Incidence of Social Security Contributions and Taxes
Empirical Evidence from Austria*

Rupert SENDLHOFER
Institute of Public Economics
University of Innsbruck
Universitaetsstrasse 15
A-6020 Innsbruck

e-mail: rupert.sendlhofer@uibk.ac.at
Tel.: ++43-512-507-7158
FAX: ++43-512-507-2788

February 2001

Abstract

In economic policy debates high unemployment rates in Europe are often seen as a result of increasing taxes and social security contributions. But the effects of the total tax wedge on real wages and employment depend on the economic incidence of social security financing. The paper investigates the amount of tax shifting in the case of taxes and social security contributions in Austria. The analysis of the macroeconomic process of wage formation is developed from a right-to-manage model. The estimation is based on the Engle-Granger procedure for a single cointegrating vector and on Johansen's procedure for multiple cointegrating vectors.

JEL classification: E24, H22, J32, J51
Keywords: social security contributions, payroll taxes, tax incidence

Comments on this paper are gratefully acknowledged.

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* Earlier versions of this paper were presented at the 2nd Austrian Labor Workshop, Vienna, September 2000 and at the Research Seminar of the Institute of Public Economics, University of Innsbruck, March 2000. I would like to thank the participants for their helpful comments. I am especially grateful to Clemens Fuest, Herbert Stocker, Engelbert Theurl and Hannes Winner for comments and discussions. I also thank Gudrun Biffl, Claudia Felix, Alois Guger, Helmut Hofer, Robert Koza, Maxime Ladaique, Roland Löffler, Norbert Rainer, Paul Schreyer, Oumar Traoré and Sepp Zuckerstätter for help with the data. Of course, the usual disclaimer applies.

1. Introduction

The simple hypothesis that higher taxes and social security contributions on labour result in higher real wages and higher unemployment rates is a commonplace in economic policy discussions. From a theoretical point of view, the effects of the tax wedge (taxes and social security contributions) on real wages and employment depend on the incidence of these instruments. The higher is the proportion of the tax burden borne by the employees, the lower are ceteris paribus the effects on employment. The indirect effects on the labour market via a lower demand resulting from lower net wages are not considered here. In addition, we do not take account of the fact that employers may change the product prices to shift their tax burden on to their customers, i.e. product prices are exogenously given.

This difference between the arguments in policy discussions and theoretical considerations emerges from the difference between formal and economic incidence. A critical assessment of taxes should be based on information about the economic incidence.

Therefore, in this paper we investigate who (the employer or the employee) has to bear the burden of taxes and social security contributions on labour income. With this information on the economic incidence we will be in a position to analyse the effects of higher taxes on real wages and employment. The empirical parts of the paper are concentrated on Austria.

We would be especially interested whether it makes a difference which side of the labour market (employer or employees) is taxed. Unfortunately, it is difficult to discriminate between the effects of different tax instruments due to multicollinearity and a small number of observations. In addition, tax/benefit linkages of the social security system are not considered (see OECD (1990, p. 156) on the role of benefits received).

There exist a large number of articles on the effects of taxes on the labour market. Seminal papers in this field are Bean, Layard and Nickell (1986) and Knoester and van der Windt (1987). A survey on the effects of taxes on the labour markets can be found in OECD (1994, ch. 9). Recent contributions are for example Sørensen (1997), Daveri and Tabellini (2000), and Fuest (2000). The incidence of taxes in selected OECD countries is estimated in Tyrväinen (1995). Empirical results for Austria can be found in Pichelmann (1993), Hofer and Pichelmann (1996) and Felderer, Hofer and Schuh (1999).

The structure of the paper is as follows. In section 2 we present a right-to-manage model with taxes. In section 3 we develop the corresponding empirical model. In the following sections this model is estimated with Austrian data from 1965 to 1997. In section 4 we test whether our data is non-stationary. In section 5 we estimate the wage-setting equation with the Engle-Granger procedure. In section 6 we make use of the Johansen approach. In Section 7 some conclusions are presented.

2. Economic Model

Our theoretical model consists of an aggregate wage-setting curve and an aggregate price-setting curve. The simultaneous interaction of these two curves results in equilibrium real wages and employment. The wage-setting curve is obtained, for given prices, from a wage bargaining model. The price-setting curve is based on the price setting behaviour of the firms and derived for given nominal wages. The equilibrium of real wages and employment determines the so called 'non-accelerating inflation rate of unemployment' (NAIRU). At this equilibrium level of unemployment the claims of unions and firms are consistent with one another. As a consequence the inflation rate is constant. If the price-setting behaviour of the firms is not consistent with the wage-setting behaviour of the unions inflationary or deflationary processes will take place (see Layard, Nickell and Jackman, 1991; Booth, 1995, p. 224ff.; Franz, 1999a).

The wage-setting curve can be derived from a model of the labour market. In a perfectly competitive labour market it does not make a difference which side of the market is formally taxed (see OECD, 1990; Muysken, van Veen and de Regt, 1999). The economic effects will not be altered by a change in the formal incidence. A model of this kind is not useful for our purpose. In order to analyse the economic effects of a change in the formal incidence a model of imperfect competition has to be used. There are two standard models for a labour market with imperfect competition: efficiency-wage models and wage bargaining models. Within the latter the right-to-manage kind of model is widely used. We follow this tradition in the literature.

From both standard models follows a negative slope of the wage-setting curve with respect to the unemployment rate. Relevant shift factors of the curve are, for example, an increasing mismatch on the labour market, higher unemployment benefit replacement rates or an increasing tax wedge (see Pichelmann, 1993; Franz, 1999a, p. 161).

In the following we assume a perfectly competitive product market. In this case, the price-setting curve is the traditional labour demand function. On imperfectly competitive product markets firms would set prices as a markup over marginal cost.

Typical shift factors of the price-setting curve are the technical progress and the intensity of competition on product markets (see Pichelmann, 1993; Franz, 1999a, p. 162ff.).

In addition to the price-setting and wage-setting curves we have to consider individual labour supply. Let us assume that the aggregated individual labour supply is exogenous and independent from the real consumption wage. The intersection of this labour supply curve with the price setting curve would give an equilibrium without unemployment. The NAIRU is the difference between the equilibrium on the just mentioned completely competitive labour market and the equilibrium on a labour market with a wage-setting curve as discussed above (Franz, 1999a, p. 156f.).

Graphical expositions of the model can be found in Pichelmann (1993) and Franz (1999a). We prefer to derive our theoretical predictions within a consistent model. The small open economy of our simple model consists of firms, capitalists, workers and the government (see Fuest (2000) and Fuest, Huber and Wöhlbier (2000) for an analysis of the model; see Tyrväinen (1995), p. 8ff. for a similar model).

The firms produce the consumption good Y with the two factors of production labour L and capital C. The price of the consumption good is determined on the world market, i.e. on a perfectly competitive product market.

The factor capital is internationally mobile, thereas the factor labour is internationally immobile. Each worker inelastically supplies one unit of labour. The number of firms and the price of the consumption good is normalized to unity. The production function of the firms is of the Cobb-Douglas form:

$$Y = AL^{a} K^{b} \tag{1}$$

where $A > 0, a > 0, b > 0$ and $a + b < 1$.

The decreasing returns to scale of this production function imply a profit. The below modelled bargain between representative firm and trade union is just about this profit.

The representative firm will maximize its profits P.

$$P = AL^{a} K^{b} - wL - (1 + t)K \tag{2}$$

The price of capital is normalized to unity. w is the real product wage and t is the tax on capital income. In the empirical parts of the paper the real product wage is represented by *wcr*. The maximization of (2) results in the following marginal productivity conditions:

$$w = aAL^{a-1}K^{b} \tag{3}$$

$$1+t = bAL^{a}K^{b-1} \tag{4}$$

From this conditions we can derive the factor demand functions:

$$L = L(w,(1+t)) \tag{5}$$

$$K = K(w,(1+t)) \tag{6}$$

where $L_w, L_t, K_w, K_t < 0$.

We can now define the wage elasticity of labour demand:

$$e_{L,w} \equiv L_w w / L = -\frac{1-b}{1-a-b} < 0 \tag{7}$$

In contrast to the detailed modelling of the firms and their labour demand decision the other sectors of the model are only represented through their budget restrictions. In this restrictions we have to consider the remaining tax instruments of the government.

The capitalists own the firms, get their whole income from the profits of the firms and have to pay a tax on pure economic rents with rate T. They use all of their income for private consumption C^c. There is a general consumption tax with rate q which is designed according to the destination principle. The number of capitalists is normalized to unity. The budget restriction of the representative capitalist is:

$$P\,(1-T) = C^{c}(1+q) \tag{8}$$

There are N workers in our economy. Each of them supplies one unit of labour. The labour supply is fixed and therefore completely inelastic to the wage. Ex ante (which means before the labour demand decision of the firms) all workers are identical. In order to get a labour market equilibrium with unemployment, there have to be employed and unemployed workers ex post.

Employed workers have to pay a tax with rate t on their labour income and the consumption tax. In this case the budget restriction of an employed worker is:

$$w(1-t) = C^e(1+q) \tag{9}$$

Unemployed workers get a transfer b from the government in form of unemployment benefits. It is assumed that this benefits are not taxed. Nevertheless, they have to pay the consumption tax:

$$b = C^u(1+q) \tag{10}$$

Finally, we can define the budget constraint of the government (11). General government expenditure G is exogenously given and equal to R - the tax revenue net of unemployment benefits.

$$G = R = TP + wtL + tK + \frac{q}{1+q}\left(P(1-T) + w(1-t)L + (N-L)b\right) - (N-L)b \tag{11}$$

The endogenous wage is determined in a bargain between trade unions and firms. The bargaining process consists of two steps. First, unions and firms bargain about the wage w. Second, given this wage, firms make their decisions over labour demand. Therefore this kind of model is called right-to-manage model.

As there are many firms and many unions in the economy, the tax rates and the unemployment benefits cannot be influenced by single firms or unions. The number of unions is also normalized to unity.

In addition, we assume that individuals are risk neutral in our economy, i.e. the trade union is risk neutral. This assumption makes our analysis more convenient, as the utility of workers is linear in wages and unemployment benefits (see Booth, 1995, p. 90). In this special case, the union will maximize the expected income of a representative worker. All workers (employed or unemployed) are members of the union. In (12) L/N is the employment rate and 1-L/N is the unemployment rate.

$$U = \frac{L}{N}w(1-t) + \left(1 - \frac{L}{N}\right)b \tag{12}$$

If unions and firms do not agree on the wage unemployed workers will get the transfer b and firms will not produce. Therefore, in the bargaining process the threat point of the unions is b and the threat point of the firms is 0 as there are no fixed costs.

The outcome of the bargaining process is determined by the Nash bargaining solution. The concerning Nash maximand (13) in logarithmic form combines the information on the utility function of unions and firms and the respective threat points. γ is the relative bargaining power of the trade union where $0 \leq g \leq 1$.

$$W = g \log \frac{L}{N}\left(w(1-t)-b\right)+(1-g)\log P\,(1-T) \tag{13}$$

We get the bargaining solution by maximizing (13) with respect to w under the restriction of the labour demand function (5). The resulting FOC is:

$$W_w = g\,\frac{L_w\left(w(1-t)-b\right)+L(1-t)}{L\left(w(1-t)-b\right)}-(1-g)\frac{L}{P}=0 \tag{14}$$

This FOC can be simplified using the definition of the wage elasticity of labour demand (7) and the available information on factor income $\dfrac{wL}{P}=\dfrac{a}{1-a-b}$:

$$w(1-t)=fb \tag{15}$$

where $f=\dfrac{(1-b)g+(1-g)a}{a}\geq 1$.

The simplified FOC (15) has an intuitive interpretation. The after-tax wage rate is equal to the unemployment benefit multiplied by a factor $f \geq 1$. The magnitude of f depends on the relative bargaining power of the trade union. If the union has no power at all $g = 0$ and therefore $f = 1$. In this case, the wage equals the unemployment benefit. If the union can act like a monopoly on the labour market $g = 1$ and $f = \dfrac{1-b}{a} > 1$. The monopoly union case results in the highest mark-up possible. In the likely case $0 < g < 1$ this mark-up is correspondingly lower. Summing up, it may be said that a higher bargaining power of the union will result in a higher wage rate:

$$\frac{\partial w}{\partial g}=\frac{b}{1-t}\frac{1-a-b}{a}>0 \tag{16}$$

For a given unemployment benefit the net wage is fixed within this bargaining framework. Knowing this, we can ask how the wage rate is influenced by changes in the tax rate on labour income and the unemployment benefit:

$$\frac{\partial w}{\partial t}=\frac{w}{1-t}>0 \tag{17}$$

$$\frac{\partial w}{\partial b} = \frac{f}{1-t} > 0 \tag{18}$$

An increase in the labour tax rate will cause higher wage rates. The same is true for changes of the unemployment benefit.

In the next section we will develop our empirical model based on the derived results.

3. Empirical Model

The wage-setting and the price-setting curves constitute a structural model of the labour market. It would be straightforward to estimate a simultaneous equation-model (see Booth, 1995, p. 253ff.). But as we will see in section 3 our data is non-stationary. We could work with the stationary first differences of our time series. In this case we would loose information about the long-run relationship. As we are mainly interested in the long-run we have to work with our non-stationary data. Therefore we have to look for cointegrating vectors between our variables.

The candidates for cointegrating vectors are the wage-setting and the price-setting functions. With the help of the just discussed model we can construct functions (19) and (20) which will be the bases for identification and estimation. All variables are in logarithms:

$$wcr = wcr\left(\underset{+}{pr}, \underset{-}{ur}, \underset{+}{taxt}, \underset{+}{ubr}, \underset{+}{bp}\right) \quad \text{wage-setting function} \tag{19}$$

$$wcr = wcr\left(\underset{+}{pr}, \underset{+}{ur}, \underset{0}{taxt}, \underset{0}{ubr}, \underset{0}{bp}\right) \quad \text{price-setting function} \tag{20}$$

wcr is the wage cost or the product wage in real terms. In our economic model *wcr* is represented by *w*. It will depend on the GDP deflator which is not explicitly considered in our above model of real product wage. *wcr* will also strongly be influenced by labour productivity *pr* as a result of the marginal productivity condition. This is relevant for the wage-setting and the price-setting function. In addition to this quite general variables influencing the product wage our wage bargaining model pointed out further factors: the unemployment rate (measured by *ur* according to the national definition), the tax on labour income (measured by the overall tax wedge *taxt*), the unemployment benefit (measured by the unemployment benefit replacement rate *ubr*) and the bargaining power of the trade union (*bp*; represented by *g* in the economic model). The subscripts indicate the direction of the influence derived from our economic model.

Unfortunately, we do not have appropriate time series information on the bargaining power of trade unions. We searched for proxy variables but did not find one. Tyrväinen (1995, p. 10f.) suggested do use the unemployment rate. He argumented that the bargaining power depends negatively on the unemployment rate. This argument is in contrast to the result of our theoretical model. A higher bargaining power will lead to higher real wages and therefore result in an higher unemployment rate. As we have no further information we cannot us the bargaining power as an explanatory variable.

In section 4 we will learn that one of our variables is stationary and cannot be part of a cointegrating vector between non-stationary variables. This variable is the unemployment benefit replacement rate *ubr*. Nevertheless, information on the replacement rate can be used in the short-run parts of our estimations.

Theory does not tell us much about the short run dynamics of our model in the case of deviations from the long-run equilibrium values. In this case, we have to use a model which allows to differentiate between the long and the short-run. We will use an error-correction model to estimate the long-run parameters and the short-run adjustment speed in the case of deviations from the long-run equilibrium.

In section 5 we estimate our model with the Engle-Granger procedure for a single cointegrating vector. This procedure assumes that there is only one cointegrating vector. We should carefully check if this could be our wage-setting curve. If there are in fact more vectors it is impossible to identify a single long-run relationship. To get more confidence in our results we will use also Johansen's procedure for more than one cointegrating vector in section 6. With this approach we can explicitly examine the multivariate vector space and impose identifying restrictions. As a preliminary task, we test our variables for non-stationarity in the following section.

4. Data

After a short description of our data we will check it for non-stationarity. Our data is mainly taken from the national accounts of Austria and from the WIFO database. We use the following variables for our estimations:

oilp..... wholesale price index of mineral oil products

pr .... labour productivity

taxt..... overall tax wedge

ubr ..... gross unemployment benefit replacement rate

ur .... unemployment rate

wcr..... real product wage

A detailed explanation of the variables is given in Appendix A. Lower-case letters indicate variables in logarithms. Upper-case letters will indicate variables in non-logarithmic form. In addition to the explanations in appendix A we have to discuss the construction of the overall tax wedge *taxt*. The total tax on labour income can be described by the wedge between the real product wage *wcr* and the real consumption wage *wnr*. The product wage equals the real cost of a worker to an employer and the consumption wage are the net wages and salaries in real terms. We measure the tax wedge by the ratio of the real product wage to the real consumption wage. The overall tax wedge is therefore:

$$taxt = \ln\left(\frac{WC / PGDP}{WN / CPI}\right) = \ln\left(\frac{WCR}{WNR}\right) \qquad (21)$$

The wedge includes social security contributions of employers and employees, the tax on labour income, some minor parts and differences between the price indices. A detailed calculation of *WN* can be found in Appendix B. The GDP deflator is commonly used to calculate the real product wage and the consumer price index is used for the real consumption wage. Differences between the two indices may be due to import prices and indirect taxes. Unfortunately, there is no GDP deflator without indirect taxes available for Austria. Therefore, only the influence of import prices is taken into account.

This calculation of the tax wedge is derived from the so called "Lohnschere" in Franz (1999b, p. 275). Similar definitions of the tax wedge can be found in OECD (1990, p. 173ff.) and Nickell and Layard (1999, p. 3037).

In bargaining models the structure of taxation (i.e. how progressive taxes are) can have important effects even with a constant average tax rate (see Sørensen, 1997; Pissarides, 1998; and Fuest, 2000; for an empirical approach see Tyrväinen, 1995, p. 15ff.). Our simple model is restricted to a proportional tax, therefore we do not use the marginal tax rate as a regressor. Nevertheless, this would be an interesting extension of the economic model.

In any case, we have to consider that our data on the tax wedge is influenced by the progressive Austrian income tax system. In a progressive income tax system an increase in wages will usually result in higher average tax rates. Consequently, we regard the tax wedge as endogenous in our estimations.

Our time series are data for the whole Austrian economy. It is very common to restrict the analysis to the private sector or solely to the manufacturing industry (e.g. OECD, 1990; Tyrväinen, 1995). This restriction is the result of hypotheses about different wage-setting institutions in the public sector and above-average data availability in some industries of the private sector. We estimate our NAIRU model with macroeconomic data. An estimation with sector specific data should be based on a disaggregated model of the wage-setting and price-setting process.

Checking data for non-stationarity is a rather complex task. We make use of the augmented Dickey-Fuller (ADF) test for unit roots. A first plot of the data helps to decide whether there is a linear time trend and/or a drift in the data-generating process. In addition, we have to decide how many lags of the difference term to include. In other words, we do not know the true order of the autoregressive process.

After a first visual check of the data, we follow the testing procedure described in Enders (1995, p. 257). In a systematic and explicit way we check our data for the presence of deterministic regressors (intercept, time trend). The decision on the lag length is based on the Schwarz Bayesian and the Akaike information criterion. In addition, we check whether the residuals of our tests are white noise (no autocorrelation, no heteroscedasticity and normality).

If we cannot reject the presence of a unit root in our data (in logarithms), we apply the above testing procedure to the first difference of our data and so on. The results of our tests are summarized in Table 1.

(Table 1 about here.)

The second column of Table 1 reports the ADF test statistics for the data. In brackets we give information on the assumed data generating process (presence of deterministic regressors and order of the autoregressive process). The hypothesis of a unit root can only be rejected for the gross unemployment benefit replacement rate. For all other variables the null hypothesis cannot be rejected. In the third column of Table 1 there are the results for the first differences of our data. Now, we are able to reject the hypothesis of a unit root for all our time series (on different significance levels). As a result, all our time series are expected to be integrated of order one with the exception of the gross unemployment benefit replacement rate. This rate is stationary (see column 4 of Table 1).

As already mentioned, the tests for non-stationarity are quite complex. The relatively few observations make these task even more difficult. We put great emphasis on the visual inspection of the data. These conclusions were confronted with the formal tests as described above. In a few cases we were not able to get white noise residuals. This could be true to outliers or structural breaks. But all in all, we got quite convincing results.

In the next section of the paper we will estimate our model with the Engle-Granger procedure for a single cointegrating vector.

5. Estimation results: Engle-Granger procedure

The Engle-Granger testing procedure for cointegration consists of four steps: (1) pretest of the variables for their order of integration, (2) estimation of the long-run equilibrium relationship, (3) estimation of the error-correction model, and (4) assessment of the adequacy of the model (see Enders, 1995; Patterson, 2000).

The Engle-Granger procedure assumes that there is only one cointegrating relationship. Let us accept for the moment that this is true and that the wage-setting function is this vector.

In section 4 we have already done the pretests. The gross unemployment benefit replacement rate (*ubr*) is stationary and therefore cannot be part of the cointegrating long-run relationship. We estimate the relationship in the following log-linear form (Please note, as the four variables are jointly determined each of them could be used as the dependent variable.):

$$wcr_t = b_o + b_1 pr_t + b_2 ur_t + b_3 taxt_t + e_t \tag{22}$$

If the variables are cointegrated the OLS-estimates are superconsistent, i.e. they move to their long-run values (population values) very fast. The common trend dominates the regression. In this case, the OLS-estimates are immune to endogeneity, autocorrelation and other forms of misspecification. If the variables are not cointegrated this is only a spurious regression.

We use the estimated residuals $\hat{e}_t$ of the long-run relationship to test whether the variables are cointegrated. An ADF test is performed on the residuals. We do not include an intercept term as the non-stationarity of a residual from a regression is tested. If we can reject the null hypothesis of a unit root in the residuals the variables are cointegrated of order (1,1) and we have found a cointegrating vector. Unfortunately, we cannot use the Dickey-Fuller tables. The residuals $\hat{e}_t$ are the result of a regression and therefore we would too often reject the null hypothesis of a unit root. The correct critical values are computed by Microfit using response surface estimates (see Pesaran and Pesaran, 1997, p. 291).

The results for the long-run relationship and the ADF tests on the residuals can be seen in Table 2.

(Table 2 about here.)

Note that we have restricted our sample from 1965 to 1995. Otherwise it would not have been possible to get a significant cointegrating vector. This shortcoming may be due to a structural break at the end of our sample. We will overcome this problem in section 6.

In the original regression (1) all coefficients have the expected sign and are of reasonable magnitude. The OLS estimate of the standard error of the coefficients is not consistent. Therefore we do not report standard errors in general (see Charemza and Deadman, 1997, p. 133). Standard errors are only reported for the time trend and dummy variables.

An ADF test on the residuals of estimation (1) does not reject the null hypothesis of a unit root. Consequently, we cannot assume that the variables are cointegrated. This result can follow from a number of reasons: a deterministic trend should be included, we have omitted a variable, a structural break is present or the variables are really not cointegrated.

In estimation (2) we include a time trend in our regression, but we are not in the position to reject the null hypothesis.

A natural candidate for an omitted variable is the oil price (*oilp*). In regression (3) we test whether the oil price should be part of the cointegrating vector. We use data on the wholesale price index of mineral oil products. But we still cannot reject that the residuals are non-stationary.

Perhaps, the oil price is not part of the cointegrating vector but influences the relationship in some way. In order to take account of the oil price shocks we have experimented with dummies for specific years, e.g. for 1973, 1979 and 1986. This dummies are very common in the literature, but they proved to be insignificant on the basis of F-Tests.

Furthermore, we assumed that in periods of sharply increasing oil prices the cointegrating relationship is different. We can test this hypothesis by introducing two dummy variables. The dummy *d73_81* takes on values of 1 in the years 1973 to 1981 and the dummy *d82_95* in the years 1982 to 1995. If our assumption is true the dummy *d73_81* should be significant and the dummy *d82_95* should be insignificant. In this case the cointegrating relationship would be the same before and after the period under question.

As can be seen from regression (4) the dummy *d73_81* is highly significant, whereas the other dummy proved to be irrelevant on the basis of an F-Test. With the introduction of these dummy variables our ADF test on the residuals clearly rejects the null hypothesis of a unit root. Consequently, the variables can be regarded as cointegrated.

We end up with regression (5). We include the four variables indicated by theory and the dummy for the period with sharply increasing oil prices.

Finally, we have to notice that the sign on the unemployment rate has changed. A positive coefficient is not in line with our model. In the regressions without dummies the unemployment rate has possibly catched up the effect of the structural break. The positive coefficient could be a sign of misspecification, e.g. there is more than one cointegrating vector. Or we should take account of the influence of long-term unemployment on the wage rate. Unfortunately, time series data for long-term unemployment in Austria is only availabe since 1975.

In the third step we make use of the stationary residuals of the cointegrating long-run relationship to estimate the error-correction model in reduced form. The estimation of this model will inform us about the short-run effects. In addition to the lagged first differences of the cointegrating variables we include also the first differences of the unemployment benefit replacement rate (*ubr*). Each of the jointly determined variables has an own equation. For convenience we only state the equation for $Dwcr$:

$$\begin{aligned} Dwcr_t &= a_1 + a_{wcr}\hat{e}_{t-1} + \sum_{i=1}^{p} a_{11}(i)Dwcr_{t-i} + \sum_{i=1}^{p} a_{12}(i)Dpr_{t-i} \\ &\quad + \sum_{i=1}^{p} a_{13}(i)Dur_{t-i} + \sum_{i=1}^{p} a_{14}(i)Dtaxt_{t-i} + \sum_{i=1}^{p} a_{15}(i)Dubr_{t-i} + e_{wcr(t)} \end{aligned} \tag{23}$$

The $a's$ are short-run parameters and e is a white noise disturbance. We call this model a near VAR. With the exception of the error-correction term $\hat{e}_{t-1}$ the equations are equivalent to a VAR in first differences. a_{wcr} is the speed of adjustment parameter. This parameter tells us in which way *Dwcr* responds to a deviation from the long-run equilibrium. It is known that all procedures used in estimating a VAR can also be applied in the case of a near VAR. Above all, OLS estimates are efficient. Usual test statistics are valid as all variables are stationary.

We estimated the near VAR model with the residuals from regression (5) of Table 2. We report in detail on our estimations for equation (23) in Table 3.

(Table 3 about here.)

In choosing the appropriate order of our VAR we get problems with the small number of observations. It was impossible to make a serious decision on the basis of model selection criteria. Therefore we arbitrarily decided to start with a lag length of 2. In addition, we included a constant as there is a time trend in the data.

Estimation (1) gives quite satisfying results. The residuals are white noise (no autocorrelation, normality and no heteroscedasticity). The speed of adjustment coefficient a_{wcr} has the expected negative sign and is significant at the 5 % level (t-statistic of -2.03 for a one-tail test, p-value: 0.029). The absolute value of a_{wcr} is quite high. This confirms our conclusion that there is a cointegrating relationship.

In order to get a more parsimonious estimation we check whether we can restrict the coefficients on the lagged differences of *ubr* to zero. On the basis of an F-test we cannot reject this restrictions. The variable *ubr* does neither have a short-run nor a long-run influence on the real product wage. Therefore estimation (2) does not consider *ubr* any longer. The results do not change very much. The residuals are still white noise and the coefficient a_{wcr} is nearly unchanged.

The results for the other equations of our near VAR are not so promising. Above all, we could not reject autocorrelation in the residuals for the three remaining equations. This can be due to a too short lag length, to outliers or other forms of misspecification.

As we have only few observations we are not in the position to increase the lag length. But, we tried to introduce dummy variables in our estimation. We experimented with period-specific dummies for oil-price shocks (1973, 1979) and for tax-reforms (1989). But we did not really succeed in improving the results for the remaining equations.

Although our results sime quite convincing at first sight there exist serious problems: We assumed that there is only one cointegrating vector. This assumption contradicts our theoretical NAIRU model with a wage-setting function and a price-setting function. We were only able to get a significant cointegrating vector by restricting our sample from 1965 to 1995. And some equations of the error-correction model do not have white noise residuals. In order to overcome these problems we apply the Johansen procedure for multiple cointegrating vectors in section 6.

6. Estimation results: Johansen procedure

Following the Johansen procedure we make use of maximum likelihood estimators to test for multiple cointegrating vectors (the seminal article is Johansen, 1988; introductory text-book expositions can be found in Enders, 1995, and Patterson, 2000; a recent overview is Johansen, 2000).

We estimate our model within a vector error-correction (VECM) framework. The remaining four non-stationary variables are treated as jointly endogenous. The number of variables is denoted by p. Let use define the vector $z = \{wcr, ur, pr, taxt\}$ and write the 4-dimensional VAR in the following way:

$$z_t = A_1 z_{t-1} + ... + A_k z_{t-k} + m + d_t \qquad (24)$$

m is a constant, d is a white noise residual term and k is the maximum lag length of the VAR. As we have a set of four variables the number of cointegrating relationships r can lie between zero and 3. To estimate our VAR with cointegrating relationships we have to specify it as a VECM:

$$Dz_t = G_1 Dz_{t-1} + ... + G_{k-1} Dz_{t-k+1} + P z_{t-1} + m + d_t \qquad (25)$$

P is a $p \times p$ matrix. The rank r of matrix P is equal to the number of cointegrating vectors. As our endogenous variables are all I(1) the matrix P has to be of reduced rank. We can now test the hypothesis H_1 of reduced rank of P:

$$H_1(r): \; P = ab' \qquad (26)$$

a and b are $p \times r$ matrices of full rank. The b_{ij} are the cointegrating parameters and the a_{ij} are the error correction terms (also called loadings). For illustration we can write down these matrices for an assumed rank of 2:

$$P z_{t-1} = \begin{bmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \\ a_{31} & a_{32} \\ a_{41} & a_{42} \end{bmatrix} \begin{bmatrix} b_{11} & b_{12} & b_{13} & b_{14} \\ b_{21} & b_{22} & b_{23} & b_{24} \end{bmatrix} \begin{bmatrix} wcr \\ pr \\ ur \\ taxt \end{bmatrix}_{t-1} \qquad (27)$$

In order to just identify the long run relationships we need r^2 restrictions. We will see, that our NAIRU model imposes these restrictions on the cointegrating vectors. In addition, we can test overidentifying restrictions.

To implement the Johansen procedure we have to go through the following steps: (1) pretest the variables on their order of integration, (2) determine the order of the underlying VAR, (3) decide whether to include deterministic terms (intercept, trends) in the underlying VAR, (4) estimate the VECM and determine the rank of P, (5) impose the just identifying restrictions, and (6) test overidentifying restrictions.

From section 4 we know that our four variables are I(1). In addition to the non-stationary variables we include an intercept and our dummy *d73_81* in the underlying VAR (see section 5 for a discussion of this dummy variable). We did also experiment with the non-stationary unemployment benefit replacement rate *ubr*. The inclusion of this variable did not change our results significantly, therefore we excluded it in favour of a more parsimonious estimation. The order of the VAR is decided on the basis of model selection criteria. The Akaike information and the Schwarz Bayesian criterion indicate that $k = 1$. Later on, we will check if these lag lengths are long enough to result in white noise residuals.

We know from the unit root tests that with the exception of the unemployment rate our data has deterministic trends in the level of the variables. Therefore we decide to include unrestricted intercepts in the underlying VAR. This specification will give us deterministic trends in the level of the variables if P is rank deficient. Now we can estimate the VECM and determine the rank of P. The Johansen estimates to determine the rank r are given in Table 4:

(Table 4 about here.)

The small sample distribution of the Johansen tests are not known. Reimers (1992) examined the small sample performance of the tests and proposed an adjustment of the statistics. Due to a finite sample bias the unadjusted tests too often indicate cointegration. In order to get rid of this bias the statistics have to be scaled down by $(T - pk)/T$, where T is the number of observations. In our case this adjustment factor is 0.875 and the adjusted statistics are given in Table 4.

The maximum eigenvalue and the trace test indicate that there are two cointegrating vectors. These results also hold in case of the adjusted statistics. A rank of two is consistent with our theoretical model. According to the NAIRU approach there should be two long-run relationships: a wage-setting equation and a price-setting equation.

In order to get these two relationships we have to impose restrictions. As we have two cointegrating vectors we need four just-identifying restrictions. We normalize both vectors on *wcr*. In addition, we use our a-priori and theory-based information that the elasticity of the real product wage on productivity increases is unity and that the tax wedge taxt does not cause higher prices as they are determined on a perfectly competitive product market. Our exactly identifying restrictions are: $b_{11} = -1, b_{12} = 1, b_{21} = -1, b_{24} = 0$. Imposing this restrictions we get the results given in Table 5.

(Table 5 about here.)

The results look quite good, but the standard error of b_{14} is too high to draw sensible conclusions. To get more precise results we impose and test an over-identifying restriction.

We test whether the real product wage does depend on the unemployment rate. According to the NAIRU model there should be a negative relationship. We test this information by imposing the restriction $b_{13} = 0$. Results are given in Table 6.

(Table 6 about here.)

The LR statistic for testing this restriction is 0.66995. According to the c^2 distribution with one degree of freedom this is equal to a p-value of 0.41 and therefore we do not reject the hypothesis that the unemployment rate does not have a long-run influence on the wage-setting equation (i.e. $b_{13} = 0$). As already mentioned in section 5, this result may be due to the influence of long-term unemployment.

The restriction resulted in lower standard errors of the unrestricted coefficients. Let us give a short interpretation of the results. Above all, we are interested in the coefficient on *taxt* in the wage-setting equation. The elasticity of the real product wage on the tax wedge is about 0.6 with a standard error of 0.34. This is significantly different from zero at the 5 % level for a one-tail test (H_0: $b_{14} = 0$, H_1: $b_{14} > 0$).

We do not give a thorough examination of the short-run dynamic properties of the model. But, we have to see if the residuals d are white noise and we shall analyse the speed of adjustment coefficients, i.e. the elements of matrix a. We tested the residuals of the underlying VAR on serial correlation, normality and heteroscedasticity. None of these tests indicated problems with the residuals and therefore we can assume that they are white noise. The results for the speed of adjustment coefficients are given in Table 7.

(Table 7 about here.)

The results on the short-run properties of our model are not as promising as those on the long-run relationships. The loading of the wage-setting vector in the equation of *Dwcr* has the right sign and is significantly different from zero. The absolute value of -0.32 means that it takes about three years that the real product wage has adjusted to its equilibrium value. But, the other coefficients have wrong signs, high standard errors or are much too high in absolute values. We experimented with an underlying VAR of order 2, but in this case also the long-run results deteriorate. A restriction of our data from 1965 to 1995 did not essentially improve the results. We think that the bad short-run dynamic properties of our model are due to the small number of observations. But, as we did not detect problems in the residuals with our diagnostic tests we are still confident about our long-run results.

In the concluding section 7 we will summarize our results and compare them with previous findings. Finally, we will discuss the limitations of our analysis and suggest further research questions.

7. Conclusions

In this paper we investigated the amount of tax shifting in the case of social security contributions and taxes in Austria. We have estimated a right-to-manage model of the labour market. We are especially interested in the effects of the tax wedge (real cost to the employer of hiring a worker - net real wage received by the worker) on the real product wage.

We used two different methods in estimating our model in an error-correction form. At first, we assumed that there is only one cointegrating vector (the wage-setting equation) and we followed the Engle-Granger procedure. The elasticity of the real product wage with respect to the tax wedge is about 0.15 and the speed of adjustment coefficient equals -0.44. These results are not reliable as they suffer from serious problems. In a second step, we estimated our model with the Johansen procedure. We identified two cointegrating vectors (wage-setting and price-setting equation). In this case, the coefficient of the tax wedge in the wage-setting equation is about 0.6 and the relevant speed of adjustment coefficient is -0.32. The results are quite different from those of the Engle-Granger procedure. We rely on the results of the Johansen procedure as they are consistent with our theoretical predictions and we did not detect problems in the residuals. Unfortunately, our results for the short-run properties of the model are not convincing due to the small number of observations.

Our results indicate that 60 % of an increase in the tax wedge is shifted forward on to the real product wage. This is in accordance with Pichelmann (1993) who reported a result of 60 to 70 %. In Hofer and Pichelmann (1996) a lower coefficient of 30 to 40 % was found. As there are differences in the estimation methods and in the data no clear cut conclusions can be drawn. But all these results reject the hypothesis that there is no forward shifting of taxes in the long-run (Nickell and Layard, 2000, 3058ff. give an overview on testing this hypothesis in general).

As already mentioned in the introduction, we would be especially interested whether it makes a difference which side of the labour market is taxed (sometimes called the Invariance of Incidence Proposition). But with our small number of observations and obvious multicollinearity it would be difficult to discriminate between the effects of the instruments. Nonetheless, this is an important research question and has already been discussed in the literature (e.g. Pichelmann, 1993; Tyrväinen, 1995; Hofer and Pichelmann, 1996; Steiner, 1998; Muysken, van Veen and de Regt, 1999). There are several ways to deal with the small number of observations: e.g. imposing restrictions on the coefficients or estimation as a simultaneous-equation model.

One further way is to estimate the model for a panel of OECD countries. If we succeed in pooling long-run parameters we would get more precise estimates than for a single country. In addition, we could possibly be in a position to discriminate between the effects of the instruments as we should have more variance in our observations of the tax wedge and its components. This looks like a promising way to test the Invaricance of Incidence Proposition.

References

Bean, Charles R., Richard Layard, and Stephen J. Nickell (1986) 'The Rise in Unemployment: A Multi-country Study', *Economica*, **53**, S1-S22.

Booth, Alison L. (1995) *The Economics of the Trade Union.* Cambridge: Cambridge University Press.

Charemza, Wojciech W. and Derek F. Deadman (1997) *New Directions in Econometric Practice: General to Specific Modelling, Cointegration and Vector Autoregression.* 2nd ed. Celtenham and Northampton: Edward Elgar.

Daveri, Francesco and Guido Tabellini (2000) 'Unemployment, growth and taxation in industrial countries', *Economic Policy*, **30**, 49-104.

Enders, Walter (1995) *Applied Econometric Time Series.* New York et al.: John Wiley & Sons.

Felderer, Bernhard, Helmut Hofer, and Andreas U. Schuh (1999) 'Entlastungspotentiale bei den Lohnnebenkosten: Vorschläge und Arbeitsmarktimplikationen', Institute for Advanced Studies, Research Report, Vienna.

Franz, Wolfgang (1999a) 'Arbeitslosigkeit in Deutschland und Möglichkeiten ihrer Überwindung: Eine Bestandsaufnahme und Analyse aus der Sicht der wirtschaftspolitischen Beratung', *Empirica*, **26**, 151-167.

Franz, Wolfgang (1999b) *Arbeitsmarktökonomik.* 4th ed. Berlin et al.: Springer.

Fuest, Clemens (2000) *Steuerpolitik und Arbeitslosigkeit.* Tübingen: Mohr Siebeck.

Fuest, Clemens, Bernd Huber, and Florian Wöhlbier (2000) 'Fiscal Policy and Unemployment in a Global Economy,' in Helmut Wagner, ed, *Globalization and Unemployment.* Berlin et al.: Springer.

Hofer, Helmut and Pichelmann, Karl (1996) 'Lohnbildung, Arbeitskosten und Wettbewerbsfähigkeit in Österreich', Institute for Advanced Studies, Research Report, Vienna.

Johansen, Søren (2000) 'Modelling of cointegration in the vector autoregressive model', *Economic modelling*, **17**, 359-373.

Johansen, Søren (1988) 'Statistical Analysis of Cointegrating Vectors', *Journal of Economic Dynamics and Control*, **12**, 231-254.

Knoester, Anthonie and Nico van der Windt (1987) 'Real Wages and Taxation in Ten OECD Countries', *Oxford Bulletin of Economics and Statistics,* **49**, 151-169.

Layard, Richard, Stephen Nickell, and Richard Jackman (1991) *Unemployment: Macroeconomic Performance and the Labour Market.* Oxford: Oxford University Press.

Muysken, Joan, Tom van Veen, and Erik de Regt (1999) 'Does a shift in the tax burden create employment?', *Applied Economics,* **31**, 1195-1205.

Nickell, Stephen and Richard Layard (1999) 'Labor market institutions and economic performance,' in O. Ashenfelter and D. Card, eds, *Handbook of Labor Economics*, *Volume 3.* Amsterdam et al.: North-Holland.

OECD (1990) 'Employer versus Employee Taxation: The Impact on Employment,' *OECD Employment Outlook.* Paris.

OECD (1994) *The OECD Jobs Study, Evidence and Explanations, Part II, The Adjustment Potential of the Labour Market.* Paris: OECD.

ÖSTAT (1990) *Österreichs Volkseinkommen 1989.* Wien: ÖSTAT.

Patterson, Kerry (2000) *An Introduction to Applied Econometrics: A Time Series Approach.* Houndmills and London: Macmillan.

Pesaran, M. Hashem and Bahram Pesaran (1997) *Working with Microfit 4.0: Interactive Econometric Analysis.* Oxford: Oxford University Press.

Pichelmann, Karl (1993) 'Arbeitskosten, Nettoverdienste und das Gewicht der Arbeitslosigkeit im österreichischen Lohnbildungsprozeß', Institute for Advanced Studies, Vienna.

Pissarides, Christopher A. (1998) 'The impact of employment tax cuts on unemployment and wages; The role of unemployment benefits and tax structure', *European Economic Review*, **42**, 155-183.

Reimers, Hans-Eggert (1992) 'Comparisons of tests for multivariate cointegration', *Statistical Papers*, **33**, 335-359.

Sørensen, Peter Birch (1997) 'Public Finance Solutions to the European Unemployment Problem', *Economic Policy*, **25**, 223-251.

Steiner, Viktor (1998) 'Employment and Wage Effects of Social Security Financing - An Empirical Analysis of the West German Experience and some Policy Simulations,' in John T. Addison and Paul J. J. Welfens, eds, *Labor Markets and Social Security - Wage Costs, Social Security Financing and Labor Market Reforms in Europe.* Berlin et al.: Springer.

Tyrväinen, Timo (1995) 'Real Wage Resistance and Unemployment: Multivariate Analysis of Cointegrating Relations in 10 OECD Countries', The OECD Jobs Study, Working Paper Series, No. 10, Paris.

Appendix A: Variable definitions and sources

Variable	Definition	Source
cpi	consumer price index (VPI I), 1958 = 100	WIFO database
oilp	wholesale price index of mineral oil products, 1976 = 100	WIFO database and own calculations
pgdp	GDP deflator, 1983 = 100	WIFO database and own calculations
pr	labour productivity = real GDP(1983 = 100)/ employed (employees and self-employed)	WIFO database and own calculations
taxt	overall tax wedge = $\ln\left(\frac{WC/PGDP}{WN/CPI}\right)$	own calculations
ubr	gross unemployment benefit replacement rate (average)	OECD and own calculations
ur	unemployment rate, registered unemployed as a percentage of employees (employed and unemployed)	WIFO database
w	gross wages and salaries per employee (nominal, per month, in ATS)	Östereichs Volkseinkommen and own calculations
wc	compensation per employee (nominal, per month, in ATS)	Östereichs Volkseinkommen and own calculations
wn	net wages and salaries per employee (nominal, per month, in ATS)	Östereichs Volkseinkommen and own calculations
wcr	real product wage = ln(WC/PGDP)	own calculations
wnr	real consumption wage = ln(WN/CPI)	own calculations

Notes: Lower-case letters indicate variables in logarithms. Upper-case letters indicate variables in non-logarithmic form.

Appendix B: Calculation of net wages and salaries

Compensation of employees (*WC*)

\- Employers' contributions for social security

\- Imputed contributions of government and financial integrated or public enterprises

\- Employers' contribution for private pension

= Gross wages and salaries (*W*)

\- Personal income tax (Labour income tax)

\- Employees' contributions for social security

\- Contributions to the chamber of labour (compulsory)

= Net wages and salaries (*WN*)

Source: ÖSTAT (1990, p. 73).

Table 1: Unit root tests

Variable	x	d(x)	I(x)
oilp	-1.44 [t,1]	-3.51 [n,0]***	I(1)
pr	-3.06 [t,0]	-4.37 [c,0]***	I(1)
taxt	-1.80 [t,0]	-4.04 [n,0]***	I(1)
ubr	-3.69 [t,1]**	-	I(0)
ur	-0.75 [c,3]	-2.00 [n,2]**	I(1)
wcr	-0.48 [t,2]	-2.24 [n,1]**	I(1)

Notes: All variables in logarithms. Sample from 1965 to 1997.
[z,#]: z = [n,c,t]
n.......neither constant nor trend included
c.......constant included
t........trend and constant included
#.......number of lags included
*.......significant rejection of unit root hypothesis at 10 % level
**......significant rejection of unit root hypothesis at 5 % level
***.....significant rejection of unit root hypothesis at 1 % level

Table 2: Long-run relationship

	(1)	(2)	(3)	(4)	(5)
dependent variable	*wcr*	*wcr*	*wcr*	*wcr*	*wcr*
constant	-7.092	-4.743	-6.751	-6.278	6.147
pr	0.927	0.737	0.910	0.873	0.864
ur	-0.029	-0.059	-0.018	0.058	0.054
taxt	0.307	0.318	0.155	0.154	0.148
oilp	-	-	0.022	-	-
time trend	-	0.006(0.0032)	-	-	-
d73_81	-	-	-	0.062(0.0128)	0.065 (0.0115)
d82_95	-	-	-	-0.010 (0.0177)	-
ADF test on residuals	-3.6769	-3.9931	-3.5415	-6.6559	-6.6183
order of ADF test	0	0	0	1	1
95 % critical value	-4.4821	-4.9345	-4.8967	-5.3084	-4.9140

Notes: Sample from 1965 to 1995. OLS. Standard errors are only reported for the time trend and dummy variables in parentheses.

Table 3: Error-correction model

	(1)	(2)
dependent variable	*Dwcr*	*Dwcr*
Dwcr(-1)	0.032 (0.177)	0.045 (0.168)
Dwcr(-2)	0.508 (0.180)	0.523 (0.165)
Dpr(-1)	0.086 (0.163)	0.089 (0.146)
Dpr(-2)	0.015 (0.116)	0.024 (0.110)
Dur(-1)	0.016 (0.020)	0.015 (0.019)
Dur(-2)	-0.042 (0.020)	-0.040 (0.019)
Dtaxt(-1)	-0.035 (0.126)	-0.020 (0.116)
Dtaxt(-2)	0.006 (0.124)	0.019 (0.113)
Dubr(-1)	0.018 (0.025)	-
Dubr(-2)	-0.006 (0.025)	-
constant	0.008 (0.005)	0.007 (0.005)
$\hat{e}^{-1}$	-0.422 (0.208)	-0.443 (0.196)

Notes: Sample from 1965 to 1995. OLS. Standard errors in parantheses.

Table 4: Johansen estimates

Null	Alternative	Statistic	Adjusted Statistic	95% Critical value
Cointegration LR test based on maximal eigenvalue of the stochastic matrix				
$r = 0$	$r = 1$	59.0653	51.6821	27.4200
$r \leq 1$	$r = 2$	28.8587	25.2514	21.1200
$r \leq 2$	$r = 3$	8.5985	7.5237	14.8800
$r \leq 3$	$r = 4$	1.1496	1.0059	8.0700
Cointegration LR test based on trace of the stochastic matrix				
$r = 0$	$r \geq 1$	97.6721	85.4631	48.8800
$r \leq 1$	$r \geq 2$	38.6068	33.7810	31.5400
$r \leq 2$	$r \geq 3$	9.7481	8.5210	17.8600
$r \leq 3$	$r = 4$	1.1496	1.0059	8.0700

Notes: 32 observations from 1966 to 1997. Cointegration with unrestricted intercepts and no trends in the underlying VAR. Order of VAR = 1. List of variables included in the cointegrating vector: *wcr, pr, ur, taxt*. List of I(0) variables included in the VAR: *d73_81*. List of eigenvalues in descending order: 0.84210; 0.59417; 0.23563; 0.035289.

Table 5: Exactly identified cointegrating vectors

	Vector 1	Vector 2
wcr	-1.0000 (none)	-1.0000 (none)
pr	1.0000 (none)	0.9386 (0.0920)
ur	0.0408 (0.0501)	0.0720 (0.0322)
taxt	0.2542 (0.4608)	0.0000 (none)

Note: Standard errors in parantheses.

Table 6: Over identified cointegrating vectors

	Vector 1	Vector 2
wcr	-1.0000 (none)	-1.0000 (none)
pr	1.0000 (none)	0.8961 (0.0529)
ur	0.0000 (none)	0.0719 (0.0288)
taxt	0.5958 (0.3397)	0.0000 (none)

Notes: Standard errors in parantheses. LR test of over-identifying restriction: $c_1^2 = 0.66995$ (p-value: 0.413).

Table 7: Loadings in underlying VAR

independent variable	wage-setting equation	price-setting equation
Dwcr	-0.3179 (0.0419)	0.3520 (0.0895)
Dpr	-0.3042 (0.0549)	0.0632 (0.1172)
Dur	1.2379 (0.4385)	-4.022 (0.9371)
Dtaxt	0.0523 (0.0612)	-0.0746 (0.1309)

Note: Standard errors in parantheses.

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