweek innovations imply lower (higher) hourly labor costs, which is a reasonable finding. We conclude that in our model for France a workweek reduction tends to depress employment and output, while raising hourly labor costs. Therefore worksharing in its narrow sense would not at all seem to be responsible for the employment increase that we observed in France.

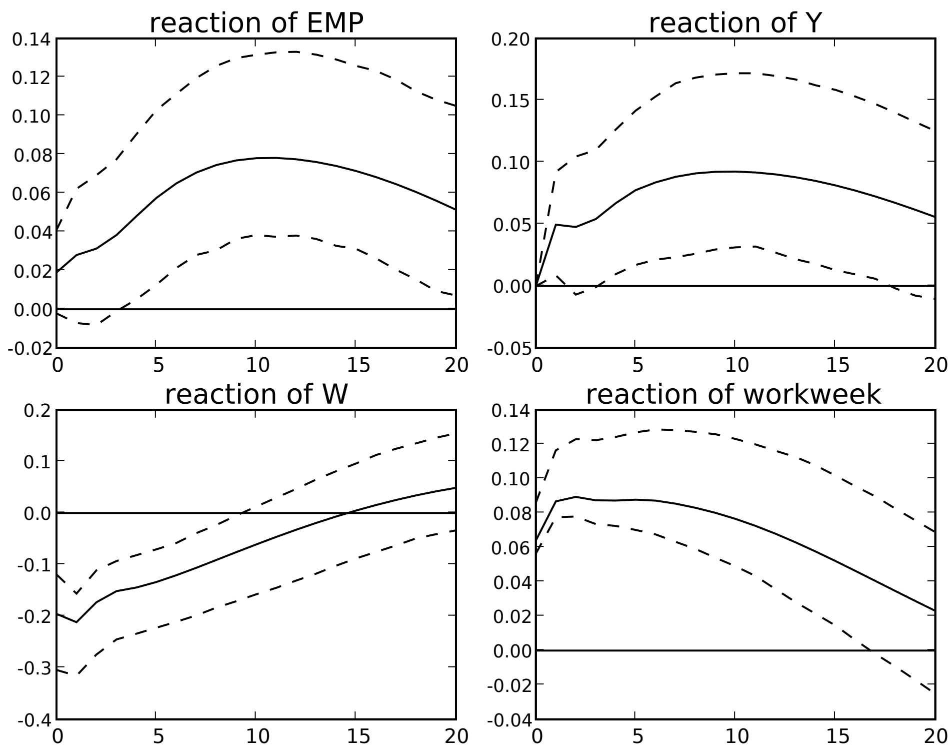
## 4.2 Variations and sensitivity analysis

Naturally the impulse responses derived in the previous subsection depend on the assumptions that were imposed in the specification of the model. Therefore we present several different specifications in order to check the robustness of the results.

First of all we investigate whether imposing the zero restrictions on $\widehat{\alpha}$, and especially weak exogeneity of $Y$, makes a difference. Figure 5 reports the results, where the same estimates of $\widehat{\beta}$ and the same identifying restrictions on $CB$ and $B$ have been used as before, now leading to an overidentification test result of 0.94 with p-value = 0.3328 ($\chi^2_1$). The impulse responses have roughly the same shape as before, but the estimation uncertainty is greater in this variant.

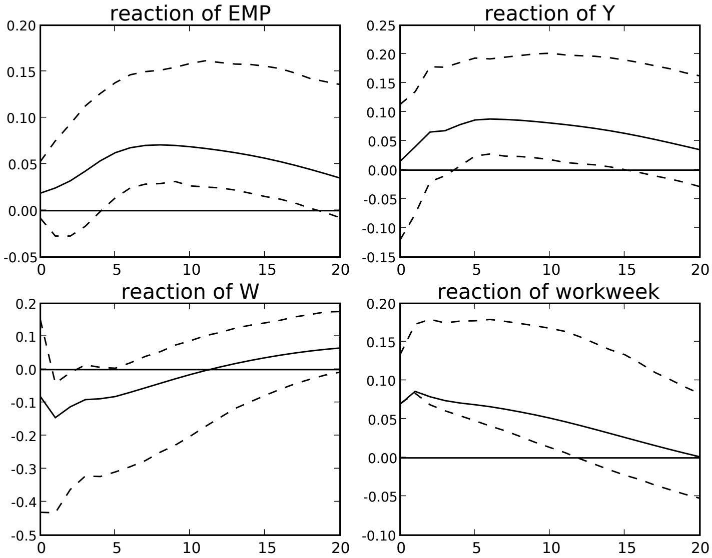
Next, we check the results of an approach without any long-run restrictions (on $CB$).

Figure 4: Responses to workweek shocks, baseline specification




**Notes:** Impulse responses –rescaled times 100– with respect to a (positive) shock to the workweek length from the SVECM with (over-) identifying restrictions (6) and (7). The confidence bands have 95% coverage and are computed with the Hall bootstrap, 1000 replications.

Figure 5: Responses to workweek shocks, loading coefficients $\widehat{\alpha}$ unrestricted

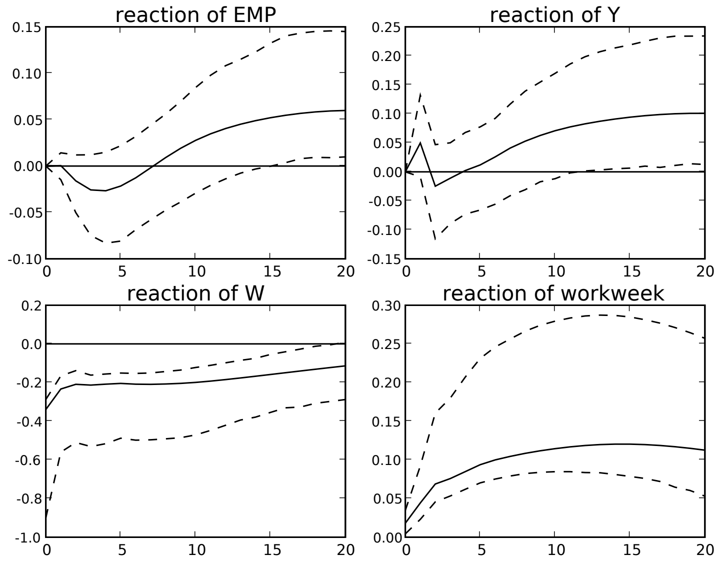



Therefore all shocks are now allowed to have permanent effects, and therefore none can be interpreted in terms of the stochastic trends or relating to the cointegration vectors anymore. Also, more restrictions in $B$ are needed then, and we specify the following restriction scheme for $B$ which embeds the previous restrictions:

$$B = \begin{bmatrix} \bullet & 0 & 0 & 0 \\ \bullet & \bullet & \bullet & 0 \\ 0 & 0 & \bullet & \bullet \\ \bullet & \bullet & 0 & \bullet \end{bmatrix}$$

The results of this identification scheme (LR test 0.5815 with p-value = 0.4457, $\chi^2_1$) are shown in figure 6. The response of labor costs is roughly the same, as is the sign of the response of the workweek itself, however not the shape. The most important difference of course concerns the response of employment up to a horizon of seven quarters; although it is clearly insignificant, the sign of the point estimate is now reversed. This result still would not really seem to support worksharing as a policy option, given that the longer-run effects (after seven quarters) are of the same "problematic" sign as in the baseline model.

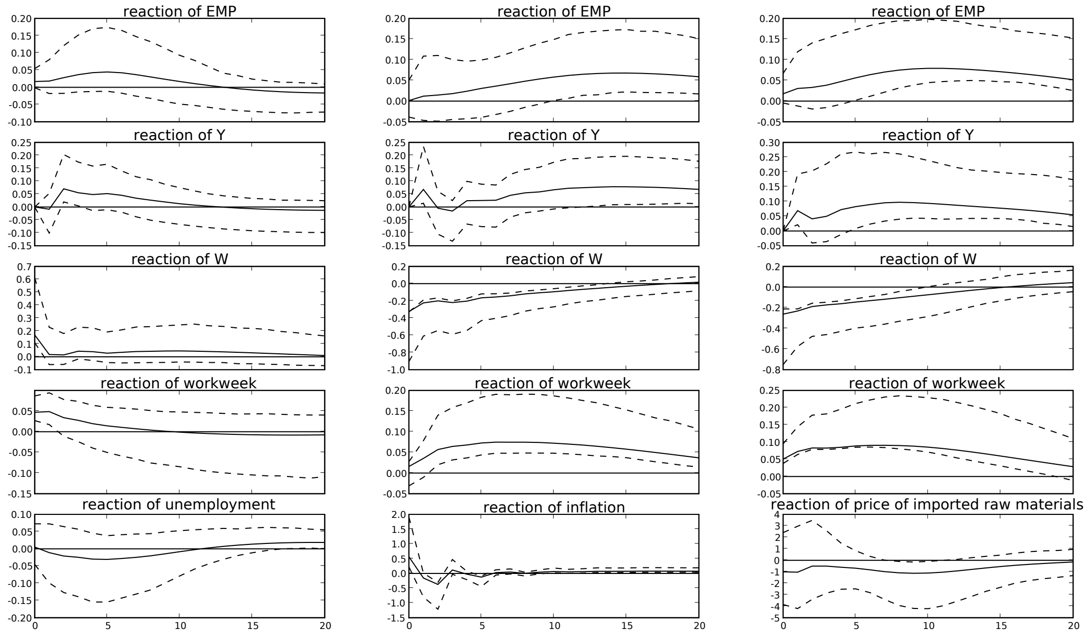
Figure 6: Responses to workweek shocks, without long-run restrictions




Finally, we address the possibility that some relevant variables may have been omitted, namely the unemployment rate (decimal between 0 and 1), the inflation rate (annualized, times four), and the price of raw material imports (log). A problem of analyzing larger systems is of course that the number of needed identifying restrictions grows exponentially, and therefore we add only one of the potentially interesting variables at a time in a sequence of new systems. The weak exogeneity of $Y$ still holds in all three models extended with unemployment, inflation, and the log raw material import price, respectively (p-values 0.164, 0.838, 0.514). We maintain two permanent shocks and therefore a cointegration rank of three in these extended models, where the last three shocks are restricted to be transitory, and $B$ embeds the previous restrictions and looks as follows:

$$B = \begin{bmatrix} \bullet & 0 & 0 & \bullet & \bullet \\ \bullet & \bullet & \bullet & \bullet & \bullet \\ \bullet & 0 & \bullet & \bullet & \bullet \\ \bullet & \bullet & \bullet & \bullet & 0 \\ \bullet & \bullet & 0 & \bullet & \bullet \end{bmatrix} \tag{8}$$

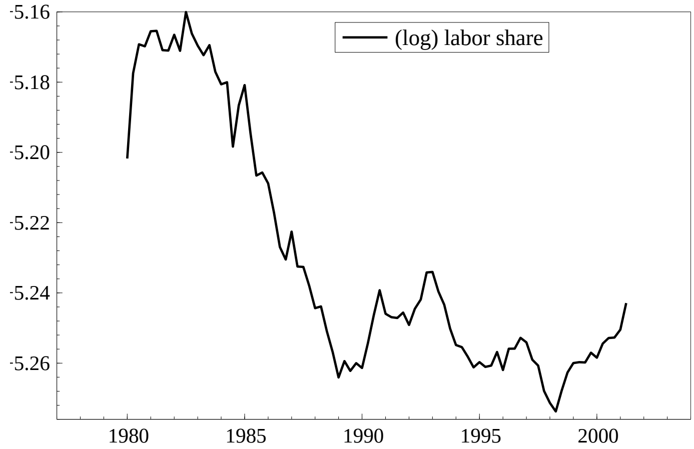
Figure 7: Responses to workweek shocks, in extended models




**Notes:** One column per extended system variant, with the respective additional variable appended at the last position.

The corresponding over-identification test results turn out as follows (p-values, $\chi^2_1$): with unemployment 0.276, with inflation 0.940, with log raw material import prices 0.634. Figure 7 shows that the responses to workweek shocks are not much changed in the models with inflation and import prices, apart from some differences concerning the confidence bands. In the model with the unemployment rate results are less similar, but are qualitatively mostly congruent with the baseline model, with the notable exception of the response of labor costs which is close to zero here. All in all, the extended five-dimensional models broadly confirm the sign of the employment response, but they also already present the well-known problems of finding plausible identifying restrictions in higher-dimensional models.

Figure 8: Non-stationary labor share in France




# 5 Discussion of the chosen method

It seems worthwhile to clarify why we did not choose alternative approaches. For example, in principle it would be desirable to use detailed micro data as demonstrated in Crépon and Kramarz (2002) for the 1982 reform. However, existing micro data is limited to samples of firms with specific characteristics.[5] These samples are not representative, and therefore a macro analysis remains useful.

Another seemingly natural approach would be to directly estimate the elasticity of labor demand with respect to the workweek length. However, the structural labor demand function in general depends on non-observable technological progress; this dependency cancels out only in the special case of a unit elasticity of factor substitution, i.e., with a Cobb-Douglas production function. Note that in this sense neither of the estimated cointegration relations constitute a structural labor demand equation. But since the Cobb-Douglas technology implies stationary factor shares, the obvious non-stationarity of the French labor share documented in figure 8 refutes this assumption. The consequence is that a direct estimate of the relevant elasticity is impossible.

Finally, the most important methodological issue is whether the shock-based evidence of impulse response analysis can be used to perform an evaluation of a policy shift. A

[5]Passeron (2002) for example only analyzes firms that took part in the preceding Aubry I scheme. Another study for selected firms is Bunel (2002), who uses the special "Passages" data set with firms that reduced their *effective* weekly working time to 35 hours (and thus excluding firms with overtime).

valid objection against our approach could in principle be due to threshold effects that stem from the existence of restructuring costs: Even given adverse effects of *temporary* workweek shocks, after *permanent* shifts firms may instead find it optimal to adapt to the new regime by bearing the one-time reorganization costs (for example by installing extra desks) and possibly increasing employment. We agree that this argument is theoretically convincing, but it again boils down to the issue of transitory versus permanent shocks which we discussed already in the context of our baseline shock identification scheme (see section 4.1). Viewed from this angle, the impulse response analysis seems adequate for our specific case, and the responses match the post-reform behavior quite well.

# 6 Conclusions

Given that the French labor market reform package at the turn of the millennium raised employment significantly (as seen in section 3), the combined measures of reducing the workweek length, lowering non-wage labor costs, and offering more intertemporal labor allocation flexibility to employers must be viewed as quite effective. In the present paper we have addressed the issue of whether the worksharing component of the reform package can be held responsible for the employment success.

Our results showed that the reduction of the workweek length is unlikely to have done the job, because the estimates imply employment *reductions* in response to workweek shortenings. This adverse effect of workweek reductions together with the additional and expected finding that they boost hourly labor costs instead suggests the following conclusions: First, the additional reform ingredient of reducing employers' social security contributions was effective and significantly increased labor demand. Secondly, the mandated reduction of the workweek alone would have depressed employment at least temporarily. It also counteracted the reduction of non-wage labor cost components such that the fall in hourly labor costs did not become significant until the end of 2000.

Finally, the increased flexibility of inter-temporal labor allocation that was granted to employers should have lowered the effective (shadow) cost of labor in view of short-term demand fluctuations at the firm level, and thus may also have helped to sustain labor demand. However, the extent of this last effect is quite uncertain because our approach does not provide direct evidence on the consequences of higher flexibility, and also because the literature on this issue appears to be inconclusive.[6]

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[6] See Wolf and Beblo (2004) for evidence of nonlinearity and for further references.

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# A Data appendix

(This description is adapted from Logeay and Schreiber 2006.) DARES (the statistical department of the French labor ministry) publishes the average working time of full-time employees in each quarter on the basis of the survey on labor activity and employment status (ACEMO) 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.)

INSEE (the French statistical office) 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 this 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. A comparison of various workweek measures is given in table 3.

GDP and the GDP deflator are taken from OECD Main Economic Indicators.

Total compensation for labor is taken from OECD Quarterly National Accounts, and real hourly labor costs (log of, $W$) are derived by using the GDP deflator and the measure of labor volume described above.

Table 3: Different measures of the average workweek

| | 1999q1 | 1999q2 | 1999q3 | 1999q4 | 2000q1 | 2000q2 | 2000q3 | 2000q4 |
|---|---|---|---|---|---|---|---|---|
| ACEMO | 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 | 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 | 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* |

**Notes:** The numbers in *italics* are quarterly growth rates in %.

ACEMO: Survey data published in the DARES database (as of November 2003); refers to firms with more than 10 employees and full-time employees.

DARES: Data calculated by DARES and also published in their database (as of November 2003), correcting the results of ACEMO for firms with less than 10 employees and part-time employees. It additionally corrects a statistical break of the definition of the working-time in 2000 induced by the Aubry II reform. Note that this series could not be used in our analysis because it only dates back to 1993.

Our data: See the text of the appendix for the exact calculation.