Price-Wage System with Taxation: Multivariate Cointegration Analysis

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Summary

The paper intends to investigate the price system and the wage equation in the presence of taxes. Price formation is analysed at three levels: producer’s prices, trade in consumer goods and, separately, in services, and at the aggregate level the cost of living index is examined. The empirical analysis is based on Polish data covering the period of transition from a centrally planned economy towards a market economy and more specifically monthly data ranging from January 1993 to December 2003. The empirical results allow to conclude that in the transition years as many as five stable long-run relationships drove inflation in Poland. Appropriate decomposition of price formation made it possible to incorporate all circumstances postulated by economic theory and to show how direct and indirect taxes impact decisions made by the employers and employees.

*JEL classification:* C32; E24; E31
1. Introduction

The prevailing paradigm for a national-economy model in Europe is that firms have mark-up opportunities in the environment of imperfectly competitive goods and labour markets. Wages are determined through bargaining and real activity is independent of the (steady-state) inflation rate (see Wallis (2004)). In addition, the model is expected to show a non-accelerating inflation rate of unemployment, NAIRU, which implies static homogeneity of the price and wage equations.

To be used in policy simulation studies, the model should distinguish between prices of goods and prices of services; additionally, influence exerted by the prices of imported consumer goods should be separated from the impact of prices of imported materials and intermediate products, as these two show significant differences in dynamics. It is also important to introduce direct and indirect taxes to the model, because both of them affect decisions that lead to the formation of wages and prices. To quantify these relationships properly, we explain price formation starting with the producer’s level, then proceed to the consumer’s level, to finally arrive at the price of living index. This approach is in the spirit of classical macromodels, which have usually been built “bottom-to-top”, that is with components summed up to combine the aggregates (see Klein at al. (1999)).

It is a well-documented phenomenon that most macroeconomic variables are first difference stationary, despite some pieces of evidence that price indices can be second difference stationary (see Juselius (1999)). Considering the circumstances, the multivariate cointegration analysis becomes a natural tool to model the price-wage nexus. Decomposition of the behaviour into long-run and short-run dynamics within the
framework of a structural vector equilibrium correction model (SVEqCM) allows to include all assumptions induced by economic theory.

In the case of a disaggregated model for the Polish economy it was not easy to successfully apply the SVEqCM, as the time series was rather limited. Of crucial importance turned out to be the strategy reducing the size of the model early in the analysis by marginalizing it (originally proposed by Greenslade et al. (2000)).

2. The long-run model’s structure

Let us start by assuming that the producer prices are determined as a mark-up on imports and unit labour costs:

\[ pr = \delta^r pm^l + (1 - \delta^r)(w - z) \]  \hspace{1cm} (1a)

where \( pm^l \) is the price of imported materials and intermediate products, \( w \) - nominal wages, and \( z \) – labour productivity. The homogeneity assumption implies that coefficients add up to unity. From this equation it follows that wages can only affect inflation, when their increase is not compensated for by productivity growth. Theoretical foundations can be traced back to the cost-push inflation hypothesis (see classic work of Tobin (1972)).

It is commonly accepted that real wages depend on both productivity, \( z \), and the labour market pressures measured by the rate of unemployment, \( U \):

\[ w - p = \delta^w z - \delta^w U. \]  \hspace{1cm} (1b)

It is possible to include in this equation some extra variables capturing wage pressures that originate from various sources. Static homogeneity of wages implies that \( \delta^w = 1 \). This function results from the accepted bargaining model of wages and prices (see Nickel
(1984), Layard et al. (1991)) and it can be regarded as a standard wage function as well (see Tobin (1995)).

In the simplest case the cost of living index, \( p \), is assumed the same as the producer price index, \( pr \). In practice, however, these two are different for four fundamental reasons. Firstly, the import of consumer goods changes the dynamics of \( p \) independently of \( pr \). Secondly, a value added tax is charged on domestic and imported goods and its rate may vary in time. Therefore, the deflator of consumer goods (commodities), \( pc \), should be expressed as a weighted sum of the deflator of imported final goods, \( pm^F \), and the producer price index, \( pr \), enlarged by a value added tax, \( tv \):

\[
pc = (1 - \delta^c_r) pm^F + \delta^c_r pr + tv.
\] (2)

Thirdly, the cost of living index, \( p \), is the weighted sum of prices of commodities and services, \( ps \):

\[
p = \delta^p_r pc + (1 - \delta^p_r) ps.
\] (3)

Fourthly, taxes are imposed on wages and this significantly modifies objectives driving the two parties involved in the wage-bargaining process. Because of the taxes, employers are concerned with the total real wage costs:

\[
w'c - p = w + tw - p
\] (4a)

where \( tw \) represents employment taxes. Employees, however, are sensitive to the real net wage:

\[
w - td - p = wc - tx - p
\] (4b)

where \( td \) stands for direct taxes and \( tx \) denotes total overheads (\( tx = tw + td \)).
All in all, we considered a system composed of equations explaining:

- producer price

\[ pr = \delta^p \cdot pm^f + (1 - \delta^p) (wc - z), \]  \hspace{1cm} (5a)

- the deflator of consumer goods (commodities)

\[ pc = (1 - \delta^c) pm^f + \delta^c pr + tv, \]  \hspace{1cm} (5b)

- the deflator of services

\[ ps = \delta^s pc + \delta^s h, \]  \hspace{1cm} (5c)

- the consumer price index

\[ p = \delta^p pc + (1 - \delta^p) ps, \]  \hspace{1cm} (5d)

- wages

\[ wc - tx - p = z - \delta^w U \]  \hspace{1cm} (5e)

where \( h \) is the share of services in total consumption and, when growing, it represents the potential market pressures. Consequently, the consumer price index can be written as:

\[ p = [\delta^p (1 - \delta^c) + \delta^s] (pr + tv + tm) + (1 - \delta^p) \delta^s h \]  \hspace{1cm} (6)

where \( tm = (1 - \delta^c) (pm^f - pr) \) and it can be interpreted as a ‘tax’ imposed due to high import prices (see also Wallis (2004)). In the long run the dynamics of services’ prices does not differ from the dynamics of commodities’ prices, \( \delta^s = 1 \), therefore everything that matters beside producer prices is the weighted share of services in total consumption and taxes:

\[ p = (pr + tv + tm) + (1 - \delta^p) \delta^s h. \]  \hspace{1cm} (7)
3. The data

The empirical investigation is based on Polish monthly, seasonaly unadjusted data covering the period from January 1993 to December 2003. It was the time when the Polish economy evolved from a centrally planned economy towards a market economy following the change of the political system. However, the first stage of the transition (i.e. years 1990-1992), when the draconian devaluation of Polish zloty and adjustment of prices took place, making the cost of living index go up over eightfold, was intentionally omitted (wider discussion in Welfe at al. (2004)). A worth noting fact is the commonly accepted view that the Polish economy started to be demand-driven only from 1993, in contrast to its supply-constrained nature in the past. Consequently, assumptions typical of the market economies can be applied. An additional, nevertheless important reason for skipping the first transition years was unavailability of the monthly data. The series are plotted in Figure 1. Small letters denote natural logarithms.

<Figure 1. around here>

This study concentrates on the price-wage system in the presence of taxes. The model is therefore built around five relationships (5a) to (5e). The set of variables includes the price index of production sold (producer prices), $pr$, the price index of consumer goods, $pc$, the price index of services, $ps$, the consumer price index (cost of living index), $p$, total wage costs in industry (current prices), $wc$.

The set of potentially weakly exogenous variables consists of the price index of imported materials and intermediate products (used in production), $pm^i$, the price index of imported consumer goods, $pm^c$, productivity in industry measured by value added per
worker (costant prices), \( z_t \), unemployment rate, \( U_t \), value added tax, \( tv_t \), taxes imposed on wages, \( t\), and the share of services in total consumption (constant prices), \( h_t \). Taxes are expressed as proportion and in fact they represent amounts contributed to the state budget.

Results presented in Table 1 prove that all variables are I(1), even though the joint ADF-KPSS test (for small sample critical values see Kęblowski, Welfe (2004)) suggests that some of them could be considered I(2). For every variable at a minimum level (varying from 3 to 12) a number of lags were chosen in the ADF to remove serial correlation and to ensure normal distribution of residuals, which was verified by the cumulated periodogram test (see Durbin (1969)).

<Table 1. around here>

4. Structural VEqCM. Model marginalisation

Since all variables proved to be I(1), the natural tool in the empirical analysis is the structural vector equilibrium correction model (SVEqCM):

\[
\Delta y_t A_0 = \bar{y}_{t-1} \Pi + \sum_{s=1}^{S-1} \Delta y_{t-s} A_s + d^T \bar{C} + \varepsilon_t
\]  

where:

\( \Pi = \Pi A_0 \), \( A_s = \Gamma_s A_0 \), \( \bar{C} = C A_0 \), \( \varepsilon_t = \xi_t A_0 \), and

\( A_0 \) - contemporaneous coefficients matrix,

\[ \bar{y}_t = [y_t, d_t'] \],

\( y_t = [y_{1t} \ldots y_{Mt}] \) - vector of \( M \) stochastic variables,
\( \mathbf{d}_t^x = [d_{1t} \ldots d_{Nt}] \) - vector of \( N \) deterministic variables,

\( \mathbf{d}_t^z = [d_{1t} \ldots d_{Pt}] \) - vector of \( P \) deterministic variables outside the cointegration relation,

\( \mathbf{\Pi} \) - total impact multipliers matrix, which can be decomposed as \( \mathbf{\Pi} = \mathbf{BA}^T \), provided the cointegrating rank of the system is \( R \ (0 \leq R < M) \),

\( \mathbf{\Gamma}_s \) - short-run parameter matrices,

\( \mathbf{C} \) - deterministic coefficients matrix,

\( \mathbf{\xi}_t = [\xi_{1t} \ldots \xi_{Mt}] \) - vector of white noise disturbances.

Results of the Monte Carlo experiments indicate that the cointegration test works better after marginalisation has been determined; but it is asymptotically irrelevant, whether the overidentifying restrictions are tested before or after the model has been marginalized and its dynamic structure set (see Greenslade et al. (2000)). Therefore, in the first step the cointegrating rank was found using a standard Johansen procedure. Following that, weak-exogeneity hypotheses were tested. Because the exogeneity tests are sensitive to the cointegrating rank, the procedure was iteratively repeated. In the third step all economic restrictions were imposed to define the long-run structural model and to allow parameter estimation. In the last step the short-run structure was found.

The starting point of the empirical analysis was a VAR model with three lags and a deterministic component. The assumption that this system provides a framework that is general enough to allow extraction of the short-run effects from the data in the course of reduced rank regression seems to be a reasonable compromise regarding the short time series, but misspecification tests were used to verify it anyway.

<Table 2. around here>
Results presented in Table 2 prove that residuals from all equations are not autocorrelated and support the choice of three lags. The non-normality of residuals from some equations is perceived as resulting from a huge number of parameters of unrestricted VAR and it does not cause problems as the main reason for it is excess kurtosis, for which reduced rank regression is robust (see Gonzalo (1994)). Furthermore, no ARCH effect was detected.

The cointegration rank was determined using two standard tests: the maximum eigenvalue and the trace. Both gave comparable results (see Table 3). However, results provided by inference based on the trace test were treated as more plausible, because Lütkepohl (2001) argues that in some cases the trace test in cointegrated systems with a number of cointegrating relations and limited samples is superior to the maximum eigenvalue test in power terms. Additionally, to reduce the small sample distortion, the Bartlett correction factor (see Johansen (2002)) and the correction derived by Ahn and Reinsel (1990) and Reimers (1992) were calculated. Inference on the cointegration rank alternated with inference on the weak-exogeneity of variables (see Tables 3 and 4, respectively). Pesaran-Shin (2000) asymptotic critical values (for 5% size) were used, which are conclusive when conditional inference is being made.

In the first step $\lambda_{\text{trace}}$, $\lambda_{\text{trace}}^{BC}$ suggested that at least five common trends exist, while $\lambda_{\text{max}}$ and $\lambda_{\text{trace}}^{AHC}$ implied as many as seven. The $LR$ test for weak-exogeneity indicated in turn that the price index of imported materials and intermediate products, productivity in industry, value added tax, the price index of imported consumer goods, and the rate of unemployment should be treated as weakly exogenous, which necessitated another
inference on the cointegration rank in a system conditional on these variables. The result of the trace test in the second iteration suggested that at least two more common trends exist, whereas $\lambda_{trace}^{AHC}$ indicate even at four additional common trends and the LR test for weak-exogeneity indicated that wage taxes and the share of services in total consumption should be considered weakly exogenous. In the last iteration, the cointegration rank was tested for a system conditional on the seven aforementioned variables and no more common stochastic trends were found.

The long-run structure was identified by subjecting parameters of matrix $B$ to exclusion and homogeneity restrictions. Fifty five restrictions were imposed altogether as can be deduced from (9a) – (9e) and they were not rejected by the LR test for over-identifying restrictions due to statistic value $\chi^2(30)=42.11$. At this stage of the analysis, matrix $A$ parameters were left unrestricted. The constant terms were decomposed into those restricted to the cointegration space and unrestricted, according to the identity exploiting an orthogonal compliment to matrix $A$ (see Johansen (1996)).

Tests of the marginalised system’s residuals presented in Table 5 confirm that the model is acceptable. Especially strong evidence in support of this opinion is results of the multivariate versions of the $LM$ test of residual autocorrelation and the Doornik-Hansen test of normality. The univariate test’s outcomes are also satisfactory, however, kurtosis shows weak signs of shocks and interventions affecting prices of consumer goods and services. Worth mentioning are high $R^2$ values across all the equations, but particularly in those explaining wages, prices of services and consumer prices.

<Table 5. around here>
The normality of the residuals from the consecutive equation, the lack of ARCH
effect and serial correlation assure that the inference made so far is valid. Besides, it
enables to seek a more parsimonious structure of parameters than the existing one.

5. The structural price-wage model. Empirical results

The maximum likelihood estimates of the long-run relationships are (absolute
values of t-statistics are bracketed under the parameters):

\[ pr = 0.313 + 0.258 \, pm^F + 0.742(we - z) \]
\[ \text{(3.79) (10.91)} \]  
\[ (9a) \]

\[ pc - tv = -0.071 + 0.137 \, pm^F + 0.863 \, pr \]
\[ \text{(2.02) (12.69)} \]  
\[ (9b) \]

\[ ps = 1.095 + pc + 2.183h \]
\[ \text{(12.06)} \]  
\[ (9c) \]

\[ p = 0.005 + 0.720 \, pc + 0.280 \, ps \]
\[ \text{(180.06) (69.94)} \]  
\[ (9d) \]

\[ wc - tx - p - z = 0.387 - 0.045U \]
\[ \text{(15.04)} \]  
\[ (9e) \]

The results are fully interpretable in economic terms and the model allows to
identify how wages induce inflation in the long-run and, conversely, how the costs of
living influence wages in the presence of exogenous taxes.

Since (8) is entirely based on stationary variables, the standard t-ratios as well as
other classical tests retain their properties and can be used to find loading coefficients
and short-run parameters that significantly differ from zero. To find as parsimonious
dynamic structure of the model as possible, the iterative procedure was applied. At first,
the parameters of loading matrix and the dynamics of endogenous variables were left
unrestricted, while regressors with the smallest absolute values of t-ratios were
iteratively removed from the rest of the short-run structure. Next, all statistically insignificant variables and unimportant cointegrating vectors were omitted. This led to the following ($t$-statistics are parenthesised):

\[
\begin{bmatrix}
\Delta p_t & \Delta p_c & \Delta p_s & \Delta p_t & \Delta w_{c_t} \\
0 & -0.034 & 0 & -0.024 & -0.505 \\
0.057 & 0.119 & 0 & 0.085 & 0.477 \\
0.025 & 0.006 & 0 & 0 & 0 \\
0 & 1.689 & 0.635 & 1.41 & 6.497 \\
0 & -0.011 & -0.007 & -0.011 & 0.052
\end{bmatrix}
\times
\begin{bmatrix}
\text{ec}_{1,t-1} & \text{ec}_{2,t-1} & \text{ec}_{3,t-1} & \text{ec}_{4,t-1} & \text{ec}_{5,t-1}
\end{bmatrix} + \begin{bmatrix}
0 & 0 & 0 & 0 & 0.537 \\
0 & 0.126 & 0.090 & 0.119 & 0 \\
0 & 0 & 0 & 0 & 0.0003 \\
0 & 0 & 0 & 0 & 0.046
\end{bmatrix}
\]

\[
\begin{bmatrix}
\Delta z_t & \Delta v_t & \Delta U_t \\
0 & 0 & 0 & 0 & 0.537 \\
0 & 0.126 & 0.090 & 0.119 & 0 \\
0 & 0 & 0 & 0 & 0.0003 \\
0 & 0 & 0 & 0 & 0.046
\end{bmatrix}
\times
\begin{bmatrix}
\Delta p_{t-1} & \Delta p_{c_{t-1}} & \Delta p_{s_{t-1}} & \Delta p_{t_{t-1}} & \Delta v_{t-1} & \Delta U_{t-1}
\end{bmatrix}
\]

\[
\begin{bmatrix}
0.131 & 0 & 0 & 0 & 0 \\
0 & 0.107 & 0.265 & 0 & 0 \\
0 & 0 & 0.211 & 0 & 0 \\
0 & 0 & 0 & 0.212 & 0 \\
0 & 0 & 0 & -0.012 & 0 \\
0 & 0 & 0 & 0.001 & -0.065
\end{bmatrix}
\times
\begin{bmatrix}
\Delta p_{c_{t-2}} & \Delta p_{t_{t-2}} & \Delta w_{c_{t-2}} & \Delta p_{m_{t-2}} & \Delta U_{t-2}
\end{bmatrix}
\]

\[
\begin{bmatrix}
0 & 0.021 & 0 & 0 & 0 \\
0 & 0 & 0 & 0.019 & 0 \\
0 & 0 & 0 & 0 & -0.146 \\
0 & -0.037 & 0 & -0.027 & 0 \\
0 & 0 & -0.001 & 0 & 0.036
\end{bmatrix}
\times
\begin{bmatrix}
\Delta p_{c_{t-2}} & \Delta p_{t_{t-2}} & \Delta w_{c_{t-2}} & \Delta p_{m_{t-2}} & \Delta U_{t-2}
\end{bmatrix}
\]

\[
\begin{bmatrix}
m93.7_t & m9495.3_t & m95.1_t & m95.7_t & m9698.1_t & m0003.1_t
\end{bmatrix} \times
\begin{bmatrix}
0.015 & -0.006 & 0.015 & 0 & 0 \\
0.014 & 0 & 0.014 & 0.004 & 0 \\
0 & -0.015 & 0 & -0.011 & 0 \\
0 & 0 & 0.022 & 0.006 & -0.035 \\
0 & 0 & 0 & 0 & -0.042
\end{bmatrix}
\]

\[
\begin{bmatrix}
m4_t & m8_t & m9_t & m10_t & m11_t
\end{bmatrix}
\begin{bmatrix}
0 & 0 & 0.0003 & 0 & 0 \\
0.002 & 0 & 0 & 0 & -0.015 \\
0 & 0 & 0 & 0 & -0.027 \\
0 & 0 & 0 & 0 & -0.041 \\
0 & 0 & -0.0003 & 0 & 0
\end{bmatrix}
\]

\[
\begin{bmatrix}
0.001 & 0 & 0 & 0 & 0.022
\end{bmatrix} + \begin{bmatrix}
u_{1,t} & u_{2,t} & u_{3,t} & u_{4,t} & u_{5,t}
\end{bmatrix}
\]

where \( \text{ec}_1, \text{ec}_2, \text{ec}_3, \text{ec}_4 \) and \( \text{ec}_5 \) are cointegrating vectors as specified by (9a), (9b), (9c), (9d) and (9e), respectively. The \( m4, m8, m9, m10 \) and \( m11 \) are seasonal centred dummies for the appropriate months. Other variables that start with “m” are also dummies, for example \( m9495.3 \) takes ones in March in years 1994 and 1995. It is worth stressing that various ways of eliminating insignificant regressors led to very similar structures.

Interestingly, in the equation of the producer prices only two cointegrating vectors identified as the long-run equations for consumer goods prices and prices of services were significant, while the cointegrating vector defining the producer prices’ disequilibrium was not. In fact, this allows to state that producers are more likely to adjust themselves to pressures arising from the consumer prices than to their own prices,
which can serve as indirect evidence that the Polish economy is demand driven indeed. In the remaining equations, the structure of loading coefficients is rather rich and parameters located on the diagonal of matrix $A$ are non-zero, so it is hard to give the economic interpretation.

Constants outside the cointegration space proved to be statistically significant only in the equation explaining producer prices and wages, but even then their values are very small. This means that none of the variables explained by the model shows an autonomous growth.

6. Conclusions

A multivariate cointegration analysis enabled to build a disaggregated model of the price-wage behaviour with nonstationary variables under the presence of direct and indirect taxes. In addition, the model meets all economic assumptions about the long-run behaviour of the interesting variables. In particular, static homogeneity is implicit in both price and wage equations. As many as five identified cointegrating vectors proved to be significant for explaining variations in the modelled variables in the presence of quite rich short-run dynamics.

References


Figure 1. Seasonally unadjusted monthly data, January 1993 to December 2003
Table 1. Inference on the order of integration

<table>
<thead>
<tr>
<th>variable</th>
<th>hypotheses</th>
<th>t statistic of the ADF test</th>
<th>LM statistic of the KPSS test</th>
<th>ADF test conclusion, limit distribution size - 5%</th>
<th>KPSS test conclusion, limit distribution size - 5%</th>
<th>ADF-KPSS joint test conclusion, exact distribution</th>
<th>K-S statistics of the cumulated periodogram test</th>
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<td>I(1)</td>
<td>I(1)</td>
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Table 2. Misspecification tests

<table>
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<tr>
<th>equation</th>
<th>K-S statistics of Durbin test</th>
<th>kurtosis</th>
<th>normality, $\chi^2(2)$</th>
<th>ARCH(3), $\chi^2(3)$</th>
<th>ARCH(12), $\chi^2(12)$</th>
<th>$R^2$</th>
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<tbody>
<tr>
<td>$pr_t$</td>
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<td>16.156</td>
<td>28.402</td>
<td>0.903</td>
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<td>3.006</td>
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<td>1.327</td>
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<td>$tv_t$</td>
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<td>20.600</td>
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Table 3. Inference on the cointegration rank

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<th>$\lambda_{\text{max}}$</th>
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<th>$\lambda_{\text{BC}}^{\text{trace}}$</th>
<th>$\lambda_{\text{AHC}}^{\text{trace}}$</th>
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<tbody>
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<td>591.54</td>
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<td>11</td>
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<td>469.99</td>
<td>454.32</td>
<td>341.68</td>
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<td>10</td>
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<td>370.37</td>
<td>361.90</td>
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<td>9</td>
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<td>283.84</td>
<td>276.78</td>
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<td>8</td>
<td>62.12</td>
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<td>151.02</td>
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<td>99.70</td>
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<td>2.36</td>
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weakly exogenous variables: pm$t_i$, $z_t$, $t_{vt}$, pm$F_t$, $U_t$

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<th></th>
<th>$\lambda_{\text{max}}$</th>
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<th>$\lambda_{\text{BC}}^{\text{trace}}$</th>
<th>$\lambda_{\text{AHC}}^{\text{trace}}$</th>
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<tbody>
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<td>9.64</td>
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weakly exogenous variables: pm$t_i$, $z_t$, $t_{vt}$, pm$F_t$, $U_t$, $t_{xt}$, $h_i$

<table>
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<th>$\lambda_{\text{max}}$</th>
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<th>$\lambda_{\text{BC}}^{\text{trace}}$</th>
<th>$\lambda_{\text{AHC}}^{\text{trace}}$</th>
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Table 4. Inference on weak-exogeneity

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<td>$p_t$</td>
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<td>$pr_t$</td>
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<tr>
<td>$wc_t$</td>
<td>29.28</td>
</tr>
<tr>
<td>$pm_{I_t}$</td>
<td>16.66</td>
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<tr>
<td>$tv_t$</td>
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<tr>
<td>$Ut$</td>
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<tr>
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</table>

weakly exogenous variables:
$pm_{I_t}, z_t, tv_t, pm_{I_t}, Ut$

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<th>variable</th>
<th>$\chi^2(5)$</th>
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<td>$p_t$</td>
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<td>$h_t$</td>
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Table 5. Residual analysis

<table>
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<tr>
<th>equation</th>
<th>K-S statistics of Durbin test</th>
<th>kurtosis</th>
<th>normality, $\chi^2(2)$</th>
<th>ARCH(3), $\chi^2(3)$=</th>
<th>ARCH(12), $\chi^2(12)$=</th>
<th>$R^2$</th>
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</thead>
<tbody>
<tr>
<td>$\Delta pr_t$</td>
<td>0.094</td>
<td>3.198</td>
<td>1.038</td>
<td>10.068</td>
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<td>$\Delta pc_t$</td>
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<td>$\Delta ps_t$</td>
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<tr>
<td>$\Delta pt_t$</td>
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<td>0.770</td>
<td>2.421</td>
<td>25.106</td>
<td>0.962</td>
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</table>

**MULTIVARIATE TESTS**

- residual autocorrelation LM1 $\chi^2(25) = 31.233$ $p$-value = 0.18
- residual autocorrelation LM4 $\chi^2(25) = 18.976$ $p$-value = 0.80
- normality: LM $\chi^2(10) = 8.243$ $p$-value = 0.61