ASYMMETRIC PRICE TRANSMISSION IN FOOD SUPPLY CHAINS: IMPULSE RESPONSE ANALYSIS BY LOCAL PROJECTIONS

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Abstract
In this paper we set out Jorda's (2005) method of local projections by which nonlinear impulse responses can be computed without the need to specify and estimate the underlying nonlinear dynamic system. The method is used to compute price reaction functions that show how the prices of the different stages in the supply chain dynamically respond to one another and whether or not these responses reveal any asymmetric patterns. Empirical applications for the US pork-meat and broiler composite chains illustrate the convenience of the method.

1. Introduction
The relationship between producer prices and consumer prices receives lots of attention by practitioners and scientists because of its diagnostic capability to reveal market imperfections. Indicative of excess profits would be a large, persistent and/or diverging difference between the two prices. Visual inspection of the time series graphs of both prices in one diagram is likely to be a sufficient device for detecting such a pattern. Less straightforward to find out about is the well-known phenomenon of asymmetric price adjustment according to which, for example, transmission of producer price increases to retail prices is of a greater magnitude and occurs more quickly than transmission of producer price decreases. In addition to a graphical analysis, multivariate time-series models will be needed to identify the diverse transitory and persistent dynamics in the price series. Moreover, to capture asymmetric responses, a threshold or nonlinear specification of the time-series model must be considered. Such time-series models, in contrast to traditional linear models like a Vector Auto-Regression (VAR), do usually not allow for simple derivation of the point and interval estimates of the impulse response functions to assess the short-, intermediate- and long-run price reactions to a change in one of the supply chain prices.

Recently, however, Jorda (2005) has introduced the method of local projections by which nonlinear impulse responses can be computed without the need to specify and estimate the underlying nonlinear dynamic system. In this paper we set out Jorda's method to compute price reaction functions that show how the prices of the different stages in the supply chain dynamically respond to one another and whether or not these responses reveal any asymmetric patterns. Empirical applications for the US pork-meat and broiler composite chains illustrate the convenience of the method.

2. Impulse responses by local projections
Traditional impulse responses are multi-period ahead predictions computed on the basis of a model in which the coefficients have been estimated by using a sample of time series observations. Unfortunately, neither a within nor out-of-sample multi-period ahead prediction performance evaluation is usually presented before an impulse response analysis is conducted. Consequently, models yielding inadequate predictions beyond a certain prediction horizon are often used for computing impulse responses at prediction horizons much farther away. Recently, Jorda (2005) introduced the method of local projections for deriving impulse responses which may solve this problem to some extent as impulse responses by local projections are, in fact, within sample direct multi-step forecasts and hence, are utilising more information than just using the sample observations for estimation of the model parameters. The following bi-variate producer-retailer

price models illustrate the projection method. Let $P_{p,t}$ be the producer price and $P_{r,t}$ the retail price in period t. Then, the linear projection model of order one for the retail price is given by

(1) $$P_{r,t+1} = \alpha_r^{(1)} + \beta_{rp}^{(1)}P_{p,t} + \beta_{rr}^{(1)}P_{r,t} + u_{r,t+1}$$

where the residuals $u_{r,t+1}$ are Gaussian white noise. Given that the linear projection model for the producer price is also of order one, then SUR is not needed as simple OLS regression to (1) already yields an efficient estimate of $\beta_{rp}^{(1)}$, which is the one-period ahead (as indicated by the superscript index (1)) direct impulse response of the retail price after a producer-price-specific one-unit shock in period t. The two-periods ahead impulse response $\beta_{rp}^{(2)}$ is then consistently estimated by the OLS regression

(2) $$P_{r,t+2} = \alpha_r^{(2)} + \beta_{rp}^{(2)}P_{p,t} + \beta_{rr}^{(2)}P_{r,t} + u_{r,t+2}$$

etc. Hence, by a separate regression for each prediction horizon h ($h = 1, 2, \ldots$) the retail price $P_{r,t+h}$ is projected onto the information set including all observations on the retail and producer prices up to and including period t. The estimates $\beta_{rp}^{(1)}, \beta_{rp}^{(2)}, \ldots, \beta_{rp}^{(h)}, \ldots, \beta_{rp}^{(H)}$ form the impulse responses from period $t + 1$ to period $t + H$ displaying how the retail price reacts to a producer-price specific one-unit shock in period t. Notice that the residuals $u_{r,t+h}$ are a moving average of the prediction errors from time $t + 1$ to $t + h$. Although these errors are uncorrelated with the regressors, which are dated t, so that the impulse responses are *consistently* estimated, *efficient* estimates can only be obtained if we take the moving average structure explicitly into account. This may complicate the estimation of the projection, but Jorda (2005) reports that only little loss of efficiency results when performing the projection regressions with a heteroskedasticity and auto-correlation (HAC) robust estimator like the one provided by Newey and West (1987), which is nowadays available in many standard regression packages like the EViews 6.0 software that we used for our computations.

So far we have considered local-*linear* projections. More flexible specifications are straightforward to apply, like the following threshold model which allows for asymmetric impulse responses

(3) $$\begin{aligned}\Delta P_{r,t+h} &= (\alpha_r^{(h)-} + \beta_{rp}^{(h)-}\Delta P_{p,t})I(\Delta P_{p,t} \le 0) + (\alpha_r^{(h)+} + \beta_{rp}^{(h)+}\Delta P_{p,t})I(\Delta P_{p,t} > 0) \\ &\quad + \beta_{rr}^{(h)}\Delta P_{r,t} + u_{r,t+h}\end{aligned}$$

for $h = 1, 2, \ldots$, where $\Delta P_{p,t} \equiv P_{p,t} - P_{p,t-1}$, $E(\Delta P_{p,t}) = 0$, and $I(\cdot)$ is the indicator function such that $I(a) = 1$ if a is true, else $I(a) = 0$. In (3) the impulse responses triggered by a negative producer-price-specific one-unit shock, given by the $-\beta_{rp}^{(h)-}$ estimates, do not have to be just the opposite of the impulse responses after a positive producer-price-specific one-unit shock as provided by the $\beta_{rp}^{(h)+}$ estimates. In the next section we employ the threshold in model (3) to assess the asymmetry in the producer-retailer price transmission.

3. Empirical applications

To illustrate the method of local projections for estimating impulse responses, two empirical cases are considered for which montly prices ($ cents per pound, retail weight equivalent) are obtained from USDA for the sample period January 1990 up to and including December 2008. For the first case we study the relationship between the wholesale price and the retail price of broiler composite. The composite wholesale and retail prices are a weighted average of whole chicken prices and prices for parts. The weights are based on estimates of the percentage of chicken sold as parts versus whole.

The second empirical application concerns the pork chain of which we consider the relationships between the farm price, the wholesale price and the retail price. Clearly, with two prices the broiler case is less involved than the analysis of the three prices in the pork chain. Therefore, we start with the broiler case before applying the method to the pork chain.

Figure 1. Monthly wholesale and retail prices of broilers in the US in $ cents per pound, retail weight equivalent for the period January 1990 up to and including December 2008

Figure 1 displays the time series of the two broiler composite prices. From visual inspection and formal unit root and cointegration tests, using the Johansen (1995) procedure, it appears that both prices are integrated of order one and not cointegrated. Consequently, we base our local projections on a VAR model in first differences. To test for asymmetry, we allow each price coefficient and the intercept term in the local projection regressions to differ between positive and negative first differences of the variable to which the coefficient is attached. Furthermore, we impose the contemporaneous identification restriction according to which the current retail price is always based on the current wholesale price. Then, the retail price projection regressions become

$$(4)\qquad P_{r,t+h} - P_{r,t-1} = \alpha_r + \sum_{s=1}^{11} \delta_{rs} D_{st} + \sum_{i=0}^{m_r}\{(\alpha_{rwi}{}^{(h)-} + \beta_{rwi}{}^{(h)-} \Delta P_{w,t-i}) I(\Delta P_{w,t-i} \le 0) + \beta_{rwi}{}^{(h)+} \Delta P_{w,t-i}\, I(\Delta P_{w,t-i} > 0)\} + \sum_{i=1}^{n_r}\{(\alpha_{rri}{}^{(h)-} + \beta_{rri}{}^{(h)-} \Delta P_{r,t-i}) I(\Delta P_{r,t-i} \le 0) + \beta_{rri}{}^{(h)+} \Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + u_{r,t+h}$$

for $h = 0, 1, \ldots$, where $P_{w,t}$ is the wholesale price and the D_{st} are seasonal dummies. Notice that the endogenous variable, $P_{r,t+h} - P_{r,t-1}$, represents the accumulated predictions, since $P_{r,t+h} - P_{r,t-1} = \Delta P_{r,t} + \Delta P_{r,t+1} + \ldots + \Delta P_{r,t+h}$. Consequently, the parameters $-\beta_{rw0}{}^{(h)-}$ and $\beta_{rw0}{}^{(h)+}$ represent the accumulated retail-price-change impulse responses and hence, the impulse responses of the retail price level, after a negative and positive one-unit shock, respectively, in the wholesale price in period t. Similarly, the price projection regressions for the wholesale price that we use are given by

$$(5)\qquad P_{w,t+h} - P_{w,t-1} = \alpha_w + \sum_{s=1}^{11} \delta_{ws} D_{st} + \sum_{i=1}^{m_w}\{(\alpha_{wwi}{}^{(h)-} + \beta_{wwi}{}^{(h)-} \Delta P_{w,t-i}) I(\Delta P_{w,t-i} \le 0) + \beta_{wwi}{}^{(h)+} \Delta P_{w,t-i}\, I(\Delta P_{w,t-i} > 0)\} + \sum_{i=1}^{n_w}\{(\alpha_{wri}{}^{(h)-} + \beta_{wri}{}^{(h)-} \Delta P_{r,t-i}) I(\Delta P_{r,t-i} \le 0) + \beta_{wri}{}^{(h)+} \Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + u_{w,t+h}$$

for $h = 1, 2, \ldots$ to estimate the wholesale price level impulse responses $-\beta_{wr1}{}^{(h)-}$ and $\beta_{wr1}{}^{(h)+}$ that are triggered by a negative and positive one-unit shock, respectively, in the retail price in period t. Notice that these impulse responses are zero in period t.

To directly estimate the impulse responses in the retail price level as a consequence of a one-unit shock in the retail price itself in period t, we have to re-specify the expression in (4) to obtain

$$(6)\qquad P_{r,t+h} - P_{r,t-1} = \alpha_r{*} + \sum_{s=1}^{11} \delta_{rs}{*}D_{st} + \sum_{i=1}^{m_r^*}\{(\alpha_{rwi}{}^{(h)-}{*} + \beta_{rwi}{}^{(h)-}{*}\Delta P_{w,t-i})I(\Delta P_{w,t-i} \le 0) + \beta_{rwi}{}^{(h)+}{*}\Delta P_{w,t-i}\, I(\Delta P_{w,t-i} > 0)\} + \sum_{i=1}^{n_r^*}\{(\alpha_{rri}{}^{(h)-}{*} + \beta_{rri}{}^{(h)-}{*}\Delta P_{r,t-i})I(\Delta P_{r,t-i} \le 0) + \beta_{rri}{}^{(h)+}{*}\Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + u_{r,t+h}{*}$$

for h = 1, 2, … so that the estimates of $-\beta_{rr1}{}^{(h)-}{*}$ and $\beta_{rr1}{}^{(h)+}{*}$ are the impulse responses we look for as in (6) the link between the most recent lag in the retail price change and the unlagged wholesale price change has been eliminated by taking out the latter term. For a similar reason, but now regarding the contemporaneous relationship between the first differences of the wholesale price and those of the retail price, we re-specify (5) by taking out the one-period lag of the retail price change, obtaining

$$(7)\qquad P_{w,t+h} - P_{w,t-1} = \alpha_w{*} + \sum_{s=1}^{11} \delta_{ws}{*}D_{st} + \sum_{i=1}^{m_w^*}\{(\alpha_{wwi}{}^{(h)-}{*} + \beta_{wwi}{}^{(h)-}{*}\Delta P_{w,t-i})I(\Delta P_{w,t-i} \le 0) + \beta_{wwi}{}^{(h)+}{*}\Delta P_{w,t-i}\, I(\Delta P_{w,t-i} > 0)\} + \sum_{i=2}^{n_w^*}\{(\alpha_{wri}{}^{(h)-}{*} + \beta_{wri}{}^{(h)-}{*}\Delta P_{r,t-i})I(\Delta P_{r,t-i} \le 0) + \beta_{wri}{}^{(h)+}{*}\Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + u_{w,t+h}{*}$$

for h = 1, 2, … to estimate $-\beta_{ww1}{}^{(h)-}{*}$ and $\beta_{ww1}{}^{(h)+}{*}$ as the impulse responses in the wholesale price level initiated by a one-unit shock in the wholesale price itself in period t.

A maximum lag length of 6 appears to reduce the residual term to white noise. Next, we use the AIC model selection criterion to determine the number of lags m and n in each of the equations (4)-(7) at $h = 0$. In fact, one could repeat this for each prediction horizon h = 1, 2, …, but for our empirical application we assume the selected numbers for m and n to be representative for each h. The standard errors of the impulse response coefficients could be estimated by a HAC robust estimator. For our computations we choose another, but very easy to implement, approach. Each time a regression is run with $h > 0$, we first run the same regression for $h - 1$ and insert the residual term of this regression as a regressor in the regression with h. In this way the dynamics in between t and $t + h$ are captured, leaving the coefficient estimates unchanged and consistent as they were, but reducing their standard errors towards efficiency levels.

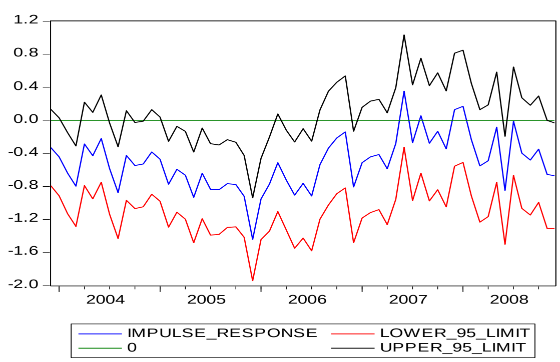

Figure 2. Monthly impulse responses retail price obtained by the sum of the retail price impulse responses after a negative one-unit shock in the wholesale price in December 2003 and the retail price impulse responses after a positive one-unit shock in the wholesale price in December 2003 (broilers in the US, responses in $ cents per pound, retail weight equivalent)

Figure 2 presents the net result of the impulse responses of the retail price to negative and positive shocks in the wholesale price, taking December 2003 as the month in which the price shocks occur and computing the impulse responses for the period thereafter (i.e., January 2004 - December 2008). After one year, in 2005, the retail price becomes significantly lower unil the second half of 2006. Since then, with a few exceptions, the retail price does not significantly differ from pre-shock levels. To see whether or not wholesalers have to pay the bill of these lower retail prices, we have to look at the impulse responses of the wholesale price itself to the same shocks that triggered the retail price impulse reponses in Figure 2. The ones of the wholesale price are displayed in Figure 3. As for the retail price, lower levels also show up for the wholesale price in 2005.

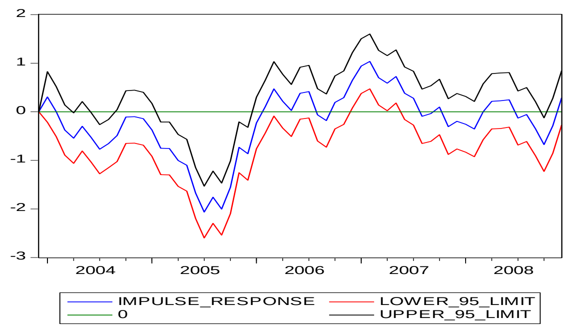

Figure 3. Monthly impulse responses wholesale price obtained by the sum of the wholesale price impulse responses after a negative one-unit shock in the wholesale price in December 2003 and the wholesale price impulse responses after a positive one-unit shock in the wholesale price in December 2003 (broilers in the US, responses in $ cents per pound, retail weight equivalent)

However, to assess the net result for the retail-wholesale price spread we should not only take into account the responses triggered by wholesale price shocks, but also those initiated by the positive and negative shocks in the retail price. Moreover, we have to consider the shocks that are representative. For this we can take the standard deviation of the errors of the reduced-form equations: the standard error of the residuals of equation (5), denoted σ_w, for the wholesale price, and the standard error of the residuals of equation (6), denoted σ_r*, for the retail price. As a consequence, the net retail price impulse responses can be computed as $\sigma_w(\beta_{rw0}{}^{(h)+} - \beta_{rw0}{}^{(h)-}) + \sigma_r{}^*(\beta_{rr1}{}^{(h)+*} - \beta_{rr1}{}^{(h)-*})$ and the net wholesale price responses are derived as $\sigma_w(\beta_{ww0}{}^{(h)+*} - \beta_{ww0}{}^{(h)-*}) + \sigma_r{}^*(\beta_{wr1}{}^{(h)+} - \beta_{wr1}{}^{(h)-})$. These net responses are presented in Figure 4 and Figure 5 and provide clear evidence that during the last quarter of 2004 and the whole of 2005 the wholesale price decreases significantly more than the retail price. On average the retail-wholesale price spread is 4.47 cents higher with a maximum of 8.68 cents in July 2005. On average the widening of the retail-wholesale price spread amounts to 6.57 per cent of the wholesale price and 2.57 per cent of the retail price level that period. To compare, the retail-wholesale price spread is 164 per cent of the wholesale price so that the widening of 6.57 per cent of the wholesale price seems ignorable. Nevertheless, in a sector characterised by a saturated market and price-inelastic consumer demand, which may explain the lower or non-significant net retail price responses, a 6.57 per cent margin on the wholesale price as extra profit for the retail stage vis-à-vis the wholesalers could well be quite considerable when compared to the assumed low profit margins in the broiler wholesale business.

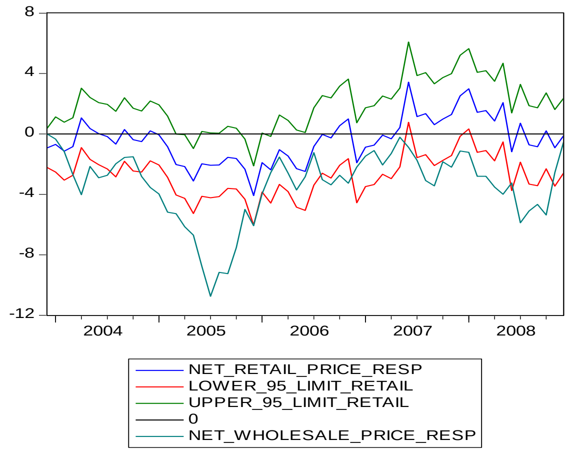

Figure 4. **Monthly impulse responses retail price (including its 95% confidence interval) and wholesale price after positive and negative one-standard deviation shocks in the wholesale and retail prices in December 2003 (broilers in the US, responses in $ cents per pound, retail weight equivalent)**

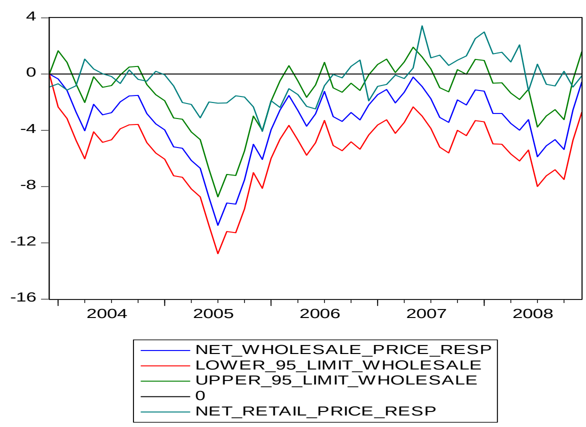

Figure 5. **Monthly impulse responses wholesale price (including its 95% confidence interval) and retail price after positive and negative one-standard deviation shocks in the wholesale and retail prices in December 2003 (broilers in the US, responses in $ cents per pound, retail weight equivalent)**

We now turn to the pork chain in the U.S. Three prices are considered: the farm price, the wholesale price and the retail price. The time series graphs of these prices are presented in Figure 6. As for the broiler price series the formal unit root and cointegration tests by Johansen (1995) do not find any cointegration and conclude that the series are integrated of order one such that the unconditional mean of each price in first-differences is equal to zero. Next, we examine the contemporaneous causal relationships according to which the first-differences of the retail price depend on the first-differences of the wholesale price and, in turn, the first-differences of the wholesale price depend on the first-differences of the farm price. In a schematic presentation this comes down to: $\Delta P_{f,t} \rightarrow$

Figure 6. Monthly farm, wholesale and retail prices of pork in the US in $ cents per pound, retail weight equivalent for the period January 1990 – December 2008

$\Delta P_{w,t} \rightarrow \Delta P_{r,t}$. To check that this causal ordering is compatible with the data, we follow Swanson and Granger (1997) in using the residuals from the VAR by which we tested for cointegration and order of integration, to perform the following regressions

Table 1a. Estimates of Equation (8a)

Dependent Variable: $u_{r,t}$; Method: OLS

Sample: 1990M01 2003M11; Included observations: 167

Variable	Coefficient	Std. Error	t-Statistic	Prob.
Intercept	1.34E-16	0.172628	7.78E-16	1.0000
$\hat{u}_{w,t}$	0.254899	0.096911	2.630239	0.0093
$\hat{u}_{f,t}$	-0.127044	0.075819	-1.675623	0.0957

R-squared	0.051826	Mean dependent var	7.98E-17
Adjusted R-squared	0.040263	S.D. dependent var	2.277157
S.E. of regression	2.230844	Akaike info criterion	4.460438
Sum squared resid	816.1731	Schwarz criterion	4.516450
Log likelihood	-369.4466	Hannan-Quinn criter.	4.483172
F-statistic	4.481990	Durbin-Watson stat	2.014959
Prob(F-statistic)	0.012730		

(8a) $\hat{u}_{r,t} = \delta_{rw}\hat{u}_{w,t} + \delta_{rf}\hat{u}_{f,t} + e_{r,t}$

(8b) $\hat{u}_{w,t} = \delta_{wf}\hat{u}_{f,t} + e_{w,t}$

where $\hat{u}_{r,t}$, $\hat{u}_{w,t}$ and $\hat{u}_{f,t}$ are the estimated residuals from the VAR for $\Delta P_{r,t}$, $\Delta P_{w,t}$ and $\Delta P_{f,t}$. The causal ordering $\Delta P_{f,t} \rightarrow \Delta P_{w,t} \rightarrow \Delta P_{r,t}$ complies with finding that δ_{rw} and δ_{wf} are significantly larger than zero, while δ_{rf} should be zero. The regression results, presented in Tables 1a and 1b, confirm these estimates, at least, at the 5 per cent significance level.

Table 1b. Estimates of Equation (8b)

Dependent Variable: $u_{w,t}$

Sample: 1990M01 2003M11; Included observations: 167

Variable	Coefficient	Std. Error	t-Statistic	Prob.
Intercept	-5.54E-16	0.138675	-3.99E-15	1.0000
$\hat{u}_{f,t}$	0.694398	0.028058	24.74897	0.0000

R-squared	0.787785	Mean dependent var	-2.39E-16
Adjusted R-squared	0.786498	S.D. dependent var	3.878417
S.E. of regression	1.792071	Akaike info criterion	4.016524
Sum squared resid	529.9003	Schwarz criterion	4.053865
Log likelihood	-333.3798	Hannan-Quinn criter.	4.031680
F-statistic	612.5116	Durbin-Watson stat	2.023429
Prob(F-statistic)	0.000000		

The following projection regressions are used to directly estimate the impulse responses. For the retail price we perform the regressions

$$(9)\quad \begin{aligned} P_{r,t+h} - P_{r,t-1} = {} & \alpha_r + \textstyle\sum_{s=1}^{11} \delta_{rs}D_{st} + \sum_{i=0}^{m_r}\{(\alpha_{rwi}{}^{(h)-} + \beta_{rwi}{}^{(h)-}\Delta P_{w,t-i})I(\Delta P_{w,t-i} \le 0) \\ & + \beta_{rwi}{}^{(h)+}\Delta P_{w,t-i}\, I(\Delta P_{w,t-i} > 0)\} + \textstyle\sum_{i=1}^{n_r}\{(\alpha_{rri}{}^{(h)-} + \beta_{rri}{}^{(h)-}\Delta P_{r,t-i})I(\Delta P_{r,t-i} \le 0) \\ & + \beta_{rri}{}^{(h)+}\Delta P_{r,t-i}\ I(\Delta P_{r,t-i} > 0)\} + \textstyle\sum_{i=1}^{l_r}\{(\alpha_{rfi}{}^{(h)-} + \beta_{rfi}{}^{(h)-}\Delta P_{f,t-i})I(\Delta P_{f,t-i} \le 0) \\ & + \beta_{rfi}{}^{(h)+}\Delta P_{f,t-i}\ I(\Delta P_{f,t-i} > 0)\} + u_{r,t+h} \end{aligned}$$

for $h = 0, 1, \ldots$ to obtain $-\beta_{rw0}{}^{(h)-}$ and $\beta_{rw0}{}^{(h)+}$ which are the retail price impulse responses after a negative and positive one-unit shock, respectively, in the wholesale price. To compute the retail price impulse responses after a shock in the retail price itself, we use the regressions

$$(10)\quad \begin{aligned} P_{r,t+h} - P_{r,t-1} = {} & \alpha_r{}^* + \textstyle\sum_{s=1}^{11} \delta_{rs}{}^*D_{st} + \sum_{i=1}^{m_r^*}\{(\alpha_{rwi}{}^{(h)-*} + \beta_{rwi}{}^{(h)-*}\Delta P_{w,t-i})I(\Delta P_{w,t-i} \le 0) \\ & + \beta_{rwi}{}^{(h)+*}\Delta P_{w,t-i}\, I(\Delta P_{w,t-i} > 0)\} + \textstyle\sum_{i=1}^{n_r^*}\{(\alpha_{rri}{}^{(h)-*} + \beta_{rri}{}^{(h)-*}\Delta P_{r,t-i})I(\Delta P_{r,t-i} \le 0) \\ & + \beta_{rri}{}^{(h)+*}\Delta P_{r,t-i}\ I(\Delta P_{r,t-i} > 0)\} + \textstyle\sum_{i=1}^{l_r^*}\{(\alpha_{rfi}{}^{(h)-*} + \beta_{rfi}{}^{(h)-*}\Delta P_{f,t-i})I(\Delta P_{f,t-i} \le 0) \\ & + \beta_{rfi}{}^{(h)+*}\Delta P_{f,t-i}\ I(\Delta P_{f,t-i} > 0)\} + u_{r,t+h}{}^* \end{aligned}$$

for $h = 1, 2, \ldots$ to estimate $-\beta_{rr1}{}^{(h)-*}$ and $\beta_{rr1}{}^{(h)+*}$ which are the retail price impulse responses after a negative and positive one-unit shock, respectively, in the retail price. Finally, for a shock in the farm price the following regressions are run

$$(11)\quad \begin{aligned} P_{r,t+h} - P_{r,t-1} = {} & \alpha_r{}^{**} + \textstyle\sum_{s=1}^{11} \delta_{rs}{}^{**}D_{st} + \sum_{i=1}^{m_r^{**}}\{(\alpha_{rwi}{}^{(h)-**} + \beta_{rwi}{}^{(h)-**}\Delta P_{w,t-i})I(\Delta P_{w,t-i} \le 0) \\ & + \beta_{rwi}{}^{(h)+**}\Delta P_{w,t-i} I(\Delta P_{w,t-i} > 0)\} + \textstyle\sum_{i=1}^{n_r^{**}}\{(\alpha_{rri}{}^{(h)-**} + \beta_{rri}{}^{(h)-**}\Delta P_{r,t-i})I(\Delta P_{r,t-i} \le 0) \\ & + \beta_{rri}{}^{(h)+**}\Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + \textstyle\sum_{i=0}^{l_r^{**}}\{(\alpha_{rfi}{}^{(h)-**} + \beta_{rfi}{}^{(h)-**}\Delta P_{f,t-i})I(\Delta P_{f,t-i} \le 0) \\ & + \beta_{rfi}{}^{(h)+**}\Delta P_{f,t-i}\, I(\Delta P_{f,t-i} > 0)\} + u_{r,t+h}{}^{**} \end{aligned}$$

for $h = 0, 1, \ldots$ to compute $-\beta_{rf0}{}^{(h)-**}$ and $\beta_{rf0}{}^{(h)+**}$ which are the retail price impulse responses after a negative and positive one-unit shock, respectively, in the farm price.

The impulse responses of the wholesale price are estimated by running the following regressions

(12a) $P_{w,t+h} - P_{w,t-1} = \alpha_w + \sum_{s=1}^{11} \delta_{ws} D_{st} + \sum_{i=1}^{m_w}\{(\alpha_{wwi}{}^{(h)-} + \beta_{wwi}{}^{(h)-}\Delta P_{w,t-i})\, I(\Delta P_{w,t-i} \le 0)$
$+ \beta_{wwi}{}^{(h)+}\Delta P_{w,t-i}\, I(\Delta P_{w,t-i} > 0)\} + \sum_{i=1}^{n_w}\{(\alpha_{wri}{}^{(h)-} + \beta_{wri}{}^{(h)-}\Delta P_{r,t-i}) I(\Delta P_{r,t-i} \le 0)$
$+ \beta_{wri}{}^{(h)+}\Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + \sum_{i=0}^{l_w}\{(\alpha_{wfi}{}^{(h)-} + \beta_{wfi}{}^{(h)-}\Delta P_{f,t-i})\, I(\Delta P_{f,t-i} \le 0)$
$+ \beta_{wfi}{}^{(h)+}\Delta P_{f,t-i}\, I(\Delta P_{f,t-i} > 0)\} + u_{w,t+h}$ for $h = 0, 1, \ldots$

(12b) $P_{w,t+h} - P_{w,t-1} = \alpha_w{}^* + \sum_{s=1}^{11} \delta_{ws}{}^* D_{st} + \sum_{i=1}^{m_w^*}\{(\alpha_{wwi}{}^{(h)-*} + \beta_{wwi}{}^{(h)-*}\Delta P_{w,t-i}) I(\Delta P_{w,t-i} \le 0)$
$+ \beta_{wwi}{}^{(h)+*}\Delta P_{w,t-i}\, I(\Delta P_{w,t-i} > 0)\} + \sum_{i=1}^{n_w^*}\{(\alpha_{wri}{}^{(h)-*} + \beta_{wri}{}^{(h)-*}\Delta P_{r,t-i}) I(\Delta P_{r,t-i} \le 0)$
$+ \beta_{wri}{}^{(h)+*}\Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + \sum_{i=1}^{l_w^*}\{(\alpha_{wfi}{}^{(h)-*} + \beta_{wfi}{}^{(h)-*}\Delta P_{f,t-i}) I(\Delta P_{f,t-i} \le 0)$
$+ \beta_{wfi}{}^{(h)+*}\Delta P_{f,t-i}\, I(\Delta P_{f,t-i} > 0)\} + u_{w,t+h}{}^*$ for $h = 1, 2, \ldots$

(12c) $P_{w,t+h} - P_{w,t-1} = \alpha_w{}^{**} + \sum_{s=1}^{11} \delta_{ws}{}^{**} D_{st} + \sum_{i=1}^{m_w^{**}}\{(\alpha_{wwi}{}^{(h)-**} + \beta_{wwi}{}^{(h)-**}\Delta P_{w,t-i}) I(\Delta P_{w,t-i} \le 0)$
$+ \beta_{wwi}{}^{(h)+**}\Delta P_{w,t-i} I(\Delta P_{w,t-i} > 0)\} + \sum_{i=2}^{n_w^{**}}\{(\alpha_{wri}{}^{(h)-**} + \beta_{wri}{}^{(h)-**}\Delta P_{r,t-i}) I(\Delta P_{r,t-i} \le 0)$
$+ \beta_{wri}{}^{(h)+**}\Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + \sum_{i=1}^{l_w^{**}}\{(\alpha_{wfi}{}^{(h)-**} + \beta_{wfi}{}^{(h)-**}\Delta P_{f,t-i}) I(\Delta P_{f,t-i} \le 0)$
$+ \beta_{wfi}{}^{(h)+**}\Delta P_{f,t-i}\, I(\Delta P_{f,t-i} > 0)\} + u_{w,t+h}{}^{**}$ for $h = 1, 2, \ldots$

to obtain the estimates $-\beta_{wf0}{}^{(h)-}$ and $\beta_{wf0}{}^{(h)+}$ from (12a) which are the wholesale price impulse responses triggered by a negative and positive one-unit shock, respectively, in the farm price, to obtain the coefficients $-\beta_{wr1}{}^{(h)-*}$ and $\beta_{wr1}{}^{(h)+*}$ from (12b) representing the impulse responses in the wholesale price generated by a negative and positive one-unit shock, respectively, in the retail price, and, to obtain the $-\beta_{ww1}{}^{(h)-**}$ and $\beta_{ww1}{}^{(h)+**}$ as the wholesale price impulse responses initiated by a negative and positive one-unit shock, respectively, in the wholesale price itself. Lastly, the impulse responses of the farm price are estimated by the regressions

(13a) $P_{f,t+h} - P_{f,t-1} = \alpha_f + \sum_{s=1}^{11} \delta_{fs} D_{st} + \sum_{i=1}^{m_f}\{(\alpha_{fwi}{}^{(h)-} + \beta_{fwi}{}^{(h)-}\Delta P_{w,t-i}) I(\Delta P_{w,t-i} \le 0)$
$+ \beta_{fwi}{}^{(h)+}\Delta P_{w,t-i} I(\Delta P_{w,t-i} > 0)\} + \sum_{i=1}^{n_f}\{(\alpha_{fri}{}^{(h)-} + \beta_{fri}{}^{(h)-}\Delta P_{r,t-i}) I(\Delta P_{r,t-i} \le 0)$
$+ \beta_{fri}{}^{(h)+}\Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + \sum_{i=1}^{l_f}\{(\alpha_{ffi}{}^{(h)-} + \beta_{ffi}{}^{(h)-}\Delta P_{f,t-i})\, I(\Delta P_{f,t-i} \le 0)$
$+ \beta_{ffi}{}^{(h)+}\Delta P_{f,t-i}\, I(\Delta P_{f,t-i} > 0)\} + u_{f,t+h}$ for $h = 1, 2, \ldots$

(13b) $P_{f,t+h} - P_{f,t-1} = \alpha_f{}^* + \sum_{s=1}^{11} \delta_{fs}{}^* D_{st} + \sum_{i=1}^{m_f^*}\{(\alpha_{fwi}{}^{(h)-*} + \beta_{fwi}{}^{(h)-*}\Delta P_{w,t-i})\, I(\Delta P_{w,t-i} \le 0)$
$+ \beta_{fwi}{}^{(h)+*}\Delta P_{w,t-i} I(\Delta P_{w,t-i} > 0)\} + \sum_{i=2}^{n_f^*}\{(\alpha_{fri}{}^{(h)-*} + \beta_{fri}{}^{(h)-*}\Delta P_{r,t-i}) I(\Delta P_{r,t-i} \le 0)$
$+ \beta_{fri}{}^{(h)+*}\Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + \sum_{i=1}^{l_f^*}\{(\alpha_{ffi}{}^{(h)-*} + \beta_{ffi}{}^{(h)-*}\Delta P_{f,t-i})\, I(\Delta P_{f,t-i} \le 0)$
$+ \beta_{ffi}{}^{(h)+*}\Delta P_{f,t-i}\, I(\Delta P_{f,t-i} > 0)\} + u_{f,t+h}{}^*$ for $h = 1, 2, \ldots$

(13c) $P_{f,t+h} - P_{f,t-1} = \alpha_f{}^{**} + \sum_{s=1}^{11} \delta_{fs}{}^{**} D_{st} + \sum_{i=2}^{m_f^{**}}\{(\alpha_{fwi}{}^{(h)-**} + \beta_{fwi}{}^{(h)-**}\Delta P_{w,t-i}) I(\Delta P_{w,t-i} \le 0)$
$+ \beta_{fwi}{}^{(h)+**}\Delta P_{w,t-i} I(\Delta P_{w,t-i} > 0)\} + \sum_{i=2}^{n_f^{**}}\{(\alpha_{fri}{}^{(h)-**} + \beta_{fri}{}^{(h)-**}\Delta P_{r,t-i}) I(\Delta P_{r,t-i} \le 0)$
$+ \beta_{fri}{}^{(h)+**}\Delta P_{r,t-i}\, I(\Delta P_{r,t-i} > 0)\} + \sum_{i=1}^{l_f^{**}}\{(\alpha_{ffi}{}^{(h)-**} + \beta_{ffi}{}^{(h)-**}\Delta P_{f,t-i}) I(\Delta P_{f,t-i} \le 0)$
$+ \beta_{ffi}{}^{(h)+**}\Delta P_{f,t-i}\, I(\Delta P_{f,t-i} > 0)\} + u_{f,t+h}{}^{**}$ for $h = 1, 2, \ldots$

where $-\beta_{fr1}{}^{(h)-}$ and $\beta_{fr1}{}^{(h)+}$ in (13a) are the farm-price impulse responses after a negative and positive one-unit shock, respectively, in the retail price, $-\beta_{fw1}{}^{(h)-*}$ and $\beta_{fw1}{}^{(h)-*}$ in (13b) are the farm-price impulse responses triggered by a negative and positive one-unit shock, respectively, in the wholesale price and, finally, $-\beta_{ff1}{}^{(h)-**}$ and $\beta_{ff1}{}^{(h)+**}$ are the farm-price impulse responses initiated by a negative and positive one-unit shock, respectively, in the farm price itself.

Like Figure 4 and Figure 5 for the broiler sector, with respect to the pork chain Figures 7-9 present the net impulse responses for the retail price (= $\sigma_w{*}(\beta_{rw0}{}^{(h)+} - \beta_{rw0}{}^{(h)-})$ $+\sigma_r{*}(\beta_{rr1}{}^{(h)+}{*}-\beta_{rr1}{}^{(h)-}{*})+\sigma_f\,(\beta_{rf0}{}^{(h)+}{**}-\beta_{rf0}{}^{(h)-}{**})$), wholesale price (=$\sigma_f\,(\beta_{wf0}{}^{(h)+}-\beta_{wf0}{}^{(h)-})$ $+\sigma_r{*}(\beta_{wr1}{}^{(h)+}{*}-\beta_{wr1}{}^{(h)-}{*})+\sigma_w{*}(\beta_{ww1}{}^{(h)+}{**}-\beta_{ww1}{}^{(h)-}{**})$) and farm price (=$\sigma_r{*}(\beta_{fr1}{}^{(h)+}-\beta_{fr1}{}^{(h)-})$ $+\sigma_w{*}(\beta_{fw1}{}^{(h)-}{*}-\beta_{fw1}{}^{(h)-}{*})+\sigma_f\,(\beta_{ff1}{}^{(h)+}{**}-\beta_{ff1}{}^{(h)-}{**})$), where the representative shocks are the standard deviations $\sigma_r{*}$, $\sigma_w{*}$ and σ_f of the residual terms in the equations (10), (12b) and (13a), respectively. According to Figures 7-9 all three prices become significantly higher, the retail price during the years 2004-2006 and the farm and wholesale prices during most of the years 2005-2006. Furthermore, in Figure 7 we see that the retail-wholesale price spread becomes significantly higher after one year and stays significantly so during the years 2005 and 2006. In contrast, the wholesale-farm price spread does not significantly change.The average widening of the retail-wholesale price spread during the years 2005 and 2006 amounts to 12.35 cents per pound, which is 4.38 per cent of the retail price, 10.06 per cent of the wholesale price and 14.58 per cent of the farm price. For comparison, the retail-wholesale price spread is, on average, 129 per cent of the whole-sale price and the wholesale-farm price spread amounts to 44 per cent of the farm price. Although these percentages are much higher than those of the widening of the retail-wholesale price spread, it ultimately depends on the profit margin as a percentage of the price to know how attractive this extra retail margin is for each of the stages in the US pork chain.

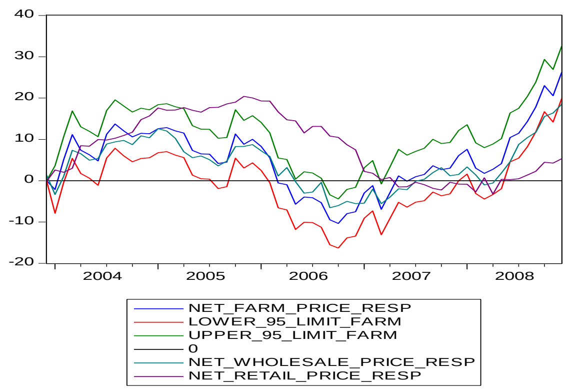

Figure 7. Impulse responses farm price (including its 95% confidence interval), wholesale price and retail price after positive and negative one-standard deviation shocks in all three prices in December 2003 (pork in the US, responses in $ cents per pound, retail weight equivalent)

4. Conclusions

In this paper we have set out Jorda's (2005) method of local projections by which nonlinear impulse responses can be computed without the need to specify and estimate the underlying nonlinear dynamic system. The method is used to compute price reaction functions that show how the prices of the different stages in the supply chain dynamically respond to one another and whether or not these responses reveal any asymmetric patterns. Empirical applications for the US pork-meat and broiler composite chains illustrate the convenience of the method. Triggered by a negative and positive one-unit shock in the retail price and wholesale price simultaneously in December 2003, we find evidence for the US broiler chain that during the last quarter of 2004 and the whole of 2005 the wholesale price decreases significantly more than the retail price. A simultaneous negative and positive one-unit shock in the retail, wholesale and farm prices in December 2003 reveals that for the US pork chain the retail-wholesale price spread becomes significantly higher after one year and stays significantly so during the years 2005 and 2006, whereas the wholesale-farm price spread does not significantly change.

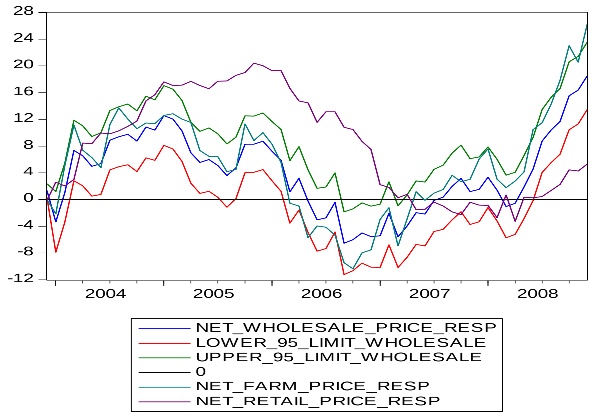

Figure 8. **Impulse responses wholesale price (including its 95% confidence interval), farm price and retail price after positive and negative one-standard deviation shocks in all three prices in December 2003 (pork in the US, responses in $ cents per pound, retail weight equivalent)**

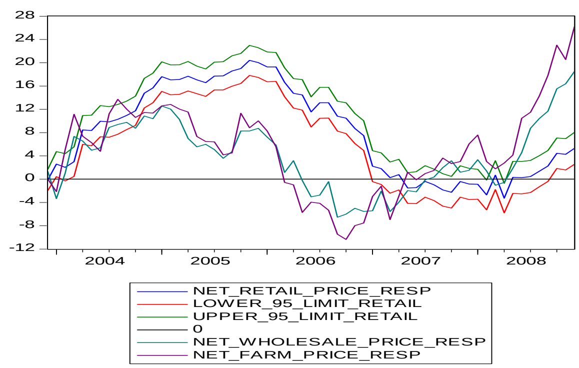

Figure 9. **Impulse responses retail price (including its 95% confidence interval), wholesale price and farm price after positive and negative shocks in all three prices in December 2003 (pork in the US, responses in $ cents per pound, retail weight equivalent)**

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