Journal of Applied Economics, Vol. VI, No. 1 (May 2003), 127-156

WHY IS FRENCH EQUILIBRIUM UNEMPLOYMENT SO HIGH? AN ESTIMATION OF THE WS-PS MODEL

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Submitted July 2000; accepted April 2002

Unemployment in France rose steadily from the early-seventies to the mid-eighties. Since the mid-eighties it has continued to experience fluctuations around a very high average level. Equilibrium unemployment theories are a useful framework within which to account for these developments. A multivariate estimation of the WS-PS model on macroeconomic quarterly data, which includes a larger number of potential unemployment determinants than earlier work, allows an enriched reading of the rise in French unemployment and of its persistence at a high level. We estimated it using a conditional VAR-ECM model, which is based upon the weak exogeneity properties of variables over the 1970-1/1996-4 period. The rise in equilibrium unemployment by 10 points in 25 years can essentially be explained by the rise in tax and social wedge, the slowdown in labour productivity and the deterioration of job security. Terms of exchange and skill mismatch account for only a slim part of the rise in equilibrium unemployment.

JEL classification codes: C32, E24
Key words: labour market, WS-PS model, equilibrium unemployment, co-integration, conditional VAR-ECM model

I. Introduction

In France, the unemployment rate has hovered at around 10% for over 20

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years. Whilst it has experienced significant fluctuations, it has always moved back to this average level. That is why special attention has been granted to structural unemployment theories and their empirical evaluation, in France, as well as in other European countries with comparable evolutions.

Equilibrium unemployment theoreticians commonly substitute a structural relation called WS for Wage Schedule (Lindbeck, 1993), for the labour supply from households in the traditional equilibrium of the labour market. The shape of this relation is deduced from theoretical models most often based on the microeconomic behaviours described by the new labour market theories (e.g. efficiency wages, bargaining models, insider/outsider approach). This relation intersects with another one describing structural price setting (PS). They jointly determine the equilibrium unemployment level that will be modified by structural shocks affecting the determinants of wage or price setting, notably oil crises, shocks on the level of direct or indirect taxes and real interest rate shocks. This sensitivity to structural shocks differentiates the approaches in terms of equilibrium unemployment, qualified as structuralism by Phelps (1994), from those in terms of natural unemployment, as defined by Friedman (1968). Moreover, it leads to a higher unemployment determinant set than the one usually considered by a Phillips' curve approach (Bean,1994). The theoretical WS-PS models have been popularised through the work of Layard, Nickell and Jackman (1991). They have now integrated employed worker heterogeneity (for an example, see Laffargue, 1995) and the dynamic aspects of wage setting (Manning, 1993; Cahuc and Zylberberg, 1998). This theoretical maturity has resulted in an impressive extension in the list of potential unemployment explanations, which both rest on an explicit microeconomic base and are connected to wage or price schedules in a general equilibrium framework.

This theoretical maturity contrasts with the state of empirical research whose purpose is to estimate the WS-PS model. The literature on this topic can be categorised into two separate groups. The univariate estimations of the WS and PS relations are compatible with a large number of unemployment equilibrium determinants, in accordance with the theory, but do not take the interdependences between variables into account. Inversely, too large a number of variables become incompatible in practice with a multivariate estimation of the WS and PS relations, yet it is more satisfactory to take the interdependences between wage and price setting into account. In French macroeconomic data, the equilibrium unemployment rise since the early seventies was thus entirely explained by real interest rate evolution, technical progress and the terms of exchange in Bonnet and Mahfouz (1996), by the evolution of the wage wedge, the replacement ratio and productivity in L'Horty and Sobczak (1997), and by the evolution of capital cost and the wage wedge in Cotis, Méary and Sobczak (1997). These multivariate estimations put the emphasis on the crucial role of some variables but do not fully explain the rise and persistence in unemployment.

The contributions of this paper are essentially threefold. To begin, it focuses on a structured theoretical setting that deals with a large number of potential equilibrium unemployment candidates. It then proceeds to build a set of original indicators for some of these determinants. Finally, it uses an econometric methodology to consider the effects of these variables simultaneously. This allows a greater understanding of the formation of equilibrium unemployment than that of the existing applied studies on French data. This reading is theoretically justified, compatible with the statistical properties of the variables considered, and validated by multivariate econometric techniques, which leads to a retrospective and quantitative explanation of French unemployment over the 1970/1-1996/4 period.

As to the econometric methodology, this paper gives an estimation of the WS-PS model on French macroeconomic data that is both in keeping with Johansen's multivariate estimation techniques and compatible with a large number of variables.[1] This re-estimation is made possible by taking the weak exogeneity properties of variables into account. The multivariate model can indeed be partitioned in two blocs whose parameters vary freely: a marginal model gathering the weakly exogenous variables for the long run parameters

[1] Our estimation is purely national and enables estimations obtained to be completed with multinational data using panel econometric techniques (cf. for example Layard, Nickell, Jackman, 1991 and Layard, Nickell, 2000). A comparative approach on international data imposes great restrictions in the construction of data that must be homogeneous between countries. In a purely national study, we do not have this constraint of data homogeneity, which allows us to construct more representative indicators of the French situation. This is, for example, the case for a complete set of SMIC hikes, replacement ratio, working hours, or progression of social wedge. These data would be either impossible to build for other countries or feebly representative of the French situation in an internationally standardised database.

of the Vector Error Correction Model (VAR-ECM), and a conditional model composed of other equations. Co-integrating vectors can then be estimated from only the conditional model, reducing the system size without losing any information from the full VAR-ECM.

Starting from a quarterly database composed of 16 series and covering the 1970/1-1996/4 period, we estimated the WS-PS model using an unrestricted VAR-ECM approach, composed of ten variables. Two co-integration relations were estimated from a partial system composed of seven equations and conditional to the three equations describing the evolution of weakly exogenous variables. These relations were identified using an approach inspired by Manning (1993), according to which productivity is not in the structural wage equation. It is important to note that the equilibrium unemployment estimation is robust with respect to that identification constraint. Finally, exclusion tests retained only five determinants in the progression of unemployment equilibrium in France: hourly productivity, through which real interest rates can have an impact; the internal terms of exchange, which essentially vary under the impact of oil crises and the exchange rate; the quit ratio; the aggregate wage wedge through which the different deduction rates can have an influence; and skill mismatch. The method used allows a calculation of the respective influences of these determinants and their retrospective contributions to unemployment development. On the other hand, the replacement ratio, which depends on the generosity of the unemployment benefit system, working hours, the French minimum wage (*SMIC*) increase and the progressiveness of the social wedge would have had a non-significant role in the evolution of equilibrium unemployment according to this estimation.

Section II provides a theoretical review of the WS-PS model. It presents the list of potential variables that can account for unemployment equilibrium, the mechanisms through which these variables have an influence and the data used in this study, which required that several original indicators be constructed for the different variables. Section III presents the model estimation results. Finally, Section IV presents our conclusions.

II. Equilibrium Unemployment Determinants and their Measures

Ideally, the richest possible theoretical model would stem from a microeconomic wage and price setting base in a dynamic framework that would take agent anticipation setting into account, as well as nominal and real rigidities and the impact of labour market institutions, such as systems of employment protection, trade union activity, active labour market policy, and so forth. Such a labour model would contain a heterogeneous factor, where all deductions and transfer systems would be modelled, including the modalities of unemployment benefit payments, their digressiveness in time and more generally, the degree of progress in the fiscal and social system. On that basis, one would deduce both a short and long term structural form of WS and PS in a general equilibrium framework to describe all determinants of equilibrium unemployment. Given all these enrichments, there is most probably no analytic solution as to the log-linearisation of structural wage and price curves. Moreover, the specification of log-non-linear structural expressions of these curves would be highly dependent on the whole successive modelling choices, and would make a non-linear estimation very delicate. In any case, writing such a full model seems impossible.

Consequently, the estimation strategy adopted here is less ambitious. From the theory, we have selected a list of variables, their expected signs, possibly some bounds for their elasticities and no more. We can then let data speak for themselves in a multivariate log-linear estimation framework.

A. Theoretical Variables

A first list of variables is given by a WS-PS model inspired by Layard, Nickell and Jackman (1991). In that model, goods markets are in imperfect competition and wages are the result of a negotiation between unions and employers, the latter maintaining their right to manage. This static homogenous labour factor model is what enables us to describe the traditional determinants of price and wage schedule and equilibrium unemployment.

In a formal definition of the value of unemployment equilibrium, one solves the system composed of the WS and PS structural equations by substituting the wage share in the added value when a Cobb-Douglas technology is used. One thus obtains a reduced form of the wage equation that defines the level of equilibrium unemployment. In the Layard, Nickell and Jackman (1991) model, this reduced form is presented as the structural form of WS. Equilibrium unemployment increases, ceteris paribus, with union power, the replacement ratio and employees' risk aversion. It decreases with the risk of becoming unemployed, with the degree of competition on the goods market, and with the labour factor efficiency parameter. It is also sensitive to the terms of exchange and to all the parameters characterising the tax system, which play a role in the wage wedge and modify the replacement ratio.

In the case of a CES production function, the structural wage equation remains the same, but it is no longer the case for the equilibrium unemployment expression, which in addition now has a productivity term whose impact depends on the substitution elasticity of factors. If factors are less substitutable than in the case of a Cobb-Douglas technology, the equilibrium unemployment elasticity to labour productivity in efficiency units is negative. An increase in productivity leads to both a wage increase and an unemployment decrease. If factors are more substitutable than in the case of a Cobb-Douglas, productivity in efficiency units has a positive impact on equilibrium unemployment. In other respects, technical progress can be seen to have no impact on equilibrium unemployment levels and to lead only to a real wage increase.

The interest rate influence goes through the productivity term. In the case of a Cobb-Douglas technology, an increase in interest rates reduces equilibrium capital intensity, decreases labour productivity, increases equilibrium labour costs and finally increases equilibrium unemployment. An increase in real interest rate always leads to a decrease in productivity, but it yields to a decrease in equilibrium unemployment if factors are more substitutable than in the case of a Cobb-Douglas technology, and to an increase in the opposite case (PS variations more than compensate those of WS in the former case). This result is not non-intuitive: when factors are slightly substitutable, a capital cost increase limits the use of all factors and thus increases equilibrium unemployment; when they are very substitutable, the substitution effect is bigger than the income effect and equilibrium employment increases.

This model can be completed by specification enrichments, which introduce new variables, by taking into account the dynamic aspects of wage and price schedules and by the introduction of labour heterogeneity. A first specification enrichment consists of introducing working hours. If hours and men are perfect substitutes concerning the technology used by firms, and if a reduction in working hours is not compensated by a rise in hourly wages, taking working hours into account would not change the PS expression. A reduction in working hours can also affect wage setting, according to the individual and union utility functions and the way this reduction is implemented (imposed or bargained). Another specification enrichment is in no longer assuming that the different deductions are flat. Then, if the progressiveness of social or fiscal deductions is taken into account, the price equation remains unchanged but wage equation is distorted, a stronger progressiveness having the same effect as a reduction of union market power in the bargaining. Moreover, in the Layard, Nickell and Jackman model (1991), a φ parameter is introduced to weight unemployment rates in the expression of the employed workers' withdrawal in the bargaining. This parameter represents the risk of becoming unemployed as a function of unemployment rate. Unemployment risk can also be measured in reference to the short length unemployment rate or to the quit ratio extracted from data flows on the labour market. This latter extension is also essential when the dynamic aspects of wage setting are taken into account. Finally, taking employed worker heterogeneity into account leads to other enrichments in the understanding of employment setting. If one distinguishes between different qualifications, one takes the consequences of the skill mismatch on the labour market into account.

All in all, the initial theoretical model and its enrichments lead the price and wage schedule to depend on apparent labour productivity or on the real interest rate, on the price-elasticity of demand, on the efficiency of the labour factor (which corresponds in a Cobb-Douglas production function to the share of wages in added value) and on working hours. As far as real wage setting is concerned, it depends on the unemployment rate, on union bargaining power, on the degree of competition in the goods market, on employed workers' risk aversion, on the replacement ratio, on the wage wedge and its components, on working hours, on wage wedge progressiveness, on the quit ratio and on the skill mismatch. Equilibrium unemployment depends on all these determinants as soon as their elasticities differ in the price and wage equations.

B. Indicators for those Variables

The empirical evaluation of equilibrium unemployment is faced with a data deficit. Some determinants of the WS-PS models are not directly observable and cannot therefore be found in any existing database. This is the case of price elasticity for goods demand, which embodies the degree of competition between offers on the product markets. It is also the case of the mark-up battle between employed workers' and employers' representatives in wage bargaining, of employed workers' risk aversion or of their psychological discount rate. Other theoretical determinants of equilibrium unemployment can be observed in a more or less direct way, but are not the subject of standardised statistic series (as is true in the case of replacement ratio or of wage wedge progressiveness, for instance). Given this data deficit problem, one answer is to build indicators for these variables. The asset of building indicators is to produce new statistics containing information on market labour evolution.

Most traditional data consists of gross wages, prices, added value, and rates of unemployment. We used the average gross hourly wage rate in the non-financial non-agricultural manufacturing sectors, which was extracted from quarterly accounts. This is also the case for consumption prices, and for added value prices and employment, which were all re-calculated for the non-financial non-agricultural manufacturing sectors. Two apparent labour productivity indicators were used: productivity per capita, which is the ratio of added value to employed workers, and hourly productivity, which is the ratio of per capita productivity to working hours.

Working hours are the synthetic indicator calculated by the French Ministry of Labour. It takes part-time job development into account, which has been promoted over the recent period by state specific assistance (a basic reduction of social wedge to share part-time jobs, some modalities of social contribution reduction on low wages that were encouraging part-time jobs). This indicator dropped throughout the nineties, falling more sharply after 1993, because of the accelerated diffusion of part-time jobs. This indicator is closer to the average working hours really performed by workers.

Real interest rate is the price of public and semi-public bonds. Its direct introduction into a price equation justifies itself when one considers the capital setting as endogenous and when one considers the existence of an asymmetry in capital and labour mobility. In the case of a small open economy in a perfectly integrated worldwide capital market, the interest rate is fixed from abroad and involves capital intensity and equilibrium productivity, which is decisive for price behaviour. An increase in interest rates reduces equilibrium capital intensity, which leads to a decrease in equilibrium labour costs and to a rise in unemployment (PS is horizontal and moves downwards).

The global wage wedge is composed of the internal terms of exchange, which are the ratio of consumption prices to producer prices, and of the social and fiscal wedge, which is itself composed of the social wedge (employers' and employees' contribution rates) and of the fiscal wedge (*VAT*, income tax rate). Employers' and employees' contribution rates (*CSE* and *CSS*) are extracted from social scales, applied to medium wage and given the same progression as the social security ceiling. Direct or indirect (Personal Income Tax and *VAT*) income tax rates, are taken from the databases of the French Ministry of Finances. Theoretically, only the deductions that are not considered by employed workers as benefits or postponed income compensations exert an upward pressure on labour cost and equilibrium unemployment.

For the replacement ratio, we used the indicator created by the Unédic (1997), which is an average of the situations of all unemployed workers at a given date. An extension of unemployment duration leads to a replacement rate reduction, which provides a satisfactory result. This quarterly indicator has been available since 1986. For previous years, we used the unemployment benefit scales applied to the situation of a medium unemployed worker whose period out of work is given by long series employment surveys (we also assumed a 6-12 month affiliation duration). Spontaneously, the two series were very close in 1986. The replacement ratio was clearly on the decrease after the 1992 reform of unemployment benefits.

To measure the quit ratio, which includes the risk of losing one's job and can be linked with the systems of labour protection, we used the transition rate between employment and unemployment, extracted from employment survey, and made it quarterly by a simple linear interpolation. It is important to notice that this rate is not directly connected to the unemployment rate: more intensive flows from employment to unemployment do not imply an increase of unemployment rate, since transitions from inactivity can decrease and exit employment rate can rise. Inversely, an employment flow reduction to unemployment does not imply an unemployment decrease, since these flows can be compensated by an increase of the transitions from inactivity to unemployment, or by a reduction of unemployment exits to employment or inactivity. This transition rate from employment to unemployment is an approximate measure of the probability of being laid off, which can vary in an inverse way to unemployment rate.

Employed workers' bargaining power is one of the parameters on which we have very little information. Instead of using a simple trend or a unionisation rate, whose reading is complex in the case of France, we have used the complete set of hikes given to the minimum wage (*SMIC*). It is an indirect proxy, whose justification is less to demonstrate the wage scale rigidity when the *SMIC* is increased, than to synthetically sum up the evolution of the general climate around wage setting.

The progressiveness of the wage wedge (*PROG*) is calculated here using the residual progressiveness indicator proposed by Jakobsson (1976). The progressiveness of the contributions of employers and employees are calculated separately and the aggregate indicator is obtained by summation.

The mismatch indicator (*MM*) is the semi-variance of relative employment rates by qualification, whose theoretical reading is given by Jackman, Layard and Savouri (1991): when wage curves are convex, a greater dispersal of unemployment rates induces an upward pressure on wages, which leads to a higher equilibrium unemployment rate. Sneessens's indicator (1994) is also tested. It deals with the ratio of the share of qualified employed workers in employment to their share in the labour force.

Other institutional variables could be taken into account when dealing with international approaches using panel data estimation techniques. Thus, centralism of wage bargaining, the systems of labour protection (for the part that does not affect the quit ratio) and active labour market policy can influence wages and unemployment formation. Without any time series data available for these variables, these determinants will be included in our econometric estimation by the constant, or, if they have varied across time, by the trend of our wage and price equations.

III. WS-PS Model Estimation

This section describes the statistical properties of the series as well as the results of the unrestricted VAR-ECM modelling that we finally adopted.

A. Univariate Properties of the Series

The database is composed of 15 quarterly series. It concerns the non-agricultural manufacturing sector and covers the 1970-1 to 1996-4 period. Deduction rates can be regrouped in two levels of aggregation, adding four indicators more.

The first step in the analysis was simply to look at the data univariate properties and to determine the degree to which they were integrated. Theoretically, a process is either I(0), I(1) or I(2). Nevertheless, in practice, many variables or variable combinations are borderline cases, so that distinguishing between a strongly autoregressive I(0) or I(1) process (interest rates are a typical example), or between a strongly autoregressive I(1) or I(2) process (nominal prices are a typical example) is far from easy. We therefore applied sequences of standard unit root tests, i.e. the augmented Dickey Fuller tests, namely the Jobert, 1992, procedure, as well as the Schmidt and Phillips, 1992, test and the Kwiatkowsky, Phillips and Shin (KPSS), 1992, test, to investigate which of the I(0), I(1), I(2) assumptions is most likely to hold true. The results of the Jobert procedure, Schmidt and Phillips' test and the KPSS tests are shown in Table 1. Note that all variables were transformed in natural logarithm, and in what follows lower-case letters denote the natural logarithm of the corresponding variable.Most variables seemed well characterised as an I(1) process, some with non-zero drift. Nevertheless, concerning *u*, *cp*, *pc-p* and *tr*, the results given by the different tests were not all concomitant and did not allow us to decide between an I(0) or I(1) process: they diverged on the number of lags to introduce to have white noise residuals, and on the applied unit root test. The fact that real wages were I(1) supported the estimation of a real model. While considering wages and prices separately, one was likely to introduce variables I(2) in estimations that would not be compatible with the econometric methodology adopted here. Moreover, this would strongly complicate the partition between marginal and conditional models and would not consequently permit us to provide an enriched reading of unemployment formation. Besides, econometric estimations available in France highlight the unit indexing of wages on prices,

Table 1. Unit Root Test Results

Non-agricultural	Jobert Tests			Schmidt-Phillips Tests			KPSS Tests		
manufacturing sectors	Bic	Hannan	Kmax	Bic	Hann.	Kmax	0	4	8
w - *p*: real labour cost	I(1)	I(1)	I(1)	I(1)+T	I(1)+T	I(1)+T	2.52 I(1)	0.53 I(1)	0.31 I(1)
prodh: hourly productivity	I(1)+T	I(1)+T	I(1)+T	I(1)+T	I(1)+T	I(1)+T	2.16 I(1)	0.49 I(1)	0.30 I(1)
tr: replacement rate	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	1.38 I(1)	0.30 I(1)	0.19 I(1)
cp: complete set of SMIC hikes	I(1)	I(0)+C	I(1)+T	I(1)	I(0)+T	I(1)+T	2.16 I(1)	0.46 I(1)	0.28 I(1)
r: real interest rate	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	0.82 I(1)	0.20 I(1)	0.14 (?)
ec: quit ratio	I(1)	I(1)	I(1)	I(1)	I(1)+T	I(1)+T	1.92 I(1)	0.43 I(1)	0.27 I(1)
mm: mismatch	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	1.87 I(1)	0.40 I(1)	0.24 I(1)
u: unemployment rate	I(0)	I(0)	I(1)	I(1)+T	I(1)+T	I(1)+T	2.30 I(1)	0.48 I(1)	0.28 I(1)
h: working hours	I(1)	I(1)	I(1)	I(1)+T	I(1)+T	I(1)+T	2.26 I(1)	0.49 I(1)	0.29 I(1)
coin: global wage wedge	I(1)	I(1)	I(1)	I(1)+T	I(1)+T	I(1)+T	1.89 I(1)	0.41 I(1)	0.24 I(1)
pc - *p*: terms of exchange	I(0)	I(1)	I(0)	I(1)	I(1)	I(1)	1.32 I(1)	0.30 I(1)	0.18 I(1)
coinfs: fiscal and social wedge	I(1)	I(1)	I(1)	I(1)+T	I(1)+T	I(1)+T	1.80 I(1)	0.41 I(1)	0.26 I(1)
coins: social wedge	I(1)	I(1)	I(1)	I(1)+T	I(1)+T	I(1)+T	2.04 I(1)	0.48 I(1)	0.29 I(1)

Table 1. (Continued) Unit Root Test Results

Non-agricultural	Jobert Tests			Schmidt-Phillips Tests			KPSS Tests		
manufacturing	Bic	Hannan	Kmax	Bic	Hann.	Kmax	0	4	8
coinf: fiscal wedge	I(1)+C	I(1)+C	I(1)	I(1)	I(1)	I(1)	0.82 I(1)	0.21 I(1)	1.49 I(1)
css: employees' social contributions rate	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)+T	2.16 I(1)	0.50 I(1)	0.30 I(1)
cse: employers' social contributions rate	I(1)	I(1)	I(1)	I(1)+T	I(1)+T	I(1)+T	1.00 I(1)	0.24 I(1)	0.15 I(1)
vat: value added tax	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)+T	0.31 I(1)	0.08 I(0)+T	0.06 I(0)+T
tir: income tax rate	I(1)	I(1)	I(1)	I(1)+T	I(1)+T	I(1)+T	1.16 I(1)	0.25 I(1)	0.17 I(1)
prog: progressiveness of social wedge	I(1)	I(1)	I(1)	I(1)	I(1)	I(1)	1.53 I(1)	0.34 I(1)	0.21 I(1)

Notes: a) All series are in logarithm and all the tests were programmed with the GAUSS software. b) Unlike the Jobert and Schmidt-Phillips unit root tests, the null hypothesis of the KPSS test is here the deterministic non-stationarity around a linear trend against the alternative hypothesis of stochastic non-stationarity (presence of a unit root). The critical value at a 5% level is 0.463. c) The question mark "?" in some boxes indicates the difficulty in concluding between an I(0) or I(1), given that the computed test is too close to the 5% critical value.

which also justified the choice of a real model. Therefore, nominal rigidities would not explain unemployment in the long-term horizon that is ours.[2]

B. Estimation Strategy

Given that most of the series in our database are non-stationary trending variables, our analysis is conducted within a framework that allows both for non-stationary and potentially co-integrated variables. Our econometric procedure is close to the multivariate co-integrated systems analysis developed originally by Johansen (1988), then expanded and applied in Johansen (1995). It consists of full information maximum likelihood estimation (FIML) of a system characterised by r co-integrating vectors (CIVs). Under conventional hypotheses the statistical model is the following (see Rault 1997 for a detailed presentation):

$$\Delta X_t = \sum_{i=1}^{P-1} \Gamma_i \, \Delta X_{t-i} + \alpha\beta' X_{t-1} + \Phi D_t + \varepsilon_t, \qquad t = 1,..,T \qquad (1)$$

where (X_t), t = 1,...,T, is a dimensional vector process composed of stochastic variables, $\varepsilon_t \sim$ iid, N $(0_n, \Sigma)$, Γ_i, i = 1,...p-1 are (n, n) matrices, supposed constant in time, α and β are (n, r) non-singular matrices of rank $0 < r < n$, D_t is a vector of non-stochastic variables (constant drift, linear deterministic trend, ...), and Σ is a regular, positive define variance-covariance matrix.

The co-integrating vectors are the β_j columns of the β matrix. In particular, the $\beta_j' X_t$ (j = 1,.., r) can be regarded as stationary linear combinations of non-stationary variables and the α as the weights of these different combinations in each equation of the model.

Then, once the number of co-integrating vectors was determined it seemed natural to more precisely apprehend the structure of the adjustment space, spanned by the α. Applying a test on α, boils down to asking oneself if the long run relation(s) belongs to all the model equations. It deals with a weak

[2] An alternative coherent approach with nominal rigidities supposes the consideration of a modelling of variables in growth rates and not in level. This leads to an estimate of a Philips curve and not a wage curve. For an example of that estimation strategy on French data, cf. Heyer, Le Bihan and Lerais (2000).

exogeneity test of the different variables of the system for long run parameters, whose aim is to check if the sufficient condition given by Johansen (1992) checks out empirically. According to Johansen, if the (X_t) variables of the system are divided into (Y_t, Z_t), a sufficient condition for a variable (or a group of variables) Z_t to be weakly exogenous for long run parameters is that the co-integrating vectors do not belong to the model equation(s) describing the evolution of ΔZ_t. In this case, the joint density function can be factorised into two blocs whose parameters vary freely: a ΔZ_t marginal model gathering the weakly exogenous variables for the long run parameters of the VAR-ECM model, and a conditional ΔY_t model composed of the other equations. The co-integration vectors can then be estimated only from the conditional model, which enables the size of the system to be reduced without losing any information from the full VAR-ECM.[3]

Finally, once the co-integrating relationships had been identified (see Johansen and Juselius, 1994 for a detailed presentation), particular structural hypotheses on the α and β matrices could be tested using asymptotically chi-squared distributed test statistics.

C. Estimation Results

Before choosing the final model, we made much prior estimation, whose main results we can only summarise. Firstly, it was impossible to estimate a satisfactory model when the complete set of SMIC hikes and progressiveness indicators were taken into account. Moreover, it was impossible to get a satisfactory estimation when the Sneessens (1994) indicator was introduced and the estimations were made using the Jackman, Layard and Savouri (1991) indicator, which was significantly different from zero in almost all the prior estimations we made. We had to limit wage wedge split up between internal terms of exchange and fiscal and social wedge without being able to split up within the latter. In other respects, the most satisfactory models were obtained using hourly labour cost and productivity specifications (and not per capita). Finally, modelling attempts with unemployment rate rather than its logarithm were unsuccessful.

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[3] See Rault (2000) for a discussion on weak exogeneity and causality.

The model adopted was composed of ten variables (unemployment rate, hourly real cost, hourly productivity, replacement ratio, mismatch, real interest rate, quit ratio, working hours, the terms of exchange, fiscal and social wedge (which combine four deduction rates)). The variable formulation of the statistical model stated by equation (1) is given by the vector $X_t = (u, w\text{-}p, prodh, tr, mm, r, ec, h, pc\text{-}p, coinfs)'_t$. Its purpose is to study the interdependences between these variables, transformed in natural logarithm, without making any *a priori* hypothesis on the value of the elasticities linking them and to test the existence of long run relations.

Two Co-integration Relations

The lag length choice used in the specification of the unrestricted VAR-ECM model is based on the results of two information criteria (Schwarz's Bayesian information criterion and the Hannan-Quinn criterion), and on global Fisher's tests. These different methods all indicate an optimal value of two quarters. One must notice that the lag length choice used in the VAR-ECM model is a crucial stage of the analysis, since it can noticeably affect the determination of the dimension of the co-integrating space, that is, the rank of the Π matrix: simulations by Boswijk and Franses (1992), and Gonzalo (1994) show that under-fitting leads to underestimating the number of long run relations, whereas over-fitting leads to overestimating this number. Moreover, these simulations show that asymptotic distributions of the trace and eigenvalue tests proposed by Johansen (1988), can be rather bad approximations of the true small sample distributions, and should therefore be used with caution. Boswijk and Franses (1992) advocate using the corrected version of these two tests, which perform better in the case of small or medium sample size. These small sample corrected versions of test statistics denoted by $\lambda_{\max}^{adj}$ and λ_{trace}^{adj}, are obtained by pre-multiplying the usual test statistics by (T - np) instead of T, where n is the model variable number and p the VAR order.

Once the lag length used in VAR-ECM model specification has been determined, the next step is to test the number of co-integrating relationships existing between the ten variables of the system. At this stage, one aforementioned point must be emphasised: the asymptotic distributions of the co-integration tests depend on the deterministic components (which are not explicitly modelled) in the system. Specifically, these tests depend on the possible presence of a constant or linear deterministic trend in the long run relations. For instance, if the linear deterministic trend is not constrained to lie in the co-integrating space, the presence of a non-zero deterministic trend outside the long run relations indicates the presence of a quadratic trend in every component of the system taken in level, since the system is written in first differences. In the same way, if the constant is unrestricted, this modelling allows for a linear deterministic trend in the level of series.

To know how to model these deterministic components, one can possibly use the results of the sequences of standard unit root tests applied previously, especially the Schmidt-Phillips (1992) ones, which have not eliminated the possibility that some of these series have a linear drift. That's why all the co-integrating rank tests have been investigated in a system with an unrestricted constant, as well as a linear deterministic trend constrained to lie in the co-integrating space. The small sample corrected versions of the two LR test statistics (trace test and Lambda max test) and also the critical value taken from Johansen (1995), are reported in Table 2.

Table 2. Estimation of the Number of Co-integrating Relationships

Ho against Ha	$\lambda_{\max}^{adj}$		λ_{trace}^{adj}	
	Statistic	Critical value[a]	Statistic	Critical value[a]
r = 0 against r = 1	77.22 **	66.2	310.90 **	263.4
r ≤ 1 against r = 2	60.46	61.3	233.60 *	222.2
r ≤ 2 against r = 3	48.07	55.5	173.20	182.8
r ≤ 3 against r = 4	39.97	49.4	125.10	146.8
r ≤ 4 against r = 5	32.50	44.0	85.14	114.9
r ≤ 5 against r = 6	16.97	37.5	52.64	87.3
r ≤ 6 against r = 7	14.43	31.5	35.66	63.0
r ≤ 7 against r = 8	10.52	25.5	21.23	42.4
r ≤ 8 against r = 9	7.67	19.0	10.71	25.3
r ≤ 9 against r = 10	3.03	12.2	3.037	12.2

Note: [a] critical value at 5 %. ** is significant at 1% level, * is significant at 5% level.

These test statistics indicate the existence of two co-integrating relationships between the ten variables considered.[4,5] The estimation of the co-integrating vectors and of the adjustment coefficients will be given later.

Once the co-integrating rank was determined, systematic LR tests on the deterministic components were made. These tests confirmed the results and led to the acceptance of a specification of the Vector Error Correction Model (VAR-ECM), with an unrestricted constant in the short run, as well as a linear deterministic trend constrained to lie in co-integrating relationships. From here on model specification was completely determined (two lags, two co-integrating relationships and a linear deterministic trend constrained to lie in co-integrating relationships).

D. Weakly Exogenous Variables and that Excluded from Co-integrating Space

The next step is to ask oneself if some system variables can be considered as weakly exogenous for the parameters of the two co-integrating relationships found previously. If so, these parameters can then be estimated without loss of information from the more manageable conditional model, having been extracted from the full VAR-ECM model. This hypothesis of weak exogeneity is expressed by the nullity of some coefficients of the α matrix. Table 3 produces the results of these weak exogeneity tests.

The results can be synthesised as follows: at a 5 % level, one rejects the weak exogeneity of real labour cost, of unemployment rate, of working hours, of mismatch, of the terms of exchange, of hourly productivity and of quit ratio. Moreover, at a 5 % level, the joint weak exogeneity hypothesis of the remaining three variables is easily accepted by the data ($\chi^2(6) = 5.24$ (0.51)). Therefore, we chose to estimate the two long run relations from a partial VAR-ECM model composed of seven equations (*w-p*, *u*, *h*, *mm*, *pc-p*, *prodh*, *ec*), conditional to

[4] The outcome of the co-integration analysis remains unchanged if we use the critical values recently tabulated by Pesaran, Shin and Smith (1999).

[5] Given that the calculated statistical value of the $\lambda_{\max}^{adj}$ test is very close to the 5 % critical value, it is reasonable to think as economic theory suggests, that there exist two long run relationships between the considered variables: that is what it indicates in addition to the λ_{trace}^{adj} test.

Table 3. Weak Exogeneity Tests of the Different Variables for all Long Run (α and β) Parameters

Variable	Weak exogeneity	LR test statistic
w - p	rejected	χ^2 (2) = 19.13 (0.00)
u	rejected	χ^2 (2) = 11.39 (0.00)
tr	not rejected	χ^2 (2) = 2.56 (0.27)
r	not rejected	χ^2 (2) = 0.97 (0.61)
coinfs	not rejected	χ^2 (2) = 4.03 (0.13)
h	rejected	χ^2 (2) = 19.27 (0.00)
mm	rejected	χ^2 (2) = 17.23 (0.00)
pc - p	rejected	χ^2 (2) = 12.84 (0.00)
prodh	rejected	χ^2 (2) = 10.78 (0.00)
ec	rejected	χ^2 (2) = 27.98 (0.00)

Note : The number in brackets indicates the marginal asymptotic level, namely the probability of exceeding the value of the computed statistic. Thus a marginal asymptotic level of 27 % (0.27), for instance, means that for an α level smaller than 27 %, the null hypothesis Ho of weak exogeneity of the variable under study is accepted.

the three equations describing the evolution of the weakly exogenous variables (*tr*, *r*, *coinfs*).

Then a first sequence of tests was applied in order to determine if some system variables could be considered excluded from the two long run relations. The following table shows that at a 5% level, replacement rate, real interest rate and working hours do not belong to the co-integrating space. Moreover at a 5 % level, the joint exclusion hypothesis of these three variables of the co-integrating space is easily accepted by data ($\chi^2(6) = 2.30$ (0.89)). The replacement ratio and the real interest rate are thus both weakly exogenous and excluded from the co-integrating space, which in other words means that they only have an influence on the short run dynamic of the price and wage schedule.

Next it is interesting to ask oneself if there exists a variable belonging to the co-integrating space, which constitutes a co-integration relation alone. In this respect, Table 5 presents the results of the stationarity tests around a linear deterministic trend of the different variables. For instance, to test if the

Table 4.Tests of the Structure of Co-integrating Space

Variable	Belonging to co-integrating space	LR test statistic
w - p	yes	χ^2 (2) = 31.46 (0.00)
u	yes	χ^2 (2) = 15.91 (0.00)
tr	no	χ^2 (2) = 0.19 (0.90)
r	no	χ^2 (2) = 1.12 (0.57)
h	no	χ^2 (2) = 0.50 (0.77)
coinfs	yes	χ^2 (2) = 6.36 (0.04)
pc - p	yes	χ^2 (2) = 6.97 (0.03)
prodh	yes	χ^2 (2) = 6.39 (0.04)
ec	yes	χ^2 (2) = 26.15 (0.00)
trend	yes	χ^2 (2) = 6.46 (0.03)

Notes: a) Some of the results given in this table were obtained after several iterations. In fact, two weekly exogenous variables were shown moreover not to belong to the co-integrating space. We found it more logical to take these two pieces of information into account step by step, instead of directly placing these two variables in the short run. For this purpose, we first estimated a VAR-ECM in which the replacement rate only belonged in the short run dynamic, then re-tested in this framework, to see if the other variables belonged to the co-integrating space. b) The number in brackets indicates the marginal asymptotic level, namely the probability of exceeding the value of the computed statistic. Thus a marginal asymptotic level of 90 % (0.90) for instance, means that for an α level smaller than 90 %, the null hypothesis Ho of exclusion from the co-integrating space of the variable under study is accepted by the data.

unemployment rate *u* is stationary around a linear deterministic trend, one has to test if vector b’ = (0 1 0 0 0 0 0 a) belongs to the co-integrating space. The results of these tests are categorical, since they reject the stationarity hypothesis around a linear deterministic trend of the seven variables belonging to the co-integrating space in every case. Thus, the results of the stationarity tests applied in the multivariate framework, where the interdependences between variables are explicitly modelled, are concomitant with those applied previously in the univariate framework. These tests indicate that the variables are characterised by a stochastic non-stationarity (namely integrated of order 1), rather than a deterministic non-stationarity (namely stationary around a linear deterministic trend).

Table 5. Stationarity Tests of the Different Variables Around a Linear Deterministic Trend

Variable	Stationarity around a linear deterministic trend	LR test statistic
w - p	rejected	χ^2 (6) = 33.11 (0.00)
u	rejected	χ^2 (6) = 31.02 (0.00)
mm	rejected	χ^2 (6) = 52.65 (0.00)
coinfs	rejected	χ^2 (6) = 29.74 (0.00)
pc - p	rejected	χ^2 (6) = 58.59 (0.00)
prodh	rejected	χ^2 (6) = 41.84 (0.00)
ec	rejected	χ^2 (6) = 34.03 (0.00)

Table 6 gives the estimation of the two long run relations and the error correction coefficients obtained from the conditional model.

E. PS and WS Identification

Spontaneously, each of the two co-integrating vectors has an unemployment rate coefficient with an opposite sign, which indicates both a price and wage setting behaviour. Nevertheless, it is important to notice that these two co-integrating vectors have no economic meaning at this stage, and are nothing other than a vectorial basis of the co-integrating space. Strictly speaking, they are obtained as the eigenvectors of the long run Π matrix and any linear combination of these two vectors forms a new co-integrating relationship between the seven variables. These vectors then have only a purely statistical value. Econometric modelling alone does not allow the structural form of (WS) and (PS) curves to be determined ex nihilo. Therefore, it does not eliminate a theoretical consideration of the form of structural equations, but requires on the contrary, the a priori specification of identification conditions,

Table 6. Maximum Likelihood Estimations of the Normalised Co-integrating Vectors and of the Error Correction Coefficients

Variables	Normalised co-integrating vectors (β matrix)	
w - p	1.000	1.000
u	0.254	-0.506
mm	-0.083	-0.000
pc - p	-0.733	1.042
prodh	0.087	-3.012
ec	-0.403	0.260
coinfs	0.764	1.642
trend	-0.001	0.014
Variables	**Error correction coefficients (α matrix)**	
w - p	-0.091	0.087
	(-3.84)	(6.77)
u	0.047	0.155
	(1.73)	(4.50)
mm	0.294	0.054
	(3.52)	(1.20)
h	-0.062	-0.034
	(-3.52)	(-4.06)
pc - p	-0.045	0.053
	(-1.64)	(3.48)
prodh	-0.042	0.068
	(-1.96)	(4.06)
ec	0.430	0.122
	(5.10)	(2.40)

Note: The number in brackets represents the t stats.

using a theoretical model, before beginning the estimation. The identification of the two curves is investigated here using the following two theoretical restrictions: the wage determination (WS curve) is supposed to be made independently of productivity level (the Manning, 1993, identification restriction) and unemployment is not supposed to influence wage determination (PS curve). Structural forms are then obtained by calculating the two linear combinations of the estimated co-integrating vectors, which satisfy identification constraints. It must be emphasised that it is not a test, but simply a change of basis in the co-integrating space, in order to statistically distinguish between the two structural equations. Thus, these constraints do not affect the level and evolution of equilibrium unemployment estimation (which is robust to identification choice). After normalisation, the two (just) identified long run relations are given by,

$$\text{(PS)} \quad w - p = 0.055\, mm + 0.138\, pc\text{-}p + 0.944\, prodh - 0.041\, coinfs + 0.181\, ec - 0.004\, trend \qquad (2)$$

$$\text{(WS)} \quad w - p = -0.232\, u + 0.080\, mm + 0.679\, pc\text{-}p + 0.693\, coinfs + 0.384\, ec - 0.001\, trend$$

Finally, over-identifying restrictions were tested, the results are reported in Table 7: the exclusion of the fiscal and social wedge, of the terms of exchange and of the linear deterministic trend from the PS curve are accepted at a 5% level.

Additional structural hypotheses were also tested, as the exclusion of *mm* and *ec* variables from (PS), but were all rejected. The presence of these variables in price equation is not theoretically justified, which is one reason for dissatisfaction. Finally, the two over-identified long run relations are given by:

$$\text{(PS)} \quad w - p = 0.073\, mm + 0.204\, prodh + 0.230\, ec \qquad (3)$$

$$\text{(WS)} \quad w - p = -0.050\, u + 0.078\, mm + 0.117\, pc\text{-}p + 0.159\, coinfs + 0.274\, ec + 0.001\, trend$$

Table 7. Tests of Over-identifying Restrictions

Null hypothesis	Accepted hypothesis	LR test statistic
Exclusion of h from (PS) and (WS), and exclusion of *pc* - *p* from (PS)	yes	χ^2 (3) = 0.94 (0.82)
Exclusion of h from (PS) and (WS), and exclusion of *pc* - *p* and *coinfs* from (PS)	yes	χ^2 (4) = 0.95 (0.92)
Exclusion of h from (PS) and (WS), and exclusion of *pc* - *p*, *coinfs* and of the linear deterministic trend from (PS)	yes	χ^2 (5) = 6.21 (0.29)

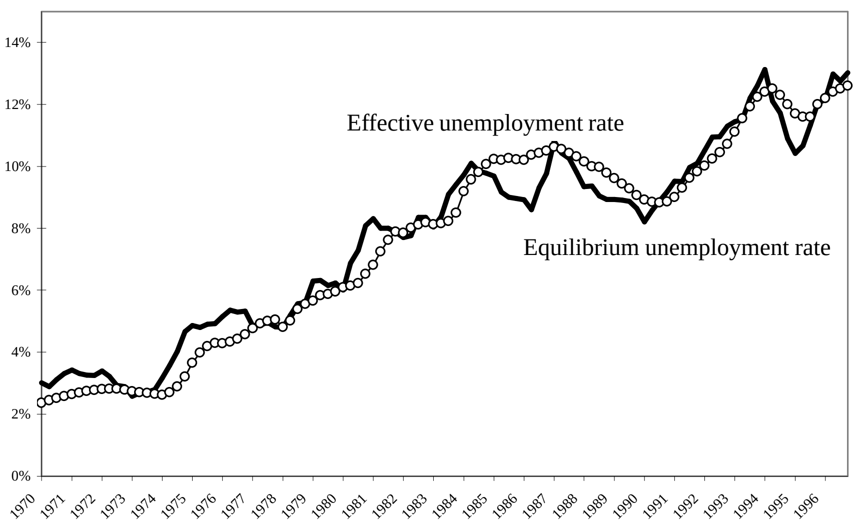
It is now possible to determine the equilibrium unemployment from the two estimated structural equations. For this purpose, one must resolve the equilibrium of the partial system of the labour market obtained. This resolution gives the following expression of equilibrium unemployment.

$$u^* = -4.1\, prodh + 2.34\, pc - p + 0.1\, mm + 0.88\, ec + 3.18\, coinfs + 0.02\, trend \tag{4}$$

All equilibrium unemployment determinants have a sign in accordance with the theoretical idea. Equilibrium unemployment decreases when productivity growth exceeds the trend, which corresponds to an annual growth rate of over 2% (this is close to the average rate of productivity growth over the period covered). Unemployment increases with the terms of exchange (the oil crisis for instance has increased unemployment, since it led to a higher rise in consumption prices than added value prices), with the growth of skill mismatch, quit ratio, fiscal and social wedge and its components. The contributions of the terms of exchange and of mismatch remain quite small (about 5% of the equilibrium unemployment increase).

Figure 1 represents effective unemployment rate and equilibrium unemployment rate. The latter is defined up to a constant, which requires a choice in reference value: we choose the 1973 average rate, so we assumed equality between effective unemployment and equilibrium unemployment in that year. Neither equilibrium unemployment nor its determinants were smoothed here.

Figure 1. Effective Unemployment Rate and Equilibrium Unemployment Rate



F. Diagnostic Tests on the Residuals

The last step is to establish whether the estimated VAR-ECM model is a reasonably congruent representation of the data. We have therefore implemented two kinds of tests: misspecification and constancy tests.

Firstly, several test statistics were calculated in order to check the quality of the multivariate estimation (Lagrange Multiplicator (LM) test and Ljung-Box test for serial correlation of order 16, ARCH (Autoregressive Conditional Heteroscedasticity) tests, Jarque-Bera normality test). The tests constitute a good way to detect possible failings of some hypotheses made during the system estimation. These tests indicate that the conditional VAR-ECM model is well behaved and not subject to misspecification, since the usual hypotheses concerning the residuals of each of the seven equations are verified (see Table 8).[6]

Table 8. Specification Tests of the Residuals of the Conditional VAR Model

Equation	LB (16)	WHITE (F-Form)	ARCH (16)	JB (2)
Dw - p	19.43	0.69	20.15	1.59
	(0.14)	(0.87)	(0.21)	(0.44)
Du	14.64	1.37	18.99	32.21
	(0.40)	(0.16)	(0.26)	(0.00)
Dmm	17.03	1.57	15.81	4.53
	(0.25)	(0.07)	(0.46)	(0.10)
Dh	24.55	0.67	24.85	61.39
	(0.03)	(0.93)	(0.07)	(0.00)
Dpc - p	30.23	0.98	23.79	4.21
	(0.007)	(0.52)	(0.09)	(0.12)
Dprodh	11.69	0.56	11.74	5.68
	(0.63)	(0.92)	(0.76)	(0.05)
Dec	21.87	1.01	13.86	75.01
	(0.08)	(0.48)	(0.60)	(0.00)

Note: The number in brackets indicates the marginal asymptotic level, namely the probability to exceed the value of the computed statistic. Thus a marginal asymptotic level of 14 % (0.14) for instance, means that for a Ho level smaller than 14 %, the null hypothesis Ho of absence of residual serial correlation of order 16 is accepted by data.

[6] The residuals of the conditional VAR-ECM model equations have good properties on the whole: they do not suffer from serial correlation, are not of ARCH type, even if they sometimes have normality problems. This lack of normality assumption in some equations is not actually very serious for the conclusions of the study, since as noted by Johansen (1995), the asymptotic properties of the Maximum Likelihood method only depend on the i.i.d assumption of the errors.

Secondly, the conditional and marginal VAR-ECM models were re-estimated by recursive least squares until 1996/4 and One Step Ahead, as well as performing Backward and Forward Chow tests in order to appreciate the parameter constancy through time. The graph examination does not reveal any particular break and was not reported here.

Thus, the misspecification and constancy tests indicate the estimated conditional VAR-ECM model to be a satisfactory representation of the data.

IV. Conclusion

One can consider a great number of possible explanations as to the rise and persistency of unemployment in France. The aim of this paper was to confront some of these determinants with data in a WS-PS model estimation framework on French macroeconomic data.

First and foremost, we chose a selection of about fifteen variables whose influence rested on explicit micro-economic bases and which was founded on a general equilibrium framework. To this first filter, of a theoretical order, a second one of a statistical order was added, resulting in the possibility of building indicators for these determinants, then a third one of an econometric order was added, resulting in the model estimation. Finally, only five variables reached the end of this procedure. The equilibrium unemployment increase in France reflects the slowing down of productivity gains, the increase of social and fiscal wedges, the deterioration in job security and in a more marginal way, the terms of exchange increase and the skill mismatch.

Considering a richer set of variables and a different methodology, this paper confirms the impact of some unemployment determinants in a unified framework, found in previous studies incorporating a limited number of candidates to explain equilibrium unemployment (Bonnet and Mahfouz, 1996; L'Horty and Sobczak, 1997; Cotis, Méary and Sobczak, 1997). It gives a main role to the rise of social and fiscal wedge, as do two of the previous studies (LS, 1997 and CMS, 1997). It is also compatible with a predominant role attributed to the influence of real interest rate, when this influence is well mediated by a downturn in productivity gain, also in keeping with the three studies. It also concludes that the terms of exchange play a role in the formation of French unemployment, like one of the studies (BM, 1996). Our empirical investigation also shows the influence of skill mismatch and of the employment protection system, via the quit ratio, which has not been obtained (nor introduced) before, in the existing applied French studies using time series. In addition, our study leads to a questioning of the influence of numerous other determinants: the replacement rate would not have had any impact on the increase of equilibrium unemployment (contrary to the LS, 1997, results), and would be the same for other determinants which were not introduced in previous studies: the lesser digressiveness of social wedge, the reduction of working hours and the minimum wage increase.

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