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NONLINEAR WAGE RESPONSES TO INTERNAL AND EXTERNAL FACTORS

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Nonlinear wage responses to internal and external factors*

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Abstract

The paper tests whether or not the effects on sectoral wages of internal and external factors depend upon the sector's relative wage position. The key hypothesis is that workers in low–wage sectors are more concerned with relative wages than workers in high wage sectors. To test the hypothesis, we make use of panel data and formulate a smooth transition regression model including relative wages as the transition variable. The empirical results provide strong evidence of nonlinear wage responses to industry profitability, outside wages and unemployment. The estimated long–run insider weight and the unemployment effect are much higher in high–wage industries than in low–wage industries. The main results are robust to alternative transformations of the unemployment rate and we also provide some evidence of nonlinear effects using regional panel data.

Keywords: Insider forces, panel data, nonlinear modelling.

JEL classification: C23, J31.

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1 Introduction

The paper investigates the importance of insider forces in the Norwegian wage formation process using industry and regional panel data. A key issue will be to test whether or not the effect of sector specific variables depend upon the sector's relative wage position. Our hypothesis is that low–wage sectors mainly act as wage followers while high–wage sectors are wage leaders. We therefore suggest that insider forces are more important in high–wage sectors than in low–wage sectors where the outside wage is hypothesized to play a dominant role. To investigate the possibility of such a nonlinear wage responses to internal and external factors, we formulate a smooth transition regression (STR) model, following the approach suggested by Granger and Teräsvirta (1993) and Teräsvirta (1998).

The basic idea is that pay is shaped by a mixture of internal and external factors. This feature is common to several models where wage setting can be regarded as a form of rent–sharing such as union bargaining models (Dunlop, 1944, Hoel and Nymoen, 1988, Nickell and Andrews, 1983), insider–outsider models (Lindbeck and Snower, 1988), and some versions of efficiency wage theories (Shapiro and Stiglitz, 1984, Agell and Lundborg, 1995a).

The presence of a permanent relation between industry wages and industry profitability is evidence against competitive forces in the labour market, but also evidence against completely centralized wage setting. In a competitive labour market, wages would be equalized, except for compensating wage differentials due to differences in workers' skill and human capital, or in working conditions. In a completely centralized wage setting system, wages are fully determined through nation–wide bargaining, and wage differentials reflect the preferences and bargaining power of the central labour market organizations.

Norway is often considered an economy with highly centralized and co–ordinated wage setting. It may therefore be argued that firm or industry specific factors are of limited importance in shaping firm or industry specific wages. However, wage bargaining in Norway takes place at both the national and the firm level. Within manufacturing, wage drift has contributed to more than 50% of total wage increases during the post war period. This implies considerable scope for firm or industry specific rent sharing.

The empirical evidence on insider forces in Norway is mixed. Holmlund and Zetterberg (1991) report small and statistically insignificant effects of industry prices and productivity upon industry wages while the estimated long–run insider weight reported in Johansen (1996, 99) approximates 20%.[1]

[1] The long–run insider weight is defined as the long–run elasticity of firm or industry

Using panel data for Norwegian manufacturing firms, Wulfsberg (1997) reports a statistically significant insider weight equal to 5% while the results in Raaum and Wulfsberg (1998) imply a long–run insider weight slightly above 30%.

The main part of the present paper makes use of the industry panel data used in Johansen (1999). A particular issue will be to test the hypothesis that industry wages respond most strongly to deviations from the outside wage when the industry wage is below the outside wage. According to Akerlof and Yellen (1990) wage comparison effects are more important for low–income workers than for high–income groups. Both survey evidence and econometric results are supportive to this hypothesis.

The rest of the paper is organized as follows. Section 2 discusses wage setting theories explaining a positive relation between industry wages and industry performance. Section 3 discusses empirical specification of the benchmark linear wage equation and formulate the nonlinear alternative. Some econometric issues are also discussed. Results for the linear model are presented in Section 4, while tests of linearity against the nonlinear alternative can be found in Section 5. Section 6 contains results for a logistic smooth transition regression (LSTR) model. In Section 7 we test the robustness of the results. In particular, some results based on *regional* panel data are reported. Section 8 concludes.

2 Theoretical background

The theoretical foundation for the empirical analysis is a firm level bargaining model similar to the model presented in Nickell and Wadhwani (1990). It is assumed that wages are determined by bargaining between the union and the firm. Prices and employment are determined unilaterally by the firm, after the wage bargain, in order to maximize profits. The firm faces a downward sloping demand schedule that is affected by random shifts. Because wages are determined prior to the revelation of actual demand, employment, prices and profits will be unknown when bargaining takes place. These variables are therefore considered as expected magnitudes.

We assume that the union is concerned with the welfare of the representative worker employed by the firm when bargaining takes place (the insider worker). We further assume that the utility of an individual insider worker increases with the real consumer wage but also with his wage *relative* to what other workers are paid.

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level wages with respect to firm or industry level prices.

If layoffs are made by random draw, the probability for the insider to stay with the firm during the next contract period decreases with the real product wage. A laid–off insider will be unemployed for a period receiving unemployment benefits. Thereafter, the individual will obtain a new job and receives the alternative wage. It is assumed that the utility of an employed worker is strictly higher than the utility of an unemployed. Further, a higher unemployment rate will increase the expected time spent unemployed. Given these assumptions, the expected utility for an insider, conditional on becoming unemployed, increases with the expected alternative wage and the benefit level, and decreases with the unemployment rate and the comparison wage.

Finally, we assume that the outcome of the local wage bargain is given by the solution of an asymmetric Nash bargaining problem. Under these assumptions we can derive the following approximation to the static nominal wage equation

$$\begin{aligned} w_{it} &= \mu_1 va_{it} + \mu_2 \left[wa_{it} + c_1 b_t + f\left(U_t\right)\right] \\ &\quad + \left(1 - \mu_1 - \mu_2\right)\left[pc_t + ti_t + tp_t\right] + \mu_0, \end{aligned} \tag{1}$$

where w_{it} is wage costs per hour worked, va_{it} is value added per hour worked, wa_t is the outside wage, b_t is the unemployment benefit replacement ratio, U_t is the unemployment rate, pc_t is the consumer price index, ti_t is the income tax rate and tp_t the payroll tax rate. Lower case letters denote log transformed variables, subscript i is used for industry and t for time period.[2]

The nominal wage equation is homogenous of degree one in the nominal explanatory variables. The insider weight, μ_1, is the main parameter of interest, and is interpreted as the long–run elasticity of industry wages with respect to industry prices and productivity. The outside wage represents the effects of both the alternative wage and the comparison wage. The final term in (1) represents a wedge effect. However, the wedge drops out if the workers and the firm are risk neutral which means that wage costs are unaffected by taxes and consumer prices, and higher tax rates are entirely borne by labour.

Above we assumed that workers are concerned with relative wages. Arguments in favour of the role of fair wage considerations and social norms in wage determination can be found in Marshall (1925), Hicks (1974), Oswald (1979), Gylfason and Lindbeck (1984), Solow (1990) and in Akerlof and Yellen (1990). Clark and Oswald (1994) give a detailed overview on theoretical models and provide econometric evidence supporting the view that relative income matters for job satisfaction. Survey evidence supporting the

[2] See appendix A.1 for definition of the variables and sources. A more detailed description of the main variables including an investigation of the temporal properties can be found in Johansen (1999).

hypothesis of interpersonal wage comparison effects can be found in Kaufman (1984), Blinder and Choi (1990) and in Agell and Lundborg (1995b). Akerlof and Yellen (1990) argue that the comparison wage effect may be asymmetric. Low–income workers are more concerned with relative wages than high-income workers, or "Those people who receive less are of comparatively little interest...; whereas those people who are paid more are of considerable interest..." (Akerlof and Yellen, 1990, p. 259). Survey evidence from Sweden, reported in Agell and Lundborg (1995b), support this hypothesis. Econometric results in Falch (1993) and Strøm (1995) also suggest that comparison effects are more important for low–paid than for high–paid workers in the Norwegian local public sector.

In this paper, the possibility of an asymmetric comparison wage effect motivates the nonlinear dynamic wage equation specified in the next section. Another argument is solidarity wage policy.

3 Empirical specification and econometric issues

3.1 Empirical specification

The replacement ratio is excluded from the empirical model, partly because previous Norwegian studies report insignificant effects[3], and partly to reduce the number of pure time series variables. Higher unemployment reduces wages by increasing the expected costs of being laid off. Nickell (1987) and Blanchflower and Oswald (1990) provide arguments in favour of a nonlinear wage curve, and the results in Johansen (1995a, 96, 99) imply that the Norwegian wage curve is highly convex. We therefore assume an inverted square specification corresponding to the preferred equation reported in Johansen (1995a, 96, 99). Finally, expectation inertia and long–term contracts are arguments for rather complicated dynamics. As a benchmark for further analysis, we utilize the results in Johansen (1999) and formulate the log–linear equilibrium correction model given by

$$\begin{aligned}
\Delta w_{it} &= \beta_{01}\left(w-wa\right)_{it-1}+\beta_{02}\left(va-wa\right)_{it-1}+\beta_{03}U_{t-1}^{-2}+\beta_{04}\Delta va_{it}+ \\
&\quad \beta_{05}\Delta va_{it-1}+\beta_{06}\Delta wa_{it}+\beta_{07}\Delta^{2}tp_{t}+\beta_{08}\Delta h_{t}+\beta_{09}\Delta U_{t}^{-2}+ \\
&\quad \beta_{010}\Delta U_{t-1}^{-2}+\beta_{011}\left(\Delta pc+\Delta ti\right)_{t-1}+\alpha_{i}+\varepsilon_{it} \\
&\equiv \boldsymbol{\beta}_{0}\mathbf{x}_{it}+\alpha_{i}+\varepsilon_{it}. \qquad (2)
\end{aligned}$$

[3] An exception is Holden and Nymoen (1998) who report a significant effect of the replacement ratio using time series data for Norway 1964–94.

Equation (2) includes industry fixed effects, α_i, to control for wage differentials due to time–invariant unobservable factors, while changes in normal working time, Δh_t, is included to control for compensation effects of reduced working time.[4] Finally, ε_{it} are the error terms, assumed to be $iid\,(0, \sigma_\varepsilon^2)$.

Turning to the nonlinear specification, our main issue will be to test whether or not industry wage responses to internal and external variables depends upon the industry's relative wage position. Given this particular issue, we make the feedback coefficients of the level variables in (2) functions of lagged levels of relative wages. The nonlinear equilibrium correction model is then given by

$$\Delta w_{it} = \boldsymbol{\beta}_0 \mathbf{x}_{it} + \boldsymbol{\theta} \widetilde{\mathbf{x}}_{it-1} G\left(\gamma, c; s_{it-1}\right) + \alpha_i + u_{it}, \tag{3}$$

where $\boldsymbol{\beta}_0 \mathbf{x}_{it}$ is defined in equation (2), $\boldsymbol{\theta} = (\theta_1,\ \theta_2,\ \theta_3)$, while $\widetilde{\mathbf{x}}_{it-1}$ is a subset of the right hand side variables in (2), defined by

$$\widetilde{\mathbf{x}}_{it-1} = \left[(w - wa)_{it-1},\ \ (va - wa)_{it-1},\ U_{t-1}^{-2}\right]^T. \tag{4}$$

$G\left(\gamma, c; s_{it-1}\right)$ is a continuous transition function, bounded between zero and unity, and $s_{it-1} = W_{it-1}/WA_{it-1}$ is the transition variable.

To estimate equation (3) we must further specify the transition function. The alternative estimated in Section 6 is the logistic smooth transition (LSTR1) function given by

$$G_1\left(\gamma, c; s_{it-1}\right) = \left[1 + \exp\left\{-\gamma\left(s_{it-1} - c\right)\right\}\right]^{-1}, \qquad \gamma > 0. \tag{5}$$

The transition function (5) is monotonically increasing in the transition variable, s_{it-1}. Further,

$$\lim_{s_{it-1} \to \infty} G_1\left(\gamma, c; s_{it-1}\right) = 1 \text{ and } \lim_{s_{it-1} \to -\infty} G_1\left(\gamma, c; s_{it-1}\right) = 0$$

for given values of γ and c. The slope parameter γ indicates how rapid the transition from zero to unity is as a function of s_{it-1}. The location parameter c determines where the transition occurs. If $\gamma \to \infty$ (3) becomes a two–regime switching regression where $s_{it-1} = c$ is the switchpoint between the regimes $\Delta w_{it} = \boldsymbol{\beta}_0 \mathbf{x}_{it} + \alpha_i + u_{it}$ and $\Delta w_{it} = \boldsymbol{\beta}_0 \mathbf{x}_{it} + \boldsymbol{\theta} \widetilde{\mathbf{x}}_{it-1} + \alpha_i + u_{it}$.

[4] The results in Johansen (1999) suggest that the log of relative wages and the log of the wage share are both stationary variables, while results in Bjørnland (1995) and Johansen (1995) indicate that the unemployment rate is $I\,(0)$, but with possible structural breaks. Assuming that the differenced variables are $I(0)$, equation (2) is balanced. The lagged wedge is excluded from equation (2) because the results in Johansen (1999) imply no long-run wedge effect. We impose equal coefficients of Δpc_{t-1} and Δti_{t-1}, a restriction that is easily accepted by data.

A monotonic transition function like (5) may not always be satisfactory. A nonmonotonic alternative to the LSTR1 function is the LSTR2 function given by

$$G_2(\gamma, c_1, c_2; s_{it-1}) = \left[1 + \exp\left\{-\gamma\left(s_{it-1} - c_1\right)\left(s_{it-1} - c_2\right)\right\}\right]^{-1}, \qquad (6)$$
$$\gamma > 0,\ c_1 \leq c_2,$$

which is symmetric about $(c_1 + c_2)/2$, and $G_2 \rightarrow 1$ for $s_{it-1} \rightarrow \pm\infty$ while the minimum value of G_2 lies between 0 and 1/2. Another nonmonotonic alternative is the exponential STR (ESTR) model defined by

$$G_3(\gamma, c, s_{it-1}) = 1 - \exp\left\{-\gamma\left(s_{it-1} - c\right)^2\right\}, \quad \gamma > 0, \qquad (7)$$

which is symmetric about c, $G_3 = 0$ for $s_{it-1} = c$, and $G_3 \rightarrow 1$ for $s_{it-1} \rightarrow \pm\infty$.[5] In section 5 we use the results for linearity testing to make a choice between these alternatives.

3.2 Econometric issues

Since the fixed effects model (2) includes lagged values of the left hand side variable, the within groups estimates are biased even if the error terms are white noise, cf. Nickell (1981). Further, because we assume that firms set prices and employment, va_{it} and va_{it-1} will be treated as endogenously determined. We also consider current values of all aggregate time series variables as expected magnitudes and therefore as potentially endogenous variables. To obtain consistent estimators we make use of the Generalized Method of Moment (GMM) estimator suggested by Arellano and Bond (1991). The model is first–differenced to remove the individual fixed effects due to potential correlation with $\mathbf{x}_{it}$. In the absence of second–order serial correlation in the transformed residuals $(\Delta\varepsilon_{it})$, w_{it-r} and va_{it-r} are valid instruments for $r \geq 2$. The set of instruments used in the estimation utilize the orthogonality restrictions between the differenced residuals and the second and the third lag of industry wages and factor income per hour. Further instruments are lagged values of the aggregate time series variables in equation (2).

More technically, the one step GMM, GMM(1), estimators of the unknown parameters $\boldsymbol{\beta}_0$ are found by minimizing the quadratic distance

$$Q\left(\widehat{\boldsymbol{\beta}}_0\right) = \left(\left(\Delta\varepsilon\right)^T Z A_N Z^T \left(\Delta\varepsilon\right)\right), \qquad (8)$$

[5]See Teräsvirta (1998, p. 511) for further details concerning the LSTR2 and ESTR models.

where $\Delta\varepsilon = \left((\Delta\varepsilon_1)^T, (\Delta\varepsilon_2)^T, .., (\Delta\varepsilon_{116})^T\right)$, $\Delta\varepsilon_i = (\Delta\varepsilon_{i68}, \Delta\varepsilon_{i69}, .., \Delta\varepsilon_{i91})^T$, Z is the matrix of instruments described above while A_N is the weighting matrix given by

$$A_N = \left(N^{-1}\sum_{i=1}^{116} Z_i^T H Z_i\right)^{-1}, \tag{9}$$

where H is a 24×24 matrix which has twos in the main diagonal, minus one in the first subdiagonals and zeros otherwise.[6]

Testing linearity against the alternative of a STR model amounts to testing the null hypothesis that $\gamma = 0$. However, the model is not identified under the null due to the nuisance parameters θ and c. Following Teräsvirta (1998) we use a Taylor series approximation about $\gamma = 0$ as a substitute to circumvent this problem. The test of linearity is based on the auxiliary regression

$$\Delta w_{it} = \boldsymbol{\beta}_0\mathbf{x}_{it} + \boldsymbol{\beta}_1\widetilde{\mathbf{x}}_{it-1}s_{it-1} + \boldsymbol{\beta}_2\widetilde{\mathbf{x}}_{it-1}s_{it-1}^2 + \boldsymbol{\beta}_3\widetilde{\mathbf{x}}_{it-1}s_{it-1}^3 + \alpha_i + v_{it}, \tag{10}$$

where $\boldsymbol{\beta}_0\mathbf{x}_{it}$ is defined by equation (2), $\boldsymbol{\beta}_i = (\beta_{i1},\ \beta_{i2},\ \beta_{i3})$, $i = 1, 2, 3$, $\widetilde{\mathbf{x}}_{it-1}$ is defined by (4) and the transition variable is given by $s_{it-1} = W_{it-1}/WA_{it-1}$. Equation (10) is estimated using the same procedure as described above with the exception that the second lag of the interaction terms are used as additional instruments. The null hypothesis of linearity i.e.:

$$\mathrm{H}_0\text{:}\quad \boldsymbol{\beta}_1 = \boldsymbol{\beta}_2 = \boldsymbol{\beta}_3 = 0,$$

against the alternative H_1: "at least one $\boldsymbol{\beta}_i \neq 0$" is tested using the Wald χ^2 test implemented in DPD by Arellano and Bond (1988).

The following sequence of null hypotheses within (10) is used to choose between alternative specifications of the transition function:

$$\begin{aligned}
\mathrm{H}_{04} &: \quad \boldsymbol{\beta}_3 = 0 \\
\mathrm{H}_{03} &: \quad \boldsymbol{\beta}_2 = 0 \mid \boldsymbol{\beta}_3 = 0 \\
\mathrm{H}_{02} &: \quad \boldsymbol{\beta}_1 = 0 \mid \boldsymbol{\beta}_2 = \boldsymbol{\beta}_3 = 0.
\end{aligned}$$

Following Teräsvirta (1998, p. 527) we apply the decision rule: "If the rejection of H_{03} is the strongest one, choose an LSTR2 (or an ESTR) model, otherwise select an LSTR1 model".

[6] See Arellano and Bond (1991) or Baltagi (1995) for details. The GMM estimator is implemented in DPD written in GAUSS, cf. Arellano and Bond (1988). We use the one-step estimator since the standard errors generated by the (asymptotically more efficient) two–step procedure are downward biased in finite samples.

The parameters in (3) with (4) and (5) are estimated by making use of a two–dimensional grid search. Giving fixed values to the parameters in the transition function (γ and c) makes (3) linear in the remaining parameters, $\boldsymbol{\beta}_0$ and $\boldsymbol{\theta}$. These parameters are estimated using GMM(1) for the given combinations of γ and c. The parameters in the transition function are chosen so as to minimize the objective function

$$Q\left(\widehat{\beta}_0, \widehat{\theta}, \widehat{\gamma}, \widehat{c}\right) = \left((\Delta u)^T Z A_N Z^T (\Delta u)\right), \tag{11}$$

where $\Delta u = \left((\Delta u_1)^T, (\Delta u_2)^T, .., (\Delta u_{116})^T\right)$, $\Delta u_i = (\Delta u_{i68}, \Delta u_{i69}, .., \Delta u_{i91})^T$, Z is the matrix of instruments which now includes the second lag of the non-linear terms, $\widetilde{x}_{it-2} G(\gamma, c; s_{it-2})$, while A_N is defined by equation (9).

4 Empirical results: The linear model

Empirical results for the linear model are given by equation (12), which reports GMM(1) estimates robust to heteroscedastisity (t–statistics in parentheses below the estimates).

$$\begin{aligned}
\widehat{\Delta w_{it}} &= \underset{(4.85)}{-0.211}(w-wa)_{it-1} + \underset{(4.55)}{0.037}(va-wa)_{it-1} + \underset{(6.12)}{0.009}U_{t-1}^{-2} \\
&\quad + \underset{(2.66)}{0.028}\Delta va_{it} - \underset{(2.01)}{0.006}\Delta va_{it-1} + \underset{(42.33)}{0.855}\Delta wa_{it} + \underset{(8.83)}{0.466}\Delta^2 tp_t \\
&\quad - \underset{(7.07)}{0.213}\Delta h_t + \underset{(2.99)}{0.0029}\Delta U_t^{-2} + \underset{(3.76)}{0.0047}\Delta U_{t-1}^{-2} \\
&\quad + \underset{(2.98)}{0.076}(\Delta pc + \Delta ti)_{t-1} \qquad (12) \\
T &= 24\ [1968-1991],\ N = 116,\ \text{Method}: GMM(1),\ Q = 0.1893, \\
\widehat{\sigma}\,(\%) &= 1.933, \quad m1 = -4.962, \quad m2 = -0.508,\ \chi^2_{\text{SAR}}[98] = 110.73.
\end{aligned}$$

The estimate of the equilibrium correction term, $(w - wa)_{it-1}$, is small but highly significant, and so are also the feedback effects of value added per worker and aggregate unemployment. Equation (12) implies a long–run insider weight equal to 0.178 with a t–value of 3.38, while the long–run unemployment coefficient is 0.04 (t–value = 4.03). The estimated insider weight is close to the corresponding estimate reported in Johansen (1996, 99).[7,8]

The estimated short–run insider effect is also well determined, but very small. The dominating determinant of industry wages is the outside wage. The short–run impact elasticity is 0.85 while the long–run elasticity approximates 0.8. These effects may reflect a mixture of alternative, comparison wage mechanisms or efficiency wage mechanisms. Most important, the estimate of wa may also reflect an effect of the centrally negotiated wage upon the final outcome for wages, after the local bargain.

Equation (12) contains statistically and numerically significant short–run effects of payroll taxes and normal working time while the estimates of ΔU_t^{-2}, ΔU_{t-1}^{-2} and $(\Delta pc + \Delta ti)_{t-1}$are all small but statistically significant from zero. [9]

Equation (12) reports several diagnostic test statistics. First, the Arellano and Bond (1991) $m2$ statistic, testing the null of no second–order serial correlation in the differenced residuals, is insignificant from zero. On the other hand, the $m1$ statistic indicate negative first–order serial correlation in the transformed residuals.[10] Taken together these results imply that the levels of the error terms are white noise. Moreover, the Sargan (1958) test for instrumental validity, $\chi^2_{\mathrm{SAR}}(98)$, is well below it's 5% critical value. Finally, Johansen (1999) provide evidence of parameter stability both over time and across industry. Nevertheless, we now turn to testing equation (12) against a general nonlinear alternative.

5 Testing linearity

When we estimated the unconstrained version of equation (10), the estimates of β_{i1}, $i = 1, 2, 3$, were all numerically small and statistically insignificant from zero. A joint test of the null that $\beta_{11} = \beta_{21} = \beta_{31} = 0$ yield $\chi^2(3) = 2.12$

[7] Johansen (1999) uses data for 117 industries for the same time period. In the present study, one industry is excluded due to very low wages during the whole sample period. The full sample estimates for the linear model are almost identical to those reported in equation (12). For the nonlinear alternative, the exact parameter estimates were somewhat affected when we included the "outlier" sector. However, all main results reported below remain also for the full sample.

[8] One should be aware that the insider weight may be overstated because of the absence of any control for the skill mix of the workforce.

[9] It should be noted that the estimated t–values of the coefficients representing aggregate effects are biased upwards due to common group effects, cf. Moulton (1986). However, most of the t–values are so high that they are likely to remain statistically significant even after any correction for common group effects.

[10] The tests for serial correlation, $m1$ and $m2$, are asymptotic normal, cf. Arellano and Bond (1991).

which is clearly insignificant (p–value $= 0.55$). The economic interpretation of this result is that the overall speed-of-adjustment of industry wages does not depend on the industry's relative wage position. We therefore impose the restrictions $\beta_{i1} = 0$, $i = 1, 2, 3$, and estimate equation (10) excluding the equilibrium correction term from the nonlinear part of the equation.

Results for testing linearity against the simplified version of equation (10) are reported in Table 1. We first see that the null of linearity is firmly rejected. The null hypothesis of no third order term can not be rejected while the null that $\boldsymbol{\beta}_2 = 0$, conditional on $\boldsymbol{\beta}_3 = 0$, is rejected at a significance level of 13%. However, the rejection of H_{02} is clearly the strongest one.

The last two rows in Table 1 report results testing the separate hypotheses that $(va - wa)_{it-1}$ and, respectively U_{t-1}^{-2} do not enter the nonlinear part of equation (10). For both these variables the null of linearity is firmly rejected against the alternative.

Table 1

6 Results for the LSTR1 model

Based on the results in Section 5 we formulate an LSTR1 model where the coefficients of $(va - wa)_{it-1}$ and U_{t-1}^{-2} are allowed to change smoothly with lagged values of relative wages. Empirical results for the LSTR1 model is given by equation (13) which reports GMM(1) estimates, robust to heteroscedasticity (t–statistics in parentheses).

$$\begin{aligned} \widehat{\Delta w_{it}} = & \left[\underset{(3.09)}{0.028}\,(va - wa)_{it-1} + \underset{(3.65)}{0.0108}U_{t-1}^{-2}\right]\{1 + \exp\left(-43.5\left(s_{it-1} - 1.057\right)\right)\}^{-1} \\ & - \underset{(5.87)}{0.318}\,(w - wa)_{it-1} + \underset{(2.33)}{0.022}\,(va - wa)_{it-1} + \underset{(2.66)}{0.0053}U_{t-1}^{-2} \\ & + \underset{(2.55)}{0.027}\Delta va_{it} - \underset{(1.61)}{0.004}\Delta va_{it-1} + \underset{(40.71)}{0.854}\Delta wa_{it} + \underset{(9.14)}{0.471}\Delta^2 tp_t \\ & - \underset{(7.22)}{0.211}\Delta h_t + \underset{(2.77)}{0.0030}\Delta U_{t-1}^{-2} + \underset{(3.50)}{0.0046}\Delta U_{t-1}^{-2} + \underset{(3.50)}{0.078}\,(\Delta pc + \Delta ti)_{t-1} \end{aligned} \tag{13}$$

$T = 24$ $[1968 - 1991]$, $N = 116$, Method: $GMM(1)$, $Q = 0.1741$,
$\widehat{\sigma}\,(\%) = 1.888$, $m1 = -4.894$, $m2 = -0.308$, $\chi^2_{\mathrm{SAR}}\,(102) = 112.86$.

From equation (13) we first note that $(va - wa)_{it-1}$ and U_{t-1}^{-2} both enter the nonlinear part of the model with statistically significant estimates.[11] Turning

[11] We also estimated a version containing the error correction term in the nonlinear part of the model. However, we could not reject the hypothesis that the speed of adjustment

to the transition function we see that the estimate of the location parameter, c, is close to unity. This means that the transition function is symmetric about a relative wage level that approximates the sample mean. The estimate of the slope parameter, γ, is rather high which implies a very steep transition function as shown in Figure 1. The value of the transition function is 0.08 for $W_{it-1}/WA_{it-1} = 1.0$ and 0.98 for $W_{it-1}/WA_{it-1} = 1.15$.[12]

Figure 1

The estimates of $(va - wa)_{it-1}$ and U_{t-1}^{-2} in the linear part of equation (13) are both smaller than the estimates in equation (12). Using the estimate of the equilibrium correction term we find that the estimated long–run insider weight is only 0.068 in the low–wage regime while the long–run unemployment coefficient is 0.017. The outside wage completely dominate the long–run path with a long–run elasticity of 0.932. The insider weight approximates 0.075 at the sample mean for relative wages, and increases to 0.158 in the high–wage regime, an estimate that is still below the estimated insider weight calculated from the linear model. The long–run unemployment coefficient approximates 0.019 at the sample mean for relative wages, which is well below the long–run estimate based on the linear model, and increases to 0.051 in the high–wage regime.

The estimated long–run insider weight is graphed against relative wages in Figure 2, while Figures 3 and 4 graph the estimated outsider weight and the long–run unemployment coefficient, respectively. The main conclusion is that the outside wage is the main long–run determinant of industry wages within low–wage industries which act as wage followers. Both internal factors, but also the state of the labour market are much more important within high–wage industries which act as wage leaders.

Figures 2, 3 and 4

A remarkable finding is that both the insider weight and the unemployment coefficient, evaluated at the sample mean for relative wages, are much smaller in the LSTR model as compared with the linear model. We further note that the estimate of the equilibrium correction term is 50% higher in the LSTR model as compared with the linear one. One possible explanation of these results is cross–industry parameter heterogeneity which is neglected in the linear model but (at least partly) taken into account in the LSTR model.

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coefficient was independent on the industry's relative wage position.

[12]The relative wage ranges from 0.50 to 1.60 with a sample mean of 1 and standard deviation of 0.14.

Problems related to estimating dynamic models for heterogenous panels are discussed in Pesaran and Smith (1995) and Pesaran *et al.* (1996). They show that the fixed effects estimator of the long–run coefficient overestimates the true long–run effect if the explanatory variable is positively autocorrelated.[13] Interestingly, Johansen (1995b) reports a speed of adjustment coefficient of 0.25 by making use of pooled panel data for 22 industries. The corresponding estimate obtained by averaging the estimates from separate industry wage regressions approximates 0.4.

The remaining estimates obtained from the nonlinear specification are all close to those reported for the linear model. The diagnostic test statistics reported below equation (13) do not indicate any mis–specification of the LSTR model. The estimated standard error of the regression is reduced from 1.933% in the linear model to 1.888% in the LSTR model. Finally, we test the null hypothesis of no remaining nonlinearity. The test is computed by running the auxiliary regression

$$\begin{aligned}\Delta w_{it} &= \boldsymbol{\beta}_0 \mathbf{x}_{it} + \theta \widetilde{x}_{it-1} G_1 (\gamma, c;\ s_{it-1}) + \boldsymbol{\beta}_1^* \widetilde{\mathbf{x}}_{it-1} s_{it-1} + \boldsymbol{\beta}_2^* \widetilde{\mathbf{x}}_{it-1} s_{it-1}^2 + \\ &\quad \boldsymbol{\beta}_3^* \widetilde{\mathbf{x}}_{it-1} s_{it-1}^3 + \alpha_i + v_{it}, \end{aligned} \tag{14}$$

where G_1 is given by equation (5). The null of no remaining nonlinearity, i.e.:

$$\mathrm{H}_0^*\colon\ \boldsymbol{\beta}_1^* = \boldsymbol{\beta}_2^* = \boldsymbol{\beta}_3^* = 0,$$

against the alternative H_1^*: "at least one $\boldsymbol{\beta}_i^* \neq 0$" and the sequence of null hypotheses within (14) are tested using the same procedure as for the linearity tests. As can be seen from Table 2, the results do not indicate any remaining nonlinearity.

Table 2

7 Some further experiments

All results reported above are based on the inverted square specification of the unemployment rate. In order to test the robustness of the results we

[13] Using the model $y_{it} = \lambda_i y_{it-1} + \beta_i x_{it} + \varepsilon_{it}$ with $\lambda_i = \lambda + \eta_{1i}$, $\beta_i = \beta + \eta_{2i}$ and $x_{it} = \mu_i (1-\rho) + \rho x_{it-1} + u_{it}$, they show that the fixed effects estimator of λ is positively biased while the fixed effects estimator of β is negatively biased when $\rho > 0$. In particular, they show that $\text{plim}\widehat{\lambda} = 1$ and $\text{plim}\widehat{\beta} = 0$ when $\rho \to 1$, irrespective of the true parameter values.

estimate the LSTR1 model with the log of the unemployment rate, u, instead of U^{-2}. Results based on this specification are given by

$$\begin{aligned}\widehat{\Delta w_{it}} = & \left[\underset{(3.38)}{0.032}\,(va-wa)_{it-1} - \underset{(2.13)}{0.0122}u_{t-1}\right]\{1+\exp\left(-64\left(s_{it-1}-1.06\right)\right)\}^{-1} \\ & - \underset{(4.95)}{0.315}\,(w-wa)_{it-1} + \underset{(2.89)}{0.027}\,(va-wa)_{it-1} - \underset{(1.94)}{0.0043}u_{t-1} \\ & + \underset{(2.68)}{0.030}\Delta va_{it} - \underset{(2.40)}{0.007}\Delta va_{it-1} + \underset{(45.49)}{0.897}\Delta wa_{it} + \underset{(8.99)}{0.425}\Delta^2 tp_t \\ & - \underset{(4.07)}{0.124}\Delta h_t - \underset{(0.32)}{0.0005}\Delta u_t - \underset{(1.92)}{0.0040}\Delta u_{t-1} + \underset{(0.24)}{0.007}\,(\Delta pc + \Delta ti)_{t-1}\end{aligned} \tag{15}$$

$T = 24\ \ [1966 - 1991]\,,\ N = 116,\ \text{Method: } GMM\,(1)\,,\ Q = 0.2250,$
$\widehat{\sigma}\,(\%) = 1.886,\ m1 = -4.650,\ m2 = -0.111,\ \chi^2_{\mathrm{SAR}}\,(102) = 110.34.$

From equation (15) we note that the estimates of $(va - wa)_{it-1}$ and u_{t-1} are both statistically significant in the nonlinear part of the model. The estimate of the slope parameter, γ, is higher as compared with the estimate based on the inverted square specification while the estimate of the location parameter, c, is almost unaffected. These results imply that the long–run insider weight increases from 0.087 in the low-wage regime to 0.189 in the high-wage regime while the absolute value of the partial long–run unemployment elasticity increases from 0.014 to 0.052.

Finally, we estimate a similar LSTR1 model using *regional* panel data. The wage variable, W_j, is now manufacturing wage costs per worker while VA_j is manufacturing value added per worker and TP_j is the payroll tax rate in municipality j. We include the county unemployment rate, U_c, while the outside wage, WA_j, is the average of manufacturing wage costs per worker outside municipality j. Results using data for 322 Norwegian municipalities for the time period 1974–1992 are given by[14]

$$\begin{aligned}\widehat{\Delta w_{jt}} = & \left[\underset{(1.11)}{0.033}\,(va-wa)_{jt-1} - \underset{(1.57)}{0.0130}u_{ct-1}\right]\{1+\exp\left(-21\left(s_{jt-1}-0.95\right)\right)\}^{-1} \\ & - \underset{(13.26)}{0.569}\,(w-tp-wa)_{jt-1} + \underset{(1.54)}{0.040}\,(va-wa)_{jt-1} + \underset{(0.12)}{0.0007}u_{ct-1} \\ & + \underset{(3.71)}{0.082}\Delta va_{jt} + \underset{(12.90)}{0.681}\Delta wa_j + \underset{(3.72)}{0.569}\Delta tp_j\end{aligned} \tag{16}$$

$T = 18\ [1974 - 92],\ N = 322,\ \text{Method: } GMM(1),\ Q = 2.11,$
$\widehat{\sigma}\,(\%) = 6.55,\ m1 = -10.35,\ m2 = -0.48,\ \chi^2_{\mathrm{SAR}}\,(123) = 132.9.$

[14] Lower case letters denote log transformed variables, see Appendix A.2 for definition and sources. The regional wage equation also contains time dummies for the wage and price freeze in 1979 and the income regulation law in 1988–89.

The estimates imply that the feedback effects of value added per worker is increasing in relative wages, $s_{jt-1} = W_{jt-1}/WA_{jt-1}$. The estimated long–run insider weight is 0.07 in the low–wage regime and 0.13 in the high–wage regime. The estimated partial long–run elasticity of county unemployment is positive in the low–wage regime and -0.02 in the high–wage regime. However, the estimated effects are not significant, neither in the linear part nor in the nonlinear part of the model. If we exclude county unemployment from the linear part, the estimate of u_{ct-1} in the nonlinear part is -0.012 with a t–value of 3.53 while the estimates of $(va - wa)_{jt-1}$ remain insignificant from zero.[15] Although the evidence is weaker, the main results based on the regional data set are consistent with those reported above using industry panel data.

8 Concluding comments

The main issue in the present paper has been to test whether or not the effects on industry wages of internal and external forces depend upon the industry's relative wage position. To do so, we make use of panel data and formulate a smooth transition regression model including relative wages as the transition variable.

The paper provides strong empirical evidence of nonlinear industry wage responses to industry profitability, outside wages and unemployment. The estimated long–run insider weight vary from 0.07 for low–wage industries to 0.16 for high–wage industries while the elasticity of the outside wage vary from 0.84 for high–wage industries to 0.93 for low–wage industries. The estimated partial unemployment effect is very small in the low–wage regime but rather strong in high–wage industries. Aggregate unemployment mainly affects wages in low–wage sectors indirectly through the effect on the outside wage. The main results are robust to alternative transformations of the rate of unemployment, and we also provide some evidence of nonlinear effects using regional panel data.

Our interpretation is that the nonlinear responses reflect asymmetric wage comparison effects. Workers in low–wage sectors are more concerned with relative wages than workers in high–wage sectors. Another possible interpretation is that the nonlinear insider effect reflect union strenght supposed to

[15] It should be noted that the estimates in the linear model given by

$$\widehat{w_{jt}} = -\underset{(12.03)}{0.55}\,(w - tp - wa)_{jt-1} + \underset{(3.55)}{0.07}\,(va - wa)_{jt-1} - \underset{(2.99)}{0.07}\,u_{ct-1} + \underset{(3.81)}{0.09}\,\Delta va_{jt} + \underset{(12.41)}{0.67}\,\Delta wa_{jt} + \underset{(3.46)}{0.54}\,\Delta tp_j,$$

are all significant from zero.

be positively correlated with wages. However, if the nonlinear insider effect reflects union strenght we should expect to find larger unemployment effects in low–wage industries than in high–wage industries. The results reported above are not consistent with such a view.

The results based on the nonlinear model imply smaller average long–run reffects of value added per hour and unemployment than the corresponding estimates obtained from the linear model. Also, the nonlinear model implies more speedy wage adjustments as compared to the results from the linear model. One possible interpretation of these findings is that neglected parameter heterogeneity makes the linear model misspecified which again produces biased estimates.

Throughout we have assumed that the relevant outside wage is the *average* wage outside the sector in question. However, since we suppose that workers in any sector are mainly concerned with their wage relative to wages in higher paid sectors an alternative modelling strategy would be to allow the outside wage differ across sectors.

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A Data sources and definitions

A.1 The industry data set

- W_i = Wage costs per manhour in industry i. Source: *National Account Statistics* (NA), Statistics Norway (SN).
- VA_i = Factor income per hour worked in industry i. Source: NA, SN.
- WA_i = Average wage costs per manhour outside industry i. Source: NA, SN.
- Pc = The official consumer price index. Source: NA, SN.
- U = Aggregate unemployment rate, per cent. Source: *Labour Market Statistics,* SN, and NA, SN.
- H = Normal working time per week in manufacturing. Source: Rødseth and Holden (1990).
- tp = log of 1+ average payroll tax rate. Source: NA, SN.
- ti = log of 1–average income tax rate on housholds. Source, NA, SN.

A.2 The regional data set

- W_j = Manufacturing wage costs per worker in municipality j. Source: *Manufacturing Statistics* (MS), SN.
- VA_j = Manufacturing value added at factor prices per worker in municipality j: Source: MS, SN.
- WA_j = Average manufacturing wage costs outside municipality j. Source: MS, SN.
- U_c = Unemployment rate in municipality j, per cent, Source: *Labour Market Statistics,* SN.
- tp_j = log of 1+payroll tax rate in municipality j. Source: Norwegian Tax Inspectorate.

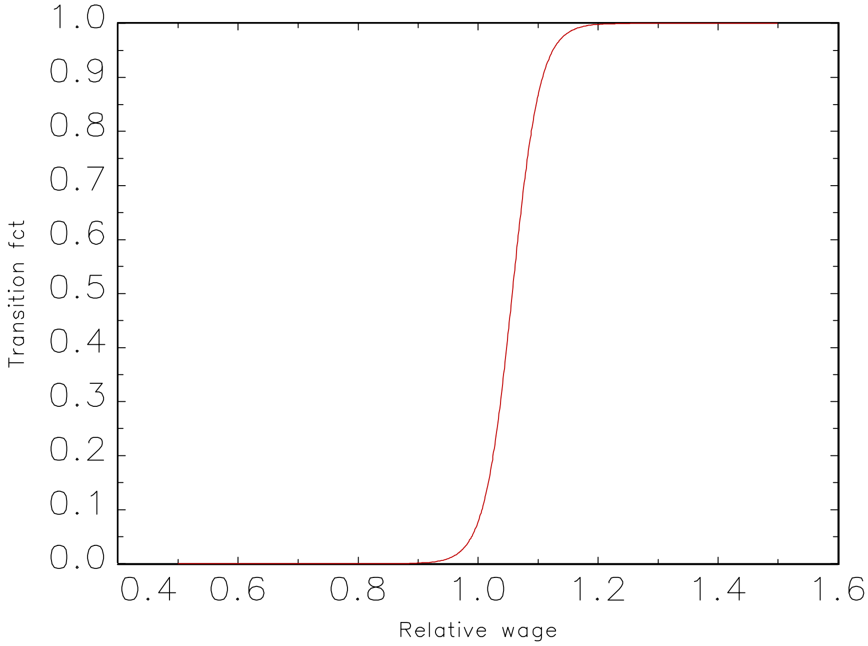

Figure 1: The transition function

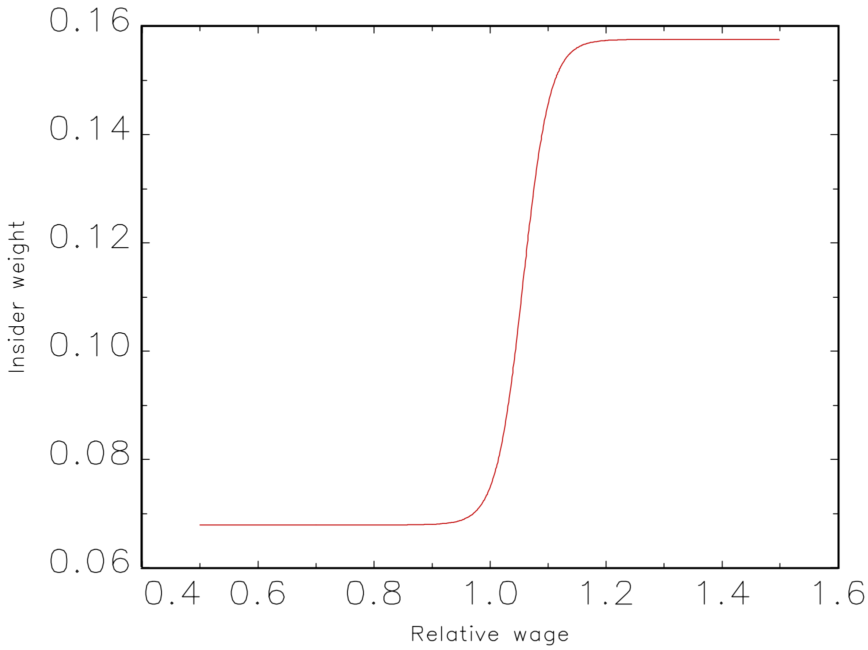

Figure 2: The estimated insider weight

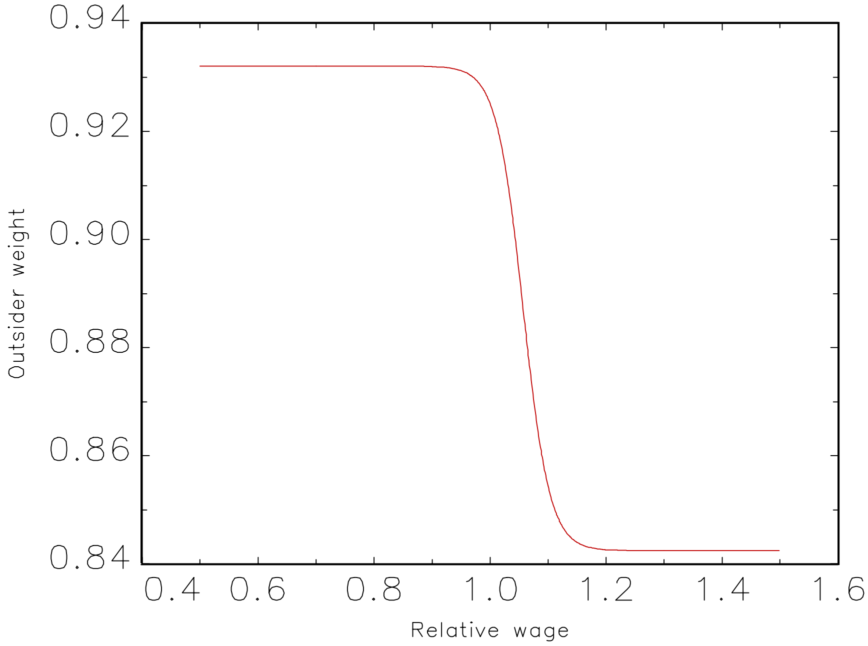

Figure 3: The estimated outsider weight

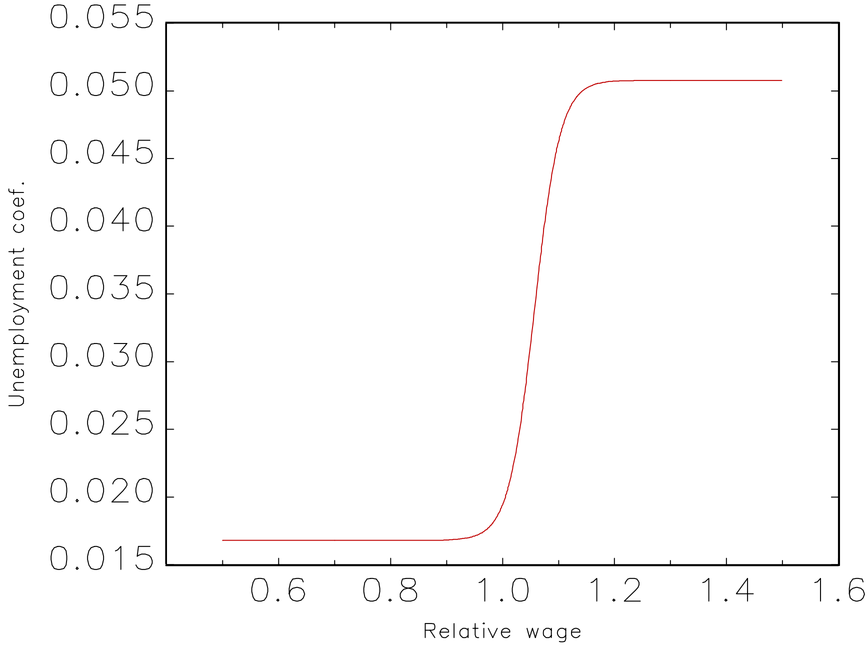

Figure 4: The estimated long-run unemployment coefficient

Table 1: Linearity tests

Hypotheses	Test values [p–values]
H_0: $\boldsymbol{\beta}_1 = \boldsymbol{\beta}_2 = \boldsymbol{\beta}_3 = 0$	$\chi^2(6) =$ 18.53 [0.005]
H_{04}: $\boldsymbol{\beta}_3 = 0$	$\chi^2(2) =$ 0.94 [0.625]
H_{03}: $\boldsymbol{\beta}_2 = 0 \mid \boldsymbol{\beta}_3 = 0$	$\chi^2(2) =$ 4.16 [0.125]
H_{02}: $\boldsymbol{\beta}_1 = 0 \mid \boldsymbol{\beta}_2 = \boldsymbol{\beta}_3 = 0$	$\chi^2(2) =$ 9.24 [0.010]
H_{0va}: $\beta_{12} = \beta_{22} = \beta_{32} = 0$	$\chi^2(3) =$ 9.37 [0.025]
$H_{0u^{-2}}$: $\beta_{13} = \beta_{23} = \beta_{33} = 0$	$\chi^2(3) =$ 12.21 [0.007]

Notes: The Table reports results for $\chi^2(df)$ Wald-tests, implemented in DPD by Arellano and Bond (1988). The first four tests are computed by running the auxiliary regression given by (10) with $\widetilde{x}_{it-1} = \left[(va - wa)_{it-1}\,,\ U_{t-1}^{-2}\right]^T$, while the last two test the hypotheses of no separate nonlinear effects in these two variables.

Table 2: Tests of no remaining nonlinearity

Hypotheses		Test values [p–values]	
H_0^*:	$\boldsymbol{\beta}_1^* = \boldsymbol{\beta}_2^* = \boldsymbol{\beta}_3^* = 0$	$\chi^2(6) =$	6.34 [0.386]
H_{04}^*:	$\boldsymbol{\beta}_3^* = 0$	$\chi^2(2) =$	1.91 [0.385]
H_{03}^*:	$\boldsymbol{\beta}_2^* = 0 \mid \boldsymbol{\beta}_3^* = 0$	$\chi^2(2) =$	1.55 [0.461]
H_{02}^*:	$\boldsymbol{\beta}_1^* = 0 \mid \boldsymbol{\beta}_2^* = \boldsymbol{\beta}_3^* = 0$	$\chi^2(2) =$	0.17 [0.919]

Notes: The Table reports results for $\chi^2(df)$ Wald-tests, implemented in DPD by Arellano and Bond (1988). The tests are computed by running the auxiliary regression given by (14) with $\widetilde{x}_{it-1} = \left[(va - wa)_{it-1},\ U_{t-1}^{-2}\right]^T$.