Basnet *et al. Agricultural and Food Economics* (2023) 11:50
https://doi.org/10.1186/s40100-023-00290-7

Agricultural and Food Economics

RESEARCH

Open Access

The distortion in the EU feed market due to import constraints on genetically modified soy

Shyam Kumar Basnet[1,4*], Ranjan Kumar Ghosh[2], Mattias Eriksson[3] and Carl-Johan Lagerkvist[4]

*Correspondence: shyam.basnet@yahoo.com

[1] Stockholm Resilience Centre, Stockholm University, Stockholm, Sweden
[2] Center for Management in Agriculture, Indian Institute of Management Ahmedabad (IIMA), Vastrapur, Ahmedabad 380015, India
[3] Department of Energy and Technology, Swedish University of Agricultural Sciences, Uppsala, Sweden
[4] Department of Economics, Swedish University of Agricultural Sciences, Uppsala, Sweden

Abstract

Feed importers in some EU member states face constraints on imports of genetically modified (GM) soy, a practice that may compromise the interests of EU livestock farmers. Using the cases of Sweden and Austria, we analyzed price transmission in the soy supply chain originating from Brazil, applying an asymmetric non-linear auto-regressive distributed lag (ARDL) model to identify short-run and long-run asymmetries. The results revealed significant asymmetric effects in how positive and negative price changes are absorbed within the feed industry. Notably, increases in the cost of Brazilian soy swiftly affect the prices for EU farmers, while cost reductions fail to trigger corresponding price decreases. Consequently, stronger constraints on GM soy imports are likely to exacerbate the competitiveness challenges faced by livestock farmers, primarily due to their reliance on non-GM soy. This implies that the restrictions on GM imports need to be relaxed or that low-cost local protein alternatives need to be developed.

Keywords: Soy supply, Price transmission, ARDL model, Genetically modified

JEL Classification: C530, Q020, Q160

Introduction

The livestock sector contributes nearly 45% of total agricultural output in the European Union (EU). To produce 150 million tons of compound feed for the EU livestock sector, 28 million tons of soymeal are required. All this is imported directly as soymeal from exporting countries or as whole soybeans that are crushed locally (Tillie and Rodríguez-Cerezo 2015), with more than 70% originating from Brazil and Argentina alone (Kroes and Kuepper 2015; Kuepper and Stravens 2022). There are concerns among EU policy-makers that high import dependency can expose the livestock sector to trade distortions and price volatility, leading to feed cost escalation (de Boer et al. 2014). This in turn can increase the production costs for livestock farmers and reduce the overall profitability of the sector.

Moreover, EU regulations require food and feed products containing more than 0.9% authorized genetically modified (GM) products to be labeled and traced, essentially resulting in segregated supply chains (European Commission 2003). The accidental presence of traces of unauthorized GM products often leads to the rejection of shipments, imposing additional costs on feed importers. According to the European Commission (2007) and Kalaitzandonakes et al. (2014), segregation costs, including contamination risks, could drive feed costs up by 300%.

Most of the studies on feed trade economics focuses on the impacts of GM trade restrictions on the exporting countries (Henseler et al. 2013; Kalaitzandonakes et al. 2014; Smith and Katovich 2017). For instance, Smith and Katovich (2017), strong GM restrictions in importing regions such as the EU have a higher negative impact on Argentinian trade than on Brazilian trade, as Argentina has been continually increasing the GM content in its soy and maize products. A study by Kalaitzandonakes et al. (2014) analyzing structural responses to asynchronicity, such as changes in trade patterns, feedstuff substitution and adjustments in primary production, found that under asynchronous approval, there will be unavoidable trade disruptions given the high costs of segregation in exporting countries. The options would then be to relax tolerance limits or to produce GM feedstuff domestically. In both cases the supply chains would have to be segregated, bringing particular costs (Desquilbet and Bullock 2009; Kalaitzandonakes 2011). Henseler et al. (2013) analyzed the impacts of potential interruptions in soy exports from exporting countries and showed that, under asynchronous approval, shipments may be rejected where there is a zero tolerance policy for non-EU-approved GM material.

On the other hand, very few studies have analyzed the impacts of soy trade constraints on the *importin*g country's feed markets and farmers. Recently, in the EU the price of imported GM-free soy has increased faster than the rate of a cryptocurrency (Jordbruksaktuellt 2021). This was primarily driven by increased demand for GM free feed in large livestock producing countries such as France, Germany and Denmark. The prices for livestock products, on the other hand, did not increase significantly. Yet, to the best of our knowledge, there has never been an examination of such disruptions, especially under conditions of import overdependence and GM restrictions.

We fill this gap by analyzing the data from Sweden and Austria, as they have adopted differing positions on the segregation of non-GM and GM soy. Sweden has *zero tolerance* for GM soy feed, with a 100% share of non-GM identity preserved in total soybean imports (Tillie and Rodríguez-Cerezo 2015). Their prices were between 5 and 35% higher than those of GM soy during 2004–2014. In contrast, a significant proportion of soymeal imports to Austria are GM soy, with only 21.8% of total soybean meal equivalent imports being non-GM identity preserved and segregated (Tillie and Rodríguez-Cerezo 2015). Since GM soymeal is the only major agricultural biotech commodity on the Austrian market, the Austrian government has even formed a task force to identify possible scenarios for the co-existence of GM, conventional non-GM and organically produced soymeal (Krautgartner 2017).

We estimated the price transmission elasticity in the soy supply from Brazil to compound feeds in Austria and Sweden. An auto-regressive distributed lag (ARDL) model was used to investigate short- and long-run asymmetries in the transmission of prices and their adjustments on the EU feed market. Using the price premium for GM soy available for other EU Member States, we simulate the potential impacts of introducing GM soy into Swedish feed markets. This helps to understand the extent of price distortions related to the segregation of non-GM and GM soy and its impact in the feed market disruptions. For robust simulated findings, we use Austrian and UK price premiums for the potential prices of GM soy in Sweden. These results, based on simulated prices of GM soy in Sweden, were corroborated by the observed co-existence of GM and conventional non-GM soy markets in Austria. The new novel contributions of this work were to establish a causal relationship between prices in originating countries and end-user compound feeds in the EU, and to determine the potential impact of GM soymeal on transmission of the upstream cost advantage in producing countries (e.g., Brazil) to livestock farmers in countries that rely solely on non-GM soy (e.g., Sweden).

In the following sections, we review the literature on asymmetric price transmission and describe the empirical estimation methodology. Subsequently, we explain the data, present the empirical results related to price transmission elasticities, and discuss key findings and important policy perspectives.

Literature review on asymmetric price transmission

The analysis of pricing behavior within supply chains is essential for understanding market efficiency and economic dynamics. Competitive markets generally assume that prices will eventually equalize across domestic markets in the long-run, though short-term deviations are expected. In contrast, the presence of market dominance within non-competitive markets can lead to a gradual and potentially incomplete transmission of prices within the supply chain. Asymmetric price transmission is a common occurrence, typically driven by the pricing strategies adopted by market leaders. In most cases, market leaders are inclined to transmit price increases more rapidly downstream in the supply chain than they are to pass on price decreases. This price transmission asymmetry is rooted in the exploitation of perceived market dominance. Market concentration and the scarcity of alternative products can exacerbate this behavior (Serra and Goodwin 2003). Empirical evidence underscores the role of market power as a catalyst for imperfections in price transmission (Meyer and Cramon-Taubadel 2004; Assefa et al. 2014; Barboza et al. 2022; Nakajima 2011).

Additionally, the price adjustment process itself can contribute to asymmetric price transmission. Decreasing prices may entail a slower adjustment process, often due to perceived high costs associated with price reductions. Multiple factors, extending beyond market power, influence this process, including market concentration, price adjustment complexities, information asymmetry, and structural barriers (Serra and Goodwin 2003; Weldegebriel 2004; Taslim and Hossain 2015; Barboza et al. 2022; 2023; Kamyabi and Chidmi 2023). Importantly, these factors are intricately linked with the challenges inherent in price adjustment, stemming from underlying disparities in adjustment costs. The interplay among these factors explains delayed price adjustments, resulting in market inefficiencies and asymmetric responses to price fluctuations.

Recent research has emphasized the complexity of price transmission dynamics and the nuances of asymmetric price adjustments (Margarido et al. 2007; Taslim and Hossain 2015; Gizaw et al. 2021; Barboza et al 2022; 2023; Kamyabi and Chidmi 2023). These studies distinguish between positive and negative price changes, examining the extent of asymmetries to gauge market imperfections. Such imperfections can arise from market dominance and the high costs associated with price adjustments. For instance, Kamaruddin et al. (2021) suggests that long-term price transmission asymmetry may be attributed to market power; while, short-term asymmetric price transmission may be influenced by underlying adjustment costs.

This study focuses on the EU's soybean market, with specific attention to Austria and Sweden. These countries show distinctive characteristics, including a high concentration of non-GM soy, structural barriers related to GM soy imports, and the co-existence of non-GM and GM traits, all of which introduce complexities into price transmission. In Sweden, the concentrated soy market may confer market dominance upon a few non-GM soy suppliers; while, in Austria GM soy may serve as an alternative to non-GM soy. In this context, the interplay between market concentration in non-GM soy and the significant costs involved in segregating non-GM and GM traits can lead to transitory disequilibria in prices. Within the broader framework of the EU soybean market, these unique factors—market concentration, structural barriers, and the presence of GM traits—introduce additional layers of complexity to the issue of price transmission. By exploring these complexities, this study contributes to our broader understanding of asymmetric price transmission and its implications for market efficiency and policy considerations within the soybean supply chain.

Methodology

Previous studies have examined the causal relationship between the prices of livestock products and compound feeds, and/or the prices of feeds and their ingredients for the different actors in the supply chain (Peeters and Surry 1997; Zhou and Koemle 2015). Since our research focus was on price transmission in the spatial dimension, from producers in one part of the world to consumers in another part, we estimated a one-way effect of price relationship. We modeled the price transmission of Brazilian soymeal to compound feed in Sweden and Austria in two stages. In the first stage, we assumed that the price of soymeal in Brazil directly affects the soymeal price in the EU as:

$$lnSoyNonGM_{cc,t} = \alpha_{1,cc} + \sum\nolimits_{pc} \beta_{pc} lnSoyNonGM_{pc,t} \sum\nolimits_{k} \psi_k ln\boldsymbol{z}_{k,cc} + e_t \tag{1}$$

where α_1 is a constant term, $SoyNonGM_t$ stands for the monthly price of non-GM soymeal at time t, the vector $\boldsymbol{z}$ represents various covariates, including prices for GM soy, seasonal dummy variables and a structural break indicator. The term e_t corresponds to an independent and identically distributed (*iid*) random error term. The index k refers to the elements in the covariate set, the index *cc* denotes the soymeal-consuming countries studied (here Sweden and Austria), and the index *pc* indicates the soybean-producing countries (here Brazil and Argentina).

We focused on Brazil and Argentina as soy producers because they have a strong comparative advantage in the GM industry and together export around 90% of South American GM soy (Smith and Katovich 2017). The magnitude of the parameters in Eq. (1) reflects elasticities or ratios of percentage changes. These parameters are comparable across the reference countries (Austria and Sweden) and can be interpreted as revealing the extent to which variables affect price transmission in terms of percentage changes.

On receiving imported soymeal, local feed companies find the optimal mix of ingredients, which determines the market price of compound feeds in Sweden and Austria. The price of compound feeds also depends, but to a lesser extent, on the price of protein ingredients (e.g., soymeal and rapeseed cake) and other cereal ingredients (e.g., maize, wheat and barley). However, soymeal is not replaceable with rapeseed cake, which contains only 20–22% protein, compared with > 70% in soymeal (Cederberg et al. 2009). Thus, in the second stage of price transmission, we considered only the effect of soymeal on compound feed prices. Assuming a perfect substitution between GM and non-GM soymeal, we modeled the transmission of soymeal prices to compound feeds as:

$$lnFeed_{cc,t} = \alpha_{2,cc} + \beta_1 lnSoyNonGM_{cc,t} + \textstyle\sum_k \gamma_k ln\boldsymbol{w}_{cc,k} + \varepsilon_t \tag{2}$$

where α_2 is a constant term, $Feed_t$ is the monthly price of compound feeds for poultry and pigs, the vector $\boldsymbol{w}$ represents the prices for GM soy and cereal ingredients such as wheat, barley and maize, k is the index for corresponding ingredients, and ε_t is the *iid* error.

To measure the asymmetric influences of $SoyNonGM_{pc}$ in Eq. (1) and $SoyNonGM_{cc}$ in Eq. (2), as in Shin et al. (2014), we decomposed the movement of $SoyNonGM_{pc}$ and $SoyNonGM_{cc}$ into positive and negative partial sums as $x_t = x_0 + x_t^+ + x_t^-$, where x_t^+ and x_t^- denote the partial sum processes of positive and negative changes in $lnSoyNonGM_{pc}$ and $lnSoyNonGM_{cc}$. The variables for partial summations are calculated as:
$x_t^+ = \sum_{i=1}^t \Delta x_i^+ = \sum_{i=1}^t \max(\Delta x_i, 0)$, and
$x_t^- = \sum_{i=1}^t \Delta x_i^- = \sum_{i=1}^t \min(\Delta x_i, 0)$.

By associating the positive and negative partial sums to the ARDL model yields the following asymmetric error correction model (AECM) for Eq. (1) (see Appendix A5 for the ARDL model in error correction form):

$$\begin{aligned}
&\Delta lnSoyNonGM_{cc,t}\\
&\quad = \alpha_{1,cc} + \sum_{pc}\sum_{i}^{q} \beta_{pc,i}^{+} \Delta lnSoyNonGM_{pc,t-i}^{+} + \sum_{pc}\sum_{i}^{q} \beta_{pc,i}^{-} \Delta lnSoyNonGM_{pc,t-i}^{-}\\
&\quad\quad + \sum_{k} \gamma_k \Delta lnz_k + \psi_{pc}^{+} lnSoyNonGM_{pc,t-1}^{+} + \psi_{pc}^{-} lnSoyNonGM_{pc,t-1}^{-}\\
&\quad\quad + \sum_{k} \psi_k lnz_k + e_{cc,t}
\end{aligned} \tag{3}$$

where the index i refers to the lag length. The coefficients β^+ and β^-, as well as ψ^+ and ψ^- measures the short- and long-run effects of positive and negative changes in $lnSoyNonGM_{pc}$, respectively.

The AECM for Eq. (2) can be written as:

$$\begin{aligned}
&\Delta lnFeed_{f,cc,t}\\
&\quad = \alpha_{2,f,cc} + \sum_{i=0}^{s} \beta_{f,cc,i}^{+} \Delta lnSoyNonGM_{cc,t-i}^{+} + \sum_{i=0}^{s} \beta_{f,cc,i}^{-} \Delta lnSoyNonGM_{cc,t-i}^{-}\\
&\quad\quad + \sum_{k}\sum_{i=1}^{l} \gamma_{f,cc,k,i} \Delta ln\boldsymbol{w}_{cc,k,t-i} + \psi_{pc}^{+} lnSoyNonGM_{cc,t-1}^{+}\\
&\quad\quad + \psi_{pc}^{-} lnSoyNonGM_{cc,t-1}^{-} + \sum_{k} \psi_{f,cc,k} ln\boldsymbol{w}_{cc,t-1} + \varepsilon_{f,cc,t}
\end{aligned} \tag{4}$$

where α_2 is a constant term and the index f refers to the price of compound feeds for poultry and pig.

The asymmetric ARDL approach follows a three-step estimation approach. First, we estimate the empirical specifications outlined in Eqs. (3) and (4) using the regressors x_t, which are decomposed into x_t^+ and x_t^-, in addition to the covariate $\boldsymbol{z}$ and $\boldsymbol{w}$. Secondly, we test the existence of a long-run relationship between the levels of the variables y_t, x_t^+ and x_t^- by testing the null hypothesis $H_0 : \psi = \psi^+ = \psi^- = 0$ using the bounds-testing procedure suggested by Pesaran et al. (2001). Finally, we applied the Wald test to examine two aspects: (i) long-run symmetry where $\psi^+ = \psi^-$, and (ii) short-run symmetry in which $\beta^+ = \beta^-$ of $SoyNonGM_{pc}$ in Eq. (5), and for $SoyNonGM_{cc}$ in Eq. (6).

Data

As input for Eqs. (1) and (2), we obtained monthly price data on compound pig and poultry feeds and their ingredients (soymeal, maize, wheat, barley, oats and rye), from the Swedish Board of Agriculture for the period January 1995 to December 2016 in the case of Sweden and from *AgarMarkt* in the case of Austria. We collected prices of compound feeds for poultry and pigs, as high-protein soybean products in the EU are mainly feed to poultry and pigs. However, we omitted the variables for maize, oats and rye in the econometric analysis, as we do not have complete data on the monthly prices for these commodities. Information on the monthly price of non-GM soymeal in Brazil and Argentina was taken from the Wageningen Economic Research database (Agrimatie 2020).

As can be seen in Fig. 1, soymeal prices in Brazil and Argentina and soymeal and compound pig and poultry feed prices in Austria and Sweden followed similar trends over the study period (1998–2016), indicating the possible existence of a price relationship in the EU soy supply chain from Brazil to the feed industry in Austria and Sweden. Some European countries, such as Austria, France, Germany and the United Kingdom (UK), have already begun to import GM soymeal. We obtained monthly price data for GM soymeal in these countries from the European Commission's Joint Research Council (JRC). For France and Germany, price data on GM soymeal are available only for short periods (2007–2012 and 2009–2014, respectively). For Austria and the UK, price data are available for a longer period (2004–2015), so we used those data to analyze the distortion effect of the import restrictions on GM soymeal in Sweden.

As Table 1 shows, soymeal was cheaper in Brazil than in Argentina over the study period, which could be one of the reasons why a large share of soymeal imported into the EU comes from Brazil. Brazil also has a lower level of adoption of GM soy varieties compared with Argentina. In Europe, the price of soymeal was considerably higher in Sweden than in other countries (Table 1), mainly because of imports of only non-GM soymeal. Similarly, cereal ingredients such as barley and wheat were

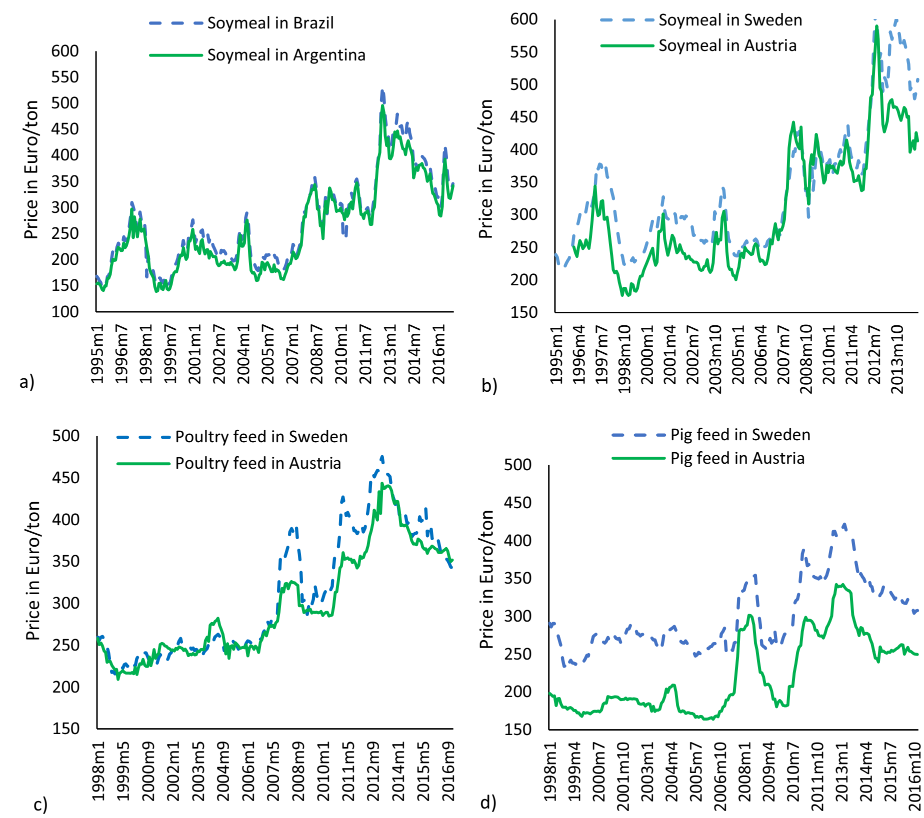

Fig. 1 Monthly price (Euro/ton) of **a** soymeal in Brazil and Argentina and **b** soymeal, **c** compound poultry feed and **d** compound pig feed in Sweden and Austria, 1998–2016. Source: authors' calculations.

Table 1 Variable abbreviations and description of prices for non-GM and GM soymeal, feed cereals and compound feeds for pigs and poultry

Variable (Euro/ton)	Price of:	Mean	Minimum	Maximum	Std.Dev
$SoyNonGM_{Brz}$	Non-GM soymeal in Brazil	303.50	238.71	376.20	29.41
$SoyNonGM_{Arg}$	Non-GM soymeal in Argentina	307.10	267.45	372.25	22.52
$SoyNonGM_{Swd}$	Non-GM soymeal in Sweden	387.68	337.80	450.80	23.03
$SoyNonGM_{AT}$	Non-GM soymeal in Austria	360.08	315.00	415.02	24.50
$SoyNonGM_{UK}$	Non-GM soymeal in the UK	377.32	335.00	472.00	34.43
$SoyGM_{AT}$	GM soymeal in Austria	319.70	282.00	370.00	19.72
$SoyGM_{UK}$	GM soymeal in the UK	335.62	304.00	412.00	21.87
$Barl_{Swd}$	Barley in Sweden	124.15	81.98	199.47	31.90
$Barl_{AT}$	Barley in Austria	117.62	74.30	216.10	33.71
$Wheat_{Swd}$	Wheat in Sweden	143.50	95.64	246.67	39.57
$Wheat_{AT}$	Wheat in Austria	120.71	65.90	218.43	38.65
$Feed(poultry)_{Swd}$	Poultry feed in Sweden	311.76	215.65	475.39	74.99
$Feed(poultry)_{AT}$	Poultry feed in Austria	297.26	209.31	443.60	63.65
$Feed(pig)_{Swd}$	Pig feed in Sweden	299.76	230.01	421.78	47.51
$Feed(pig)_{AT}$	Pig feed in Austria	221.99	163.90	342.11	50.62

Note: Prices for non-GM and GM soymeal are the nominal values presented only for the period January 2009–April 2012, as information is lacking for Germany and France beyond that point. For compound feeds and cereal ingredients, prices refer to the period January 1998-November 2016. Source: authors' calculations also more expensive in Sweden than in Austria. Generally speaking, the price of non-GM soymeal was higher than that of GM soymeal, so Austria experienced lower prices for compound feeds than Sweden (Table 1).

The empirical model

Following Goodwin and Holt (1999), Karantininis et al. (2011) and Shin et al. (2014), we applied the ARDL approach to estimate the price relationship in the EU soy supply chain. The ARDL model does not rely on the assumption of different orders of integration of the model variables. It possesses the capability to incorporate variables with varying optimal lag lengths, and it can simultaneously provide insights into both short- and long-run effects of exogenous variables, (see Meyer and Cramon-Taubadel 2004; Frey and Manera 2007). As a result, this model finds application in the analysis of supply constraints associated with both non-GM and GM soy.

To estimate the first stage price transmission process, we empirically estimate the AECM model for Eq. (3) as follows:

$$
\begin{aligned}
&\Delta lnSoyNonGM_{cc,t}\\
&\quad= \alpha_{1,cc} + \sum_{i=1}^{p} \rho_{cc,i} \Delta lnSoyNonGM_{cc,t-i} + \sum_{pc} \sum_{i}^{q} \beta^{+}_{Brz,i} \Delta lnSoyNonGM^{+}_{Brz,t-i}\\
&\qquad + \sum_{i}^{r} \beta^{-}_{Brz,i} \Delta lnSoyNonGM^{-}_{Brz,t-i} + \sum_{i}^{s} \beta_{Arg,i} \Delta lnSoyNonGM_{Arg,t-i}\\
&\qquad + \sum_{i}^{s} \beta^{GM}_{i} \Delta lnSoyGM_{cc,t-i} + \sum_{m} \delta_{cc,m} \Delta SD_{cc,m} + d_{1,cc} \Delta DU_{cc} + d_{2,cc} \Delta DB_{cc}\\
&\qquad + \psi_{1,cc} lnSoyNonGM_{cc,t-1} + \psi^{+}_{Brz} lnSoyNonGM^{+}_{Brz,t-1}\\
&\qquad + \psi^{-}_{Brz} lnSoyNonGM^{-}_{Brz,t-1} + \sum_{i=1}^{l} \psi_{Arg} lnSoyNonGM_{Arg,t-i}\\
&\quad= \sum_{i=1}^{n} \psi^{GM}_{cc,i} lnSoyGM_{cc,t-i} \sum_{i=1} \psi_{pc} lnSoyNonGM_{pc,t-1} + \sum_{m} \psi^{SD}_{cc,m} SD_{cc,m}\\
&\qquad + \psi^{DU}_{cc} \Delta DU_{cc} + \psi^{DB}_{cc} \Delta DB_{cc} + e_{cc,t}
\end{aligned}
\tag{5}
$$

where α_1 is a constant term, *SoyGM* stands GM soy prices, *SD* denotes seasonal dummies, *DB* and *DU* are dummies for a one-time structural change and a complete shift in intercept after a break date, respectively (defined mathematically as $DB = 1$ if $t = T_B + 1$ and 0 otherwise, and $DU = 1$ for $t \geq T_B$ and 0 otherwise, where T_B is the break date), index i indicates the lag length, and m refers to month for seasonal drift.

In Eq. (5), the short-run effects are captured by the coefficient of first-difference variables (i.e., ρ and β), while ψ_{pc} measures the long-run effects of all variables. In this set-up, we tested a null hypothesis of no co-integration relationship between the variables (i.e., $\psi_1 = \psi_{pc} = 0 \ \forall \ pc$) using a standard F-test proposed by Pesaran et al. (2001).

Similarly, we defined the ARDL model for second stage price transmission as:

$$
\begin{aligned}
&\Delta lnFeed_{f,cc,t} \\
&= \alpha_{2,f,cc} + \sum_{i=1}^{r} \rho_{f,cc,i} \Delta lnFeed_{f,cc,t-i} + \sum_{i=0}^{s} \beta^{+}_{f,cc,i} \Delta lnSoyNonGM^{+}_{cc,t-i} \\
&\quad + \sum_{i=0}^{s} \beta^{-}_{f,cc,i} \Delta lnSoyNonGM^{-}_{cc,t-i} + \sum_{i=1}^{s} \beta^{GM}_{f,cc,i} \Delta lnSoyGM_{cc,t-i} \\
&\quad + \sum_{k} \sum_{i=1}^{l} \gamma_{f,cc,k,i} \Delta ln\boldsymbol{w}_{cc,k,t-i} + \sum_{m} \delta_{f,cc,m} \Delta SD_{cc,m} + d_{1,f,cc} \Delta DU_{cc} \\
&\quad + d_{2,f,cc} \Delta DB_{cc} + \psi_{1,f,cc} lnFeed_{f,cc,t-1} + \psi^{+}_{f,cc} lnSoyNonGM^{+}_{cc,t-1} \\
&\quad + \psi^{-}_{f,cc} lnSoyNonGM^{-}_{cc,t-1} + \psi^{GM}_{f,cc} lnSoyGM_{f,cc,t-1} + \sum_{k} \gamma_{f,cc,k} ln\boldsymbol{w}_{cc,t-1} \\
&\quad + \sum_{m} \psi^{SD}_{f,cc,m} SD_{cc,m} + d^{DU}_{f,cc} DU_{cc} + d^{DB}_{f,cc} DB_{cc} + \varepsilon_{f,cc,t}
\end{aligned}
\tag{6}
$$

where α_2 is a constant term, the index f refers to the price of compound feeds for poultry and pig, the vector $\boldsymbol{w}$ represents the prices for cereal ingredients such as wheat, barley and maize.

As in Eq. (5), the standard F-test was also applied to Eq. (6) to test the null hypothesis of no co-integration relationship in the long-run (i.e., $\psi_1 = \psi^{NonGM} = \psi^{GM} = \gamma_k = 0$). In a reference scenario, we set $\beta^{GM} = 0$ and $\psi^{GM} = 0$, to reflect zero acceptance for GM soymeal in Sweden, but in an alternative scenario we relaxed this assumption.

In the above set-up, we used Akaike Information Criterion (AIC) to find the optimal lag length. Diagnostic tests, such as the CUSUM (cumulative sum of recursive residuals) test proposed by (Pesaran and Shin 1998), were used to examine the stability of estimated parameters in the long-run. Autocorrelation can be a serious issue in time-series variables, so we used the Durbin–Watson (DW) test to examine the severity of autocorrelation.

Before estimating the price transmission relationship, we checked the stationarity of the variables in the dataset. As we used monthly data with trend stationarity, as suggested by Lee and Strazicich (2004), we tested the unit root with a structural break. We found that the Zivot and Andrews (2002) model with a structural break in intercept was relevant for most variables (see Appendix A1). The structural breaks were detected on different dates, but a few dependent variables such as prices for non-GM soy and compound feeds had the approximate date of January 2009 for Sweden and May 2007 for Austria. Introduction of the EU CAP Health Check Reform in 2008 could explain the structural shift around that time in price variables associated with the soy supply chain (Grant 2008). As expected, we also detected seasonal drifts in the supply of soymeal for February and June in Sweden, and March and June in Austria. In general, the European market receives new harvests of rapeseeds in June–July to supply the protein feed, which can influence soymeal prices. In late autumn, soybean harvests in Europe and Canada also have an impact on EU prices for non-GM soymeal (Jordbruksaktuellt 2021). For comparison, we defined the same set of explanatory variables, including these structural breaks and seasonal dummies, in the reference and alternative scenarios.

Scenario development

To measure the effects of import constraints for GM soy on the European feed market, we defined two scenarios: (i) a reference scenario reflecting business-as-usual (BAU) conditions in Sweden without an alternative to non-GM soy for feed production, and (ii) an alternative scenario with co-existence of GM and non-GM soymeal, where we assessed the impacts of GM soymeal on price transmission to compound feeds in Sweden.

As global adoption of GM soy varieties is expanding rapidly, disruptions in GM soy imports could harm EU livestock farmers and consumers. The European Commission has recognized this risk and has provided general guidelines for nationalizing the approval process in line with the Single European Market principle, via co-existence regulation. This scenario, with changes in market structure, creates a need for analyzing the market dominance of non-GM soy and assessing the price adjustment process in the EU feed market, particularly in Sweden.

Since Sweden does not import GM soy, we generated its counterfactual prices by adding the Austrian price premium for GM soymeal to the Swedish non-GM soy prices. The price of non-GM soymeal follows a similar trend in Sweden and Austria (Fig. 1b), confirming the appropriateness of using the Austrian reference to estimate the alternative scenario of co-existence of GM and non-GM soymeal in Sweden. To assess the model robustness in the alternative scenario for Sweden, we considered the UK price premium for GM soymeal (Table 1). These results, based on simulated prices of GM soy in Sweden, were corroborated by the observed co-existence of GM and conventional non-GM soy markets in Austria. For Austria, we estimated the price transmission model with and without a supply of GM soymeal. This scenario imposed a constraint on imports of GM soy and did not include it in the econometric model. We expected that the reference scenario for Austria without GM soy demonstrates the robustness of the empirical model in determining whether absence of an alternative to non-GM soy would cause asymmetric price adjustment and affect market structure.

Results

In this section, we first present the ARDL model results for the reference scenario in Sweden and report the price transmission elasticities for the alternative scenario with co-existence of GM soymeal. We present the price transmission in the soymeal supply chain in Sweden and Austria in two stages.

Stage 1: Price transmission of soymeal

We measured the transmission of soymeal prices from Brazil to the EU feed industry, particularly Austria and Sweden using the model of Eq. (5). In the reference scenario for Sweden without GM soy, the adjustment parameter was estimated to be -0.289 (Table 2), indicating that the Swedish feed market can adjust 28.9% of the short-term disequilibrium in the price of soymeal. In the alternative scenario with GM soy

Table 2 Estimation of the short-run relationship for price transmission of soymeal

Dep. Var. (Δy_t)	$\Delta lnSoyNonGM_{Swd,t}$		$\Delta lnSoyNonGM_{AT,t}$	
	Sweden.Ref	***Sweden.Alt***	***Austria.Ref***	***Austria.Alt***
Adjustmentcoeft.	−0.289*** (0.066)	−0.283*** (0.082)	−0.357*** (0.075)	−0.400*** (0.076)
Δy_{t-1}	0.025 (0.052)	−0.027 (0.086)	–	–
Δy_{t-2}	0.144*** (0.050)	0.211** (0.061)	–	–
$\Delta lnSoyNonGM^{+}_{Brz,t-1}$	0.077* (0.088)	0.093 (0.091)	0.090 (0.086)	0.038 (0.089)
$\Delta lnSoyNonGM^{+}_{Brz,t-2}$	–	−0.004 (0.083)	–	–
$\Delta lnSoyNonGM^{+}_{Brz,t-3}$	–	−0.140* (0.074)	–	–
$\Delta lnSoyNonGM^{-}_{Brz,t-1}$	0.332*** (0.081)	0.327*** (0.081)	0.048 (0.112)	0.018 (0.112)
$\Delta lnSoyNonGM^{-}_{Brz,t-2}$	–	–	−0.196* (0.103)	−0.150* (0.090)
$\Delta lnSoyNonGM_{Arg,t-1}$	0.362*** (0.071)	0.361*** (0.071)	0.389*** (0.093)	0.342*** (0.093)
$\Delta lnSoyNonGM_{Arg,t-2}$	–	−0.063 (0.075)	0.178** (0.069)	–
$\Delta lnSoyNonGM_{Arg,t-3}$	–	–	0.140*** (0.050)	–
ΔSD_1	0.012 (0.008)	0.011 (0.007)	−0.012 (0.011)	−0.011 (0.011)
ΔSD_2	0.024*** (0.008)	0.027*** (0.008)	−0.010 (0.010)	−0.011 (0.010)
ΔDU	0.128*** (0.029)	0.127*** (0.029)	0.026 (0.016)	0.014 (0.015)
ΔDB	−0.009 (0.011)	−0.004 (0.010)	−0.002* (0.001)	−0.001* (0.001)
$\Delta lnSoyGM_{t-1}$	–	0.005 (0.053)	–	0.192* (0.116)

Note: This table presents the estimated outcomes of the model in Eq. (5). The suffix *'Ref'* and *'Alt'* in the model name indicates *'Reference'* and *'Alternative'*, respectively. The model *'Sweden.Alt'* uses the *Austrian* price premium to generate prices for GM soymeal. Values in brackets are standard error. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Source: authors' calculations

available, the adjustment coefficient was estimated to be 0.283 (see *Sweden.Alt* in Table 2). This was close to that in the reference scenario, at the estimated parameter for GM soy was statistically non-significant. The adjustment coefficient increased to 0.356 when the estimated parameter for GM soymeal was statistically significant (see in *Sweden.SUK* in Appendix A2). This implies that, with GM soymeal permitted, the Swedish feed market would become more responsive to changes in the price of soymeal. This is also confirmed with Austria, as this adjustment in price changes was increased from 35.7% to 40% in the alternative scenario (see *Austria.Ref* and *Austria.Alt* in Table 2).

In BAU conditions (only non-GM soy), the Swedish feed industry showed asymmetric responses to short- and long-term changes in soymeal prices in Brazil. In the short-run, the parameter estimate was 0.077 for positive changes (see $\Delta lnSoyNonGM^{+}_{Brz}$ in Table 2), implying an increase of 0.77% in soymeal prices with a 10% increase in Brazilian prices for non-GM soymeal. However, for a 10% decrease in

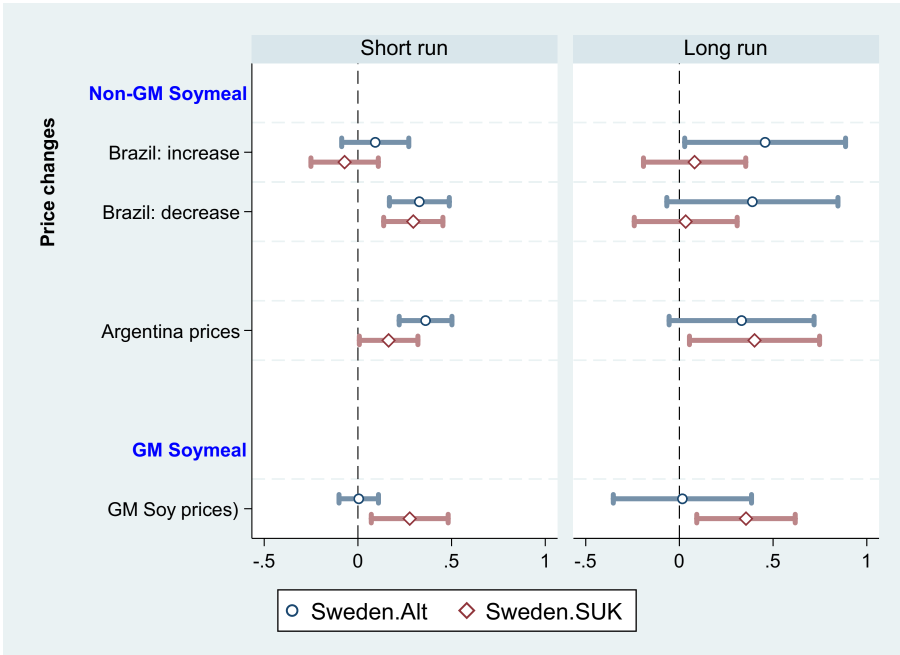

Fig. 2 Estimated model for price transmission of soymeal in Sweden. Note: The model *'Sweden.SUK'* uses the UK price premium to generate prices for GM soymeal in Sweden (see Appendix A3 for the model estimation result)

Brazilian prices, the price of soymeal in Sweden decreased by 3.32% in the reference scenario (see $\Delta lnSoyNonGM^{-}_{Brz}$ in Table 2). This shows that the Swedish feed industry is more responsive to a decrease in the Brazilian price for non-GM soymeal. Interestingly, this effect of the Brazilian price would decrease to 2.96%, even if an alternative (GM soymeal) were available in Sweden (see *Sweden.SUK* in Fig. 2 and Appendix A2).

The presence of GM soymeal could decrease prices of non-GM soymeal in the Swedish feed market due to substitution effects. Even with the influence of GM soy is statistically non-significant, we still observed a decrease in the demand of non-GM soy in Sweden, when its prices increased in Brazil (-0.140 and 0.327 in *Sweden.Alt*, see Table 2). The prices of non-GM soy would further decrease in Sweden, if the influence of GM soymeal was statistically significant (see in *Sweden.SUK* in Appendix A2). This suggests that the supply of GM soymeal may decrease the impact of rising non-GM soy prices in importing countries.

In the long-run, the model parameter of Eq. (5) for positive and negative growth in Brazilian prices of non-GM soymeal was 0.457 and 0.389 in the alternative scenario (see *Sweden.Alt* in Table 3). This is consistent with the model outcome in the reference scenario (0.405 and 0.318 in *Sweden.Ref*). As in the short-term, when the parameter for GM soymeal was statistically significant, these price effects became weak, i.e., statistically non-significant (see *Sweden.SUK* in Fig. 2 and Appendix A2). A similar trend was observed for Austria, with symmetric effects of Brazilian prices of soymeal when the influence of GM soymeal was significant. As in Sweden, if imports of GM soymeal to Austria were restricted, we might also have observed asymmetric effects on the Austrian feed market (see *Austria.Ref* in Table 3). This suggests that the availability of GM

Table 3 Estimated long-run coefficients for price transmission of soymeal

Dep. Var. (Δy_t)	$\Delta lnSoyNonGM_{Swd,t}$		$\Delta lnSoyNonGM_{AT,t}$	
	Sweden.Ref	***Sweden.Alt***	***Austria.Ref***	***Austria.Alt***
$lnSoyNonGM^{+}_{Brz,t-1}$	0.405** (0.181)	0.457** (0.219)	0.252* (0.150)	0.095 (0.224)
$lnSoyNonGM^{-}_{Brz,t-1}$	0.318* (0.191)	0.389* (0.232)	0.094 (0.240)	− 0.033 (0.209)
$lnSoyNonGM_{Arg,t-1}$	0.375** (0.177)	0.331* (0.197)	0.536* (0.270)	0.643*** (0.229)
SD_1	0.044 (0.029)	0.037 (0.030)	− 0.035 (0.031)	− 0.026 (0.027)
SD_2	0.085** (0.035)	0.094** ((0.043)	− 0.028 (0.029)	− 0.027 (0.026)
DU	0.443*** (0.151)	0.448** ((0.179)	0.073 (0.047)	0.036 (0.038)
DB	− 0.031 (0.036)	− 0.014 (0.041)	− 0.004* (0.002)	− 0.004* (0.002)
$lnSoy(GM)_{t-1}$	–	0.016 (0.188)	–	− 0.242 (0.163)
Constant	0.945*** (0.339)	0.983** (0.379)	0.776* (0.442)	0.590 (0.452)
R-square	0.793	0.802	0.7277	0.7278
Adj. *R*-square	0.774	0.772	0.6913	0.6941
DW stat	2.04	1.93	2.08	2.06
Bound test#	7.31***	6.04***	4.43***	5.13***
Ramsey test	1.23	2.16*	1.69	1.73
Wald test (Short-run)	5.73**	4.71**	0.12	0.30
Wald test (Long-run)	4.42**	2.22	3.08*	2.16
CUSUM (CUSUMQ)	S (S)	S (S)	S (S)	S (N)

Note: This table presents the estimated outcomes of the model in Eq. (5). The suffix *'Ref'* and *'Alt'* in the model name indicates *'Reference'* and *'Alternative'*, respectively. The model *'Sweden.Alt'* uses the *Austrian* price premium to generate prices for GM soymeal. Values in brackets are standard error. 'S' stands for stationarity. ***$p<0.01$, **$p<0.05$, *$p<0.1$. #*p*-values are approximated based on critical values in Kripfganz & Schneider (2019). Source: authors' calculations

soymeal can potentially decrease the level of market concentration observed in non-GM soy, thereby facilitating a faster adjustment of asymmetric effects on soymeal prices.

We also observed statistically significant effects of Argentinian soymeal prices on prices in Sweden and Austria (see Tables 2 and 3). Given the magnitude of the effects, it can be assumed that soymeal from Argentina has a greater influence on the Austrian feed market. However, Swedish imports are rather concentrated on the Brazilian supply chain.[1] Argentina supplies around 40% of soymeal to the EU (Berkhout et al. 2018), although the proportion has declined recently due to the rapid adoption of GM varieties in Argentina and corresponding decline in non-GM supply. The cheaper price of Brazilian soymeal (Table 1) could also reduce imports of Argentinian soymeal into the EU.

Based on R^2-values, the estimated model explained 79.3–80.2% of variation in the dependent variable in Sweden, and 72.8% in Austria. Most of these models were free of serial correlation.[2] Similarly, the bound tests revealed the presence of long-term

[1] In 2006, Austria and Sweden recorded 15% and 2.5% imports of soybeans from Argentina, respectively. Since 2012, Sweden has no record of direct import of soybeans from Argentina (Chatham House 2020).

[2] The DW statistic close to 2 indicates no serial correlation (Pindyck and Rubinfeld 1998).

co-integration relationships among the model variables. In most cases, the Ramsey test revealed no missing variable bias in the model specification. For *Sweden.Alt*, the test statistic showed poor fit in the empirical model but, of several possible model specifications, the selected model was best suited to the empirical data. In the reference scenario for Sweden, the F-statistics in the short- and long-term Wald tests were statistically significant, indicating asymmetric effects of positive and negative changes in the Brazilian price of soymeal. However, these effects were symmetric in the alternative scenario with a supply of GM soymeal. Finally, the CUSUM tests showed stability of the estimated parameters.

Stage 2: Price transmission from soymeal to compound feeds

Table 4 shows the regression estimation of Eq. (6) for the second stage in transmission of soymeal prices to compound feeds in Sweden and Austria. In the reference scenario, the adjustment coefficient for poultry feeds in Sweden and Austria was estimated to be 0.252 and 0.195, respectively, while in the alternative scenario it increased to 0.349 in Sweden and 0.278 in Austria (see *Sweden.Alt* and *Austria.Alt* in Table 4). This implies that the price of poultry feed would become more responsive to the market if an alternative to non-GM soymeal were available.

The reference scenario for Sweden showed that the price of poultry feed could increase by 1.89% and 4.41% in the short- and long-run, respectively, with a 10% increase in the price of non-GM soymeal (see *Sweden.Ref* in Tables 4 and 5). Interestingly, the effect was estimated to be 1.90% and 1.34% in the short- and long-run (see *Sweden.Alt* in Tables 4 and 5) when the effects of GM soymeal were statistically significant. This shows that the presence of GM soymeal can replace the use of non-GM soymeal and reduce spill-overs of soy price volatility in poultry feed. Similar results were obtained for negative changes in the price of non-GM soymeal, with a 2.32% and 4.08% decrease in poultry feed prices in the short- and long-run, respectively, with a 10% decrease in non-GM soymeal prices (see *Sweden.Ref* in Tables 4 and 5). These values would be reduced to 1.64% and 1.28% if an alternative (GM soymeal) were available. When the influence of GM soymeal was weak, i.e., statistically non-significant, the price transmission elasticities converged to the parameter estimates in the reference model (see *Sweden.SUK* in Fig. 3). This shows that the presence of GM soy can make the feed market more responsive to market prices for soymeal and reduce asymmetric price adjustments in poultry feed.

With regard to poultry feed in Austria, we observed a greater impact of negative changes in soymeal prices when GM soymeal had a statistically significant influence (see *Austria.Alt* in Table 5). As in Sweden, the effect of positive changes in soymeal prices also decreased in Austria with a supply of GM soy (0.140 and 0.128 in *Austria.Ref* and 0.187 in *Austria.Alt* in Table 4). In the long-run, the parameter estimates for both positive and negative changes in soymeal prices were statistically non-significant in the alternative scenario model, e.g., *Austria.Alt* in Table 5. It is possible that the supply of GM soymeal might have reduced the influence of non-GM soymeal on poultry feed prices, indicating potential for lower feed prices in the EU livestock sector. Statistically significant seasonal dummies and structural breaks indicated the presence of drifts in soymeal and poultry feed prices.

Table 4 Estimation of short-run relationship for price transmission of soymeal to poultry feed

Dep. Var.(Δy_t)	$\Delta lnFeed(poultry)_t$			
	Sweden.Ref	***Sweden.Alt***	***Austria.Ref***	***Austria.Alt***
Adjustmentcoeft.	− 0.252*** (0.062)	− 0.349*** (0.062)	− 0.195*** (0.062)	− 0.278*** (0.076)
Δy_{t-1}	0.308*** (0.083)	0.327*** (0.082)	− 0.277*** (0.086)	− 0.185* (0.107)
$\Delta lnSoyNonGM^{+}_{t-1}$	0.189** (0.072)	0.190*** (0.069)	0.140** (0.069)	0.187** (0.072)
$\Delta lnSoyNonGM^{+}_{t-2}$	− 0.193** (0.075)	–	− 0.063 (0.071)	–
$\Delta lnSoyNonGM^{+}_{t-3}$	–	–	0.128* (0.068)	–
$\Delta lnSoyNonGM^{-}_{t-1}$	0.232*** (0.087)	0.164* (0.091)	− 0.065 (0.084)	− 0.098 (0.088)
$\Delta lnSoyNonGM^{-}_{t-2}$	− 0.111 (0.084)	− 0.013 (0.015)	0.199*** (0.073)	0.256*** (0.089)
$\Delta lnSoyNonGM^{-}_{t-3}$	–	0.228** (0.093)	− 0.181** (0.075)	− 0.105 (0.083)
$\Delta lnWheat_{t-1}$	− 0.108** (0.046)	− 0.075 (0.046)	0.005 (0.021)	− 0.005 (0.022)
$\Delta lnWheat_{t-2}$	0.128*** (0.045)	0.101** (0.045)	–	− 0.021 (0.032)
$\Delta lnWheat_{t-3}$	0.082** (0.032)	0.062* (0.032)	–	− 0.043 (0.028)
$\Delta lnBarl_{t-1}$	0.139** (0.053)	0.118** (0.051)	0.032 (0.029)	0.066** (0.033)
$\Delta lnBarl_{t-2}$	− 0.162*** (0.052)	− 0.159*** (0.052)	− 0.031 (0.029)	− 0.020 (0.040)
$\Delta lnBarl_{t-3}$	–	–	0.073** (0.028)	0.068** (0.033)
$\Delta lnMaize_{t-1}$	–	–	0.016 (0.039)	0.002 (0.041)
$\Delta lnMaize_{t-2}$	–	–	–	0.094** (0.039)
$\Delta lnSoyGM_{t-1}$	–	− 0.109 (0.071)	–	0.040 (0.085)
$\Delta lnSoyGM_{t-2}$	–	0.111** (0.064)	–	–
$\Delta lnSoyGM_{t-3}$	–	0.118** (0.055)	–	–
ΔSD_1	− 0.011 (0.006)	− 0.013** (0.006)	0.003 (0.007)	− 0.015* (0.009)
ΔSD_2	− 0.013** (0.006)	− 0.009 (0.007)	− 0.001 (0.007)	− 0.011 (0.008)
ΔDU_B	0.020 (0.013)	0.024* (0.014)	− 0.019* (0.011)	− 0.017* (0.012)
ΔDT_B	− 0.001 (0.001)	− 0.001 (0.001)	0.000 (0.001)	− 0.000 (0.001)

Note: This table presents the estimated outcomes of the model in Eq. (6). The suffix *'Ref'* and *'Alt'* in the model name indicates *'Reference'* and *'Alternative'*, respectively. The model *'Sweden.Alt'* uses the *Austrian* price premium to generate prices for GM soymeal. Values in brackets are standard error. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Source: authors' calculations

The estimated R^2-values showed good fit of the empirical models. The bound test rejected the null hypothesis of no long-run co-integrating relationship among the model variables. With the exception of *Sweden.SUK*, the DW test did not show

Table 5 Estimated long-run coefficients for price transmission of soymeal to poultry feed

Dep. Var.(Δy_t)	$\Delta lnFeed(poultry)_t$			
	Sweden.Ref	***Sweden.Alt***	***Austria.Ref***	***Austria.Alt***
$lnSoy(NonGM)^+_{t-1}$	0.441*** (0.098)	0.134* (0.074)	0.193* (0.109)	− 0.126 (0.156)
$lnSoy(NonGM)^-_{t-1}$	0.408** (0.171)	0.128* (0.071)	0.117 (0.154)	− 0.248 (0.200)
$lnWheat_{t-1}$	− 0.295 (0.221)	− 0.017 (0.145)	0.025 (0.108)	0.132 (0.131)
$lnBarl_{t-1}$	0.355*** (0.115)	0.392*** (0.137)	0.052 (0.143)	0.106 (0.132)
$lnMaize_{t-1}$	–	–	0.354** (0.186)	0.068 (0.116)
$lnSoyGM_{t-1}$	–	0.276** (0.135)	–	0.328** (0.166)
SD_1	− 0.041 (0.028)	− 0.037* (0.019)	0.016 (0.038)	− 0.055 (0.035)
SD_2	− 0.050* (0.028)	− 0.026* (0.015)	− 0.006 (0.036)	− 0.040* (0.022)
DU	0.078 (0.059)	0.067 (0.044)	− 0.099* (0.059)	− 0.061* (0.036)
DB	− 0.001 (0.001)	− 0.001 (0.001)	0.002 (0.002)	− 0.000 (0.002)
Constant	0.970*** (0.219)	0.823*** (0.288)	0.635** (0.313)	0.582* (0.345)
R-square	0.68	0.73	0.5265	0.5931
Adj. *R*-square	0.61	0.66	0.4251	0.4544
DW stat	2.03	2.11	2.02	2.04
Bound test[#]	6.44***	6.34***	3.29**	2.09***
Ramsey test	1.53	2.05	2.06	1.19
'Wald test (Short-run)	0.10	0.40	2.77*	4.95**
Wald test (Long-run)	0.09	0.10	0.40	1.58
CUSUM (CUSUMQ)	S (S)	S (S)	S (S)	S (S)

Note: This table presents the estimated outcomes of the model in Eq. (6). The suffix *'Ref'* and *'Alt'* in the model name indicates *'Reference'* and *'Alternative'*, respectively. The model *'Sweden.Alt'* uses the *Austrian* price premium to generate prices for GM soymeal. Values in brackets are standard error. ***$p<0.01$, **$p<0.05$, *$p<0.1$. [#]*p*-values approximated based on critical values in Kripfganz & Schneider (2019). Source: authors' calculations

presence of serial correlation. Similarly, the Ramsey test accepted the null hypothesis of no model misspecification errors in empirical models, which implies a well-specified model. For the model *'Sweden.SUK'*, we made a deliberate interpretation of the estimated parameters, as this model was used only for a robustness check of the alternative scenario model in Sweden.

In the case of pig feed, for the model in Eq. (6) the adjustment parameter for Sweden was estimated to be 0.226–0.233, while it was 0.316–0.370 for Austria (see Table 6). This indicates that pig feed prices are responsive to short-term changes in the price of their ingredients in Sweden and Austria. Interestingly, we observed statistically significant effects of negative changes in soymeal prices in Austria and Sweden (see $lnSoyNonGM^-$ in Tables 6 and 7). For Sweden, the magnitude of the effect did not change significantly between the reference and alternative scenarios (0.208–0.220 in the short-term and 0.542–0.560 in the long-term). As the parameter estimates for GM soymeal were low (0.107 for the short-term in Austria (Table 6) and 0.071 for

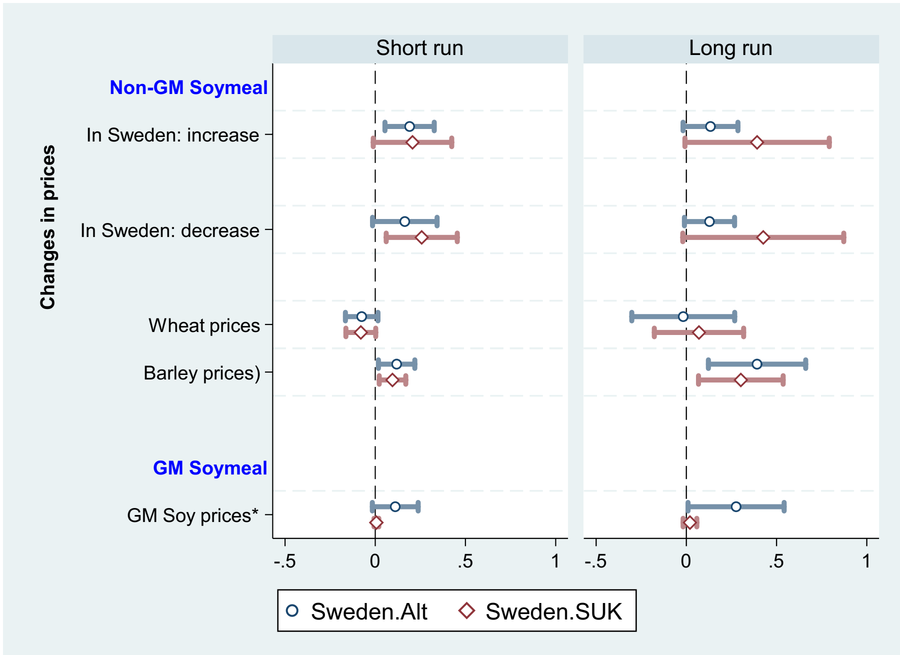

Fig. 3 Estimated model for price transmission of soymeal to poultry feed in Sweden. Note: The model *'Sweden.SUK'* uses the UK price premium to generate prices for GM soymeal in Sweden (see Appendix A3 for the model estimation result)

the long-term in Sweden (Table 7)), we assumed that pig feed prices are likely less responsive to the supply of GM soymeal.

In the long-run, the price of non-GM soymeal had asymmetric effects on the price of pig feed (Fig. 4). In Sweden, pig feed prices increased by 2.14–2.51% with a 10% increase in the price of non-GM soymeal (see $\ln SoyNonGM^{+}$ in Table 7). However, for a 10% decrease in the price of non-GM soymeal, feed prices for pigs fell by 5.42 to 5.60%. This shows that, in the long-run, pig feed prices are more elastic, as they largely react to decreases in soymeal prices. This could have resulted from the availability of GM soymeal in the alternative scenario.

The results for Austria confirmed the presence of asymmetric effects of soymeal prices, similar to Sweden. In the long-term, the effects of negative changes in soymeal prices were more pronounced in Austria (0.315 in *Austria.Ref* and 0.378 in *Austria.Alt*; see Table 7). This shows that the Austrian feed market is likely to be more responsive to changes in market prices for feed ingredients such as soymeal, wheat and barley. Unexpectedly, the short-term effect of GM soymeal was estimated to be -0.107 and statistically significant ($p < 0.1$). This meant that pig feed prices decreased by 1.07% with a 10% increase in the price of GM soymeal (see *Austria.Alt* in Table 6), perhaps because higher prices of GM soymeal encouraged feed manufacturers to use non-GM soymeal for pig feeds. The higher degree of substitutability of GM and non-GM soymeal could have resulted in a negative value of short-term price elasticity. However, in the long-run we found a positive sign for the price transmission elasticities of GM soymeal.

Table 6 Estimation of short-run relationship for price transmission of soymeal to pig feed

Dep. Var. (Δy_t)	$\Delta lnFeed(pig)_t$			
	Sweden.Ref	***Sweden.Alt***	***Austria.Ref***	***Austria.Alt***
Adjustmentcoeft	− 0.226*** (0.040)	− 0.233*** (0.047)	− 0.316*** (0.044)	− 0.370** (0.041)*
Δy_{t-1}	0.219*** (0.083)	0.225** (0.086)	0.099 (0.068)	0.141** (0.071)
$\Delta lnSoyNonGM^{+}_{t-1}$	0.056 (0.087)	0.082 (0.089)	0.023 (0.020)	0.046 (0.030)
$\Delta lnSoyNonGM^{-}_{t-1}$	0.208** (0.105)	0.220** (0.107)	0.100*** (0.030)	0.140** (0.041)*
$\Delta lnWheat_{t-1}$	− 0.063 (0.045)	− 0.053 (0.044)	0.022 (0.023)	0.029 (0.023)
$\Delta lnBarl_{t-1}$	0.149*** (0.042)	0.146*** (0.042)	0.022 (0.030)	0.037 (0.032)
$\Delta lnBarl_{t-2}$	–	–	0.031 (0.031)	–
$\Delta lnMaiz_{t-1}$	–	–	0.099** (0.045)	0.067*** (0.023)
ΔSD_1	− 0.015* (0.008)	-0.015* (0.008)	− 0.014* (0.008)	− 0.008 (0.007)
ΔSD_2	− 0.005 (0.008)	− 0.007 (0.008)	0.028*** (0.009)	0.029*** (0.007)
ΔDU_B	0.038** (0.015)	0.045*** (0.016)	0.017 (0.011)	0.016 (0.012)
ΔDT_B	0.001** (0.000)	0.002** (0.001)	0.002*** (0.000)	0.002*** (0.001)
$\Delta lnSoyGM_{t-1}$	–	− 0.112 (0.086	–	− 0.107* (0.078)

Note: This table presents the estimated outcomes of the model in Eq. (6). The suffix *'Ref'* and *'Alt'* in the model name indicates *'Reference'* and *'Alternative'*, respectively. The model *'Sweden.Alt' uses* the *Austrian* price premium to generate prices for GM soymeal. Values in brackets are standard error. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Source: authors' calculations

The estimated R^2-values showed good fit of the model. The bound test statistic was statistically significant, indicating presence of a long-term co-integration relationship between the model variables. The Wald tests showed long-term asymmetric effects of positive and negative changes in non-GM soymeal prices. The DW test statistics indicated no serious problem with serial correlation. With the exception of *Austria.Ref*, the Ramsey test accepted the null hypothesis of no misspecification in the empirical model due to missing variables. As we already had variables that capture structural breaks within the model variables, the CUSUM test showed stability of the model parameters.

Certain limitations of our models and estimations need to be considered. For example, we only analyzed Austria and Sweden, so future analyses should be extended to some other EU member states, subject to availability of data. Moreover, although we performed counterfactual scenario analysis using monthly time-series data, the simulation results may still entail idiosyncratic errors because of unaccounted factors influencing price transmission in the spatiotemporal model. Moreover, it focuses on analyzing historical monthly prices from 1998 to 2012 due to the unavailability of monthly prices for GM soy. Nevertheless, its relevance to the EU and its Member States remains significant. This study could have further benefitted

Table 7 Estimated long-run coefficients for price transmission of soymeal to pig feed

Dep. Var. (Δy_t)	$\Delta lnFeed(pig)_t$			
	Sweden.Ref	***Sweden.Alt***	***Austria.Ref***	***Austria.Alt***
$lnSoyNonGM^{+}_{t-1}$	0.251** (0.098)	0.214* (0.124)	0.072 (0.062)	0.125 (0.083)
$lnSoyNonGM^{-}_{t-1}$	0.542*** (0.174)	0.560* (0.331)	0.315*** (0.089)	0.378*** (0.111)
$lnWheat_{t-1}$	−0.278 (0.211)	−0.227 (0.203)	0.259*** (0.091)	0.237*** (0.084)
$lnBarl_{t-1}$	0.161*** (0.086)	0.267** (0.105)	0.200** (0.099)	0.247*** (0.083)
$lnMaiz_{t-1}$	–	–	0.224*** (0.081)	0.181*** (0.062)
SD_1	−0.068* (0.038)	−0.065* (0.037)	−0.046* (0.025)	−0.021 (0.020)
SD_2	−0.024 (0.036)	−0.028 (0.035)	0.088*** (0.032)	0.078*** (0.021)
DU	0.168** (0.075)	0.191** (0.086)	0.055 (0.036)	0.042 (0.032)
DB	0.004** (0.002)	0.004** (0.002)	0.005*** (0.001)	0.005*** (0.001)
$lnSoyGM_{t-1}$	–	0.071* (0.039)	–	0.014 (0.083)
Constant	0.973*** (0.185)	0.916*** (0.241)	0.818*** (0.162)	0.994*** (0.195)
R-square	0.52	0.52	0.686	0.6844
Adj. *R*-square	0.46	0.45	0.639	0.6384
DW stat	1.96	1.99	2.04	2.02
Bound test#	6.63***	5.91***	8.92***	12.42***
Ramsey test	0.38	0.31	2.20	1.94
Wald test (short-run)	0.90	0.70	10.11***	12.02***
Wald test (long-run)	6.41**	7.16***	10.76***	13.46***
CUSUM (CUSUMQ)	S (S)	S (N)	S (N)	S (N)

Note: This table presents the estimated outcomes of the model in Eq. (6). The suffix *'Ref'* and *'Alt'* in the model name indicates *'Reference'* and *'Alternative'*, respectively. The model *'Sweden.Alt'* uses the *Austrian* price premium to generate prices for GM soymeal. Values in brackets are standard error. ***$p<0.01$, **$p<0.05$, *$p<0.1$. #*p*-values are approximated based on critical values in Kripfganz & Schneider (2019). Source: authors' calculations

from the availability of long-term price data on GM soy for enabling the generation of more robust empirical estimations for the EU market.

Conclusion and implications

Brazil and Argentina are the major suppliers of soybeans and soymeal to the European feed industry. These exporting countries have been rapidly embracing GM varieties of soybean and this has created friction in their trade with EU Member States, many of which have been slow to approve imports of GM varieties into their market. Some Member States, such as Sweden, do not accept any imports of GM soy for animal feed. In recent times, the prices for non-GM soymeal has been increased sharply in the EU, which could affect the competitiveness of European livestock sector (Jordbruksaktuellt 2021). As markets for soy are also quickly emerging in China and India, EU food companies and supermarkets have been lobbying for relaxation of the current GM restrictions (Popp et al. 2013). Against this background, we analyzed the effect of the constraints on

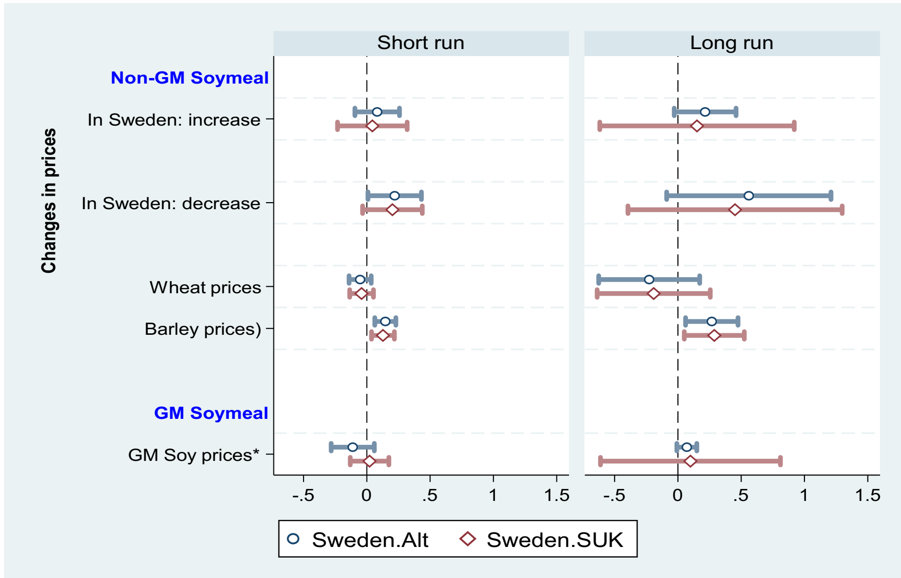

Fig. 4 Estimated model for price transmission of soymeal to pig feed in Sweden

GM imports in Sweden and compared the Swedish situation to that in Austria, which has a less restrictive policy, through estimation of price transmission elasticities of imported soy.

The results of our empirical analysis revealed asymmetric effects of Brazilian soymeal prices on the EU feed market, confirming incomplete transmission of prices in the feed industry. This means that increased costs of soymeal are passed on to consumers, but decreased costs are not passed on to the same extent. This may create an incentive for EU feed manufacturers to exercise perceived market dominance in adjusting price distortions in the soy supply chain. However, the supply of GM soymeal may reduce these asymmetries in the adjustment of soy prices by lessening the level of market concentration in non-GM soy. The substitution effects of GM soymeal in Austria and the simulated substitution effects for Sweden suggested strong asymmetric adjustments in transmission of soymeal prices to compound feeds. The lower price of GM soy could influence the price adjustment process if EU feed prices rise, while for a decrease in feed prices the supply of GM soymeal seems to be the last choice for feed manufacturers. This implies that when non-GM soy prices become competitive in South America, EU farmers may lose their competitiveness, especially in countries where GM soy imports are not accepted, e.g., Sweden.

In Austria, farmers can choose either to compete with the added value of GM-free food or to use the cheaper GM soy. In Sweden, farmers do not have the option to become competitive by using cheaper GM soy. We found symmetric effects of soy prices on the Austrian feed market, because of the alternative supply of Argentinian soymeal. A relaxation of the no-GM rule for imports to Sweden could open up trade relations with Argentina and the USA, where GM soy is adopted across their entire soybean cultivated area. Trade relations have been interrupted since 2012 because of failure of these countries in maintaining complete segregation between non-GM and GM varieties in the supply chain. Relaxing imports of GM soy would likely bring trade gains to the EU and reduce market uncertainties for livestock farmers. The other option for Swedish farmers to be competitive would be to grow low-cost protein feedstuff locally (Eriksson et al. 2018). This would of course require a major policy shift and high investments in biotech crop production, increased efficiency of production of local feedstuffs such as rapeseed cake, and improved breeding and cultivation of local protein crops.

Appendix

Appendix A1: Unit Root test with seasonality and structural break

Variables	Selected model	Breakpoint (T_B)	ADF test-stat ($t\text{-}stat_\alpha$)	Result
$ln\ SoyNonGM_{Swd}$	Perron IO model	2009M01	−3.734	N
$ln\ SoyNonGM_{Arg}$	Perron AO model	2007M04	−4.162	N
$ln\ SoyNonGM_{Brz}$	ZA model, break (intercept)	2007M06	−4.025	N
$SoyNonGM^{+}_{Brz}$	ZA model, break (intercept)	2009M01	−3.562	S
$SoyNonGM^{-}_{Brz}$	Perron IO model	2007M12	−6.001	S
$ln\ Feed(poultry)_{Swd}$	ZA model, break (intercept)	2009M09	−3.994	S
$ln\ Feed(poultry)_{AT}$	Perron AO model	2007M05	−3.066	N
$ln\ Feed(pig)_{Swd}$	Perron IO model	2009M01	−3.867	N
$ln\ Feed(pig)_{AT}$	ZA model, break (intercept)	2007M08	−4.691	S
$lnBarl_{Swd}$	ZA model, break (intercept)	2007M10	−4.007	N
$lnBarl_{AT}$	Perron AO model	2001M07	−2.952	N
$lnWheat_{Swd}$	ZA model, break (intercept)	2006M10	−4.459	N
$lnWheat_{AT}$	ZA model, break (intercept)	2006M07	−4.285	S
$lnMaiz_{AT}$	ZA model, break (intercept)	2006M08	−4.071	S

ZA Zivot and Andrews, IO Innovational outlier, AO Additive outlier, S stationary, N Non-stationary, M Month. $SoyCon^{+}_{Brz}$ and $SoyNonGM^{-}_{Brz}$ are the partial sum of positive and negative changes in $ln\ SoyCon_{Brz}$, respectively. Critical values at 5% level: −4.86 (Perron IO1), −4.44 (Perron AO), −4.80 (ZA). Source: authors' calculations.

Appendix A2: Estimation of price transmission relationship of soymeal in Sweden

***Dep. Var.* (Δy_t):**	***Model: Sweden.SUK***		
$\Delta ln\ SoyNonGM_{Swd,t}$	**Short-run**	**Long-run**	
Adjustment coeft.	−0.356*** (0.077)		–
Δy_{t-1}	0.095 (0.062)		–
Δy_{t-2}	0.271*** (0.058)		–
$\Delta ln\ SoyNonGM^{+}_{Brz,t-1}$	0.071 (0.091)	$lnSoyNonGM^{+}_{Brz,t-1}$	0.081 (0.139
$\Delta ln\ SoyNonGM^{+}_{Brz,t-2}$	−0.041 (0.076)		–
$\Delta ln\ SoyNonGM^{+}_{Brz,t-3}$	−0.214* (0.072)		–
$\Delta ln\ SoyNonGM^{-}_{Brz,t-1}$	0.296*** (0.080)	$lnSoyNonGM^{-}_{Brz,t-1}$	0.034 (0.140)

Dep. Var. **(Δy_t):**	***Model: Sweden.SUK***		
$\Delta ln\ SoyNonGM_{Swd,t}$	**Short-run**	**Long-run**	
$\Delta ln\ SoyNonGM_{Arg,t-1}$	0.163**(0.0799)	$ln\ SoyNonGM_{Arg,t-1}$	0.401** (0.177)
ΔSD_1	0.003 (0.007)	SD_1	0.008 (0.016)
ΔSD_2	0.020** (0.008)	SD_2	0.043** (0.019)
ΔDU	0.129*** (0.026)	DU	0.284*** (0.078)
ΔDB	− 0.004 (0.010)	DB	− 0.008 (0.022)
$\Delta ln\ SoyGM_{t-1}$	0.276** (0.104)	$ln\ Soy(GM)_{t-1}$	0.355** (0.134)
	–	Constant	0.523 (0.396)
R-square	0.842	Ramsey test	1.37
Adj. *R*-square	0.813	Wald test (Short-run)	12.53***
DW stat	2.07	Wald test (Long-run)	3.76
Bound test#	8.34***	CUSUM (CUSUMQ)	S (S)

This table presents the estimated outcomes of the model in Eq. (6). The model 'Sweden.SUK' uses the UK price premium to generate prices for GM soymeal. Parentheses indicate standard error. *** p < 0.01, ** p < 0.05, * p < 0.1. #p-values are approximated based on Kripfganz and Schneider (2019) critical values. Source: authors' calculations.

Appendix A3: Estimation of price transmission relationship of soymeal to poultry feed

Dep. Var. **(Δy_t):**	***Model: Sweden.SUK***		
$\Delta ln\ Feed(poultry)_t$	**Short-run**	**Long-run**	
Adjustment coeft.	− 0.316*** (0.046)	–	–
Δy_{t-1}	0.256*** (0.081)	–	–
$\Delta ln\ SoyNonGM^{+}_{t-1}$	0.206* (0.110)	$ln\ Soy(NonGM)^{+}_{t-1}$	0.392* (0.204)
$\Delta ln\ SoyNonGM^{+}_{t-2}$	− 0.238*** (0.072)	–	–
$\Delta ln\ SoyNonGM^{+}_{t-3}$	–0.129* (0.073)	–	–
$\Delta ln\ SoyNonGM^{-}_{t-1}$	0.257** (0.101)	$ln\ Soy(NonGM)^{-}_{t-1}$	0.426* (0.227)
$\Delta lnWheat_{t-1}$	− 0.079* (0.042)	$lnWheat_{t-1}$	0.069 (0.126)
$\Delta lnBarl_{t-1}$	0.096** (0.037)	$lnBarl_{t-1}$	0.302** (0.119)
$\Delta ln\ SoyGM_{t-1}$	0.006 (0.007)	$ln\ SoyGM_{t-1}$	0.020 (0.019)
ΔSD_1	− 0.009 (0.006)	SD_1	− 0.031 (0.021)
ΔSD_2	− 0.012* (0.007)	SD_2	− 0.038* (0.021)
ΔDU_B	0.025 (0.017)	DU	0.079 (0.057)
ΔDT_B	0.000 (0.001)	DB	0.000 (0.001)

Dep. Var. (Δy_t):	***Model: Sweden.SUK***		
$\Delta ln\ Feed(poultry)_t$	**Short-run**	**Long-run**	
–	–	Constant	1.194*** (0.373)
R-square	0.64	Ramsey test	1.59*
Adj. *R*-square	0.58	Wald test (Short-run)	0.13
DW stat	1.90	Wald test (Long-run)	0.06
Bound test[#]	6.94***	CUSUM (CUSUMQ)	S (S)

This table presents the estimated outcomes of the model in Eq. (6). The model 'Sweden.SUK' uses the UK price premium to generate prices for GM soymeal. Parentheses indicate standard error. *** $p<0.01$, ** $p<0.05$, * $p<0.1$. [#]p-values are approximated based on Kripfganz and Schneider (2019) critical values. Source: authors' calculations.

Appendix A4: Estimation of short-run relationship for price transmission of soymeal to pig feed

Dep. Var. (Δy_t):	**Model: Sweden.SUK**		
$\Delta ln\ Feed(pig)_t$	**Short-run**	**Long-run**	
Adjustment coeft	− 0.217*** (0.041)	–	–
Δy_{t-1}	0.170* (0.086)	–	–
$\Delta ln\ SoyNonGM^{+}_{t-1}$	0.044 (0.140)	$ln\ SoyNonGM^{+}_{t-1}$	0.214* (0.124)
$\Delta ln\ SoyNonGM^{-}_{t-1}$	0.203* (0.120)	$ln\ SoyNonGM^{-}_{t-1}$	0.560* (0.331)
$\Delta lnWheat_{t-1}$	− 0.042 (0.047)	$lnWheat_{t-1}$	− 0.227 (0.203)
$\Delta lnBarl_{t-1}$	0.128*** (0.046)	$lnBarl_{t-1}$	0.267** (0.105)
ΔSD_1	− 0.016* (0.008)	SD_1	− 0.065* (0.037)
ΔSD_2	− 0.005 (0.009)	SD_2	− 0.028 (0.035)
ΔDU_B	0.039* (0.021)	DU	0.191** (0.086)
ΔDT_B	0.001 (0.001)	DB	0.004** (0.002)
$\Delta ln\ SoyGM_{t-1}$	0.021 (0.077	$ln\ SoyGM_{t-1}$	0.071* (0.039)
Constant	–	–	0.806* (0.465)
R-square	0.50	Ramsey test	0.80
Adj. *R*-square	0.44	Wald test (short-run)	0.81
DW stat	1.95	Wald test (long-run)	1.64
Bound test[#]	4.90***	CUSUM (CUSUMQ)	S (S)

This table presents the estimated outcomes of the model in Eq. (6). The model 'Sweden.SUK' uses the UK price premium to generate prices for GM soymeal. Parentheses indicate standard error. *** $p<0.01$, ** $p<0.05$, * $p<0.1$. [#]p-values are approximated based on Kripfganz and Schneider (2019) critical values. Source: authors' calculations.

Appendix A5: The ARDL model in error correction form

Consider a model represented as follows:

$$Y_t = m + \alpha Y_{t-1} + \beta_0 X_t + \beta_1 X_{t-1} + e_t \tag{7}$$

We can express Eq. (7) in an alternative form:

$$Y_t = \frac{1}{1-\alpha L}[m + \beta_0 X_t + \beta_1 X_{t-1} + e_t] \tag{8}$$

Here, the symbol L is a lag operator, where $LY_t = Y_{t-1}$.

Analyzing short and Long-term effect of *X* on *Y*:

In the short-run, the effect of X on Y can be derived as follows: $\frac{\partial Y}{\partial X} = \beta_0$.

However, in the long-run, we assume that a one-unit change in X results in a new level, denoted as $\overline{X}$, and drives Y towards a "long-run equilibrium". We denote the long-run level of Y corresponding to $\overline{X}$ as $\overline{Y}$.

Solving for equilibrium $\overline{Y}$, we set $e_t = \overline{e} = 0$, indicating the absence of shocks in equilibrium, where $Y_t = Y_{t-1} = \overline{Y}$. This implies that L equals 1. Consequently, Eq. (7) becomes:

$$\overline{Y} = \frac{1}{(1-\alpha)}\left[m + (\beta_0 + \beta_1)\overline{X}\right] \tag{9}$$

We can observe the long-run impact of X on Y as: $\frac{\partial \overline{Y}}{\partial \overline{X}} = \frac{\beta_0+\beta_1}{1-\alpha}$.

Expressing the ARDL model in "Error Correction" form:

By defining $Y_t = Y_{t-1} + \Delta Y_t$ and $X_t = X_{t-1} + \Delta X_t$, we can restructure Equation (7) as:

$$\Delta Y_t = \beta_0 \Delta X_t - (1-\alpha)\left[Y_{t-1} - \frac{m}{1-\alpha} - \frac{\beta_0 + \beta_1}{1-\alpha} X_{t-1}\right] + e_t$$

Alternatively, we can represent it as:

$$\Delta Y_t = \beta_0 \Delta X_t - (1-\alpha)[Error]_{t-1} + e_t \tag{10}$$

Here, $Error_{t-1}$ is defined as:

$$Error_{t-1} = Y_{t-1} - \frac{m}{1-\alpha} - \frac{\beta_0 + \beta_1}{1-\alpha} X_{t-1}$$

The term $(1-\alpha)$ serves as the adjustment coefficient. In the long-run, the term within the square brackets becomes zero. When Y_{t-1} deviates from equilibrium at time $t-1$ with $\left[Y_{t-1} > \frac{m}{1-\alpha} - \frac{\beta_0+\beta_1}{1-\alpha} X_{t-1}\right]$, this term will bring Y_{t-1} back towards equilibrium, as the adjustment coefficient is negative.

Conversely, if $Y_{t-1} < \frac{m}{1-\alpha} - \frac{\beta_0+\beta_1}{1-\alpha} X_{t-1}$, the contents of square bracket are negative. In this case, the negative adjustment coefficient will push Y_t upwards toward equilibrium in the subsequent period. This mechanism explains the term inside the square bracket as "error correction term" and the coefficient as the adjustment parameter.

Acknowledgements

Not applicable

Author contributions
SKB designed the concept, collected and analyzed the data, and prepared the first draft of the manuscript. RG and ME helped in conceptualization, aided in interpreting the results and preparing the final draft of the manuscript. CJL commented on previous versions of the manuscript. All authors verified the analytical methods, contributed to the interpretation of the results and worked on the manuscript. All authors read and approved the final manuscript.

Funding
Open access funding provided by Swedish University of Agricultural Sciences. This research was financially supported by the Swedish Foundation for Strategic Environmental Research through the research program Mistra Biotech 2.

Availability of data and materials
The datasets used and/or analyzed during the current study are available from the corresponding author on reasonable request.

Declarations

Competing interests
The authors declare that they have no competing interests.

Received: 28 December 2022 Revised: 21 September 2023 Accepted: 6 November 2023
Published online: 23 November 2023

References
Agrimatie (2020) Agrarische prijzen. LEI Wageningen UR. In: agrimatie.nl/Prijzen.aspx
Assefa TT, Kuiper WE, Meuwissen MPM (2014) The effect of farmer market power on the degree of farm retail price transmission: a simulation model with an application to the Dutch ware potato supply chain. Agribusiness 34(4):424–437. https://doi.org/10.1002/agr.21371
Barboza MG, Behrendt K, Paparas D (2022) Price transmission analysis of the international soybean market in a trade war context. Economies 10(8):1–22. https://doi.org/10.3390/economies10080203
Barboza MG, Paparas D, Behrendt K (2023) Asymmetric price transmission analysis of the international soybean market. Agric Sci 14:317–334. https://doi.org/10.4236/as.2023.143020
Berkhout P, Achterbosch T, van Berkum S, Dagevos H, Dengerink J, van Duijn AP, Terluin IJ (2018) Global implications of the European Food System: A food systems approach. Wageningen
Cederberg C, Sonesson U, Henriksson M, Sund V, Davis J (2009) Greenhouse gas emissions from Swedish production of meat, milk and eggs 1990 and 2005. The Swedish Institute for Food and Biotechnology, SIK Report No 793, Gothenburg
Chatham House (2020) resourcetrade.earth. https://resourcetrade.earth/
de Boer HC, van Krimpen MM, Blonk H, Tyszler M (2014) Replacement of soybean meal in compound feed by European protein sources. Livestock Research Report 819
Desquilbet M, Bullock DS (2009) Who pays the costs of non-GMO segregation and identity preservation? Am J Agric Econ 91:656–672. https://doi.org/10.1111/j.1467-8276.2009.01262.x
Eriksson M, Ghosh R, Hansson E, Basnet S, Lagerkvist CJ (2018) Environmental consequences of introducing genetically modified soy feed in Sweden. J Clean Prod 176:46–53. https://doi.org/10.1016/j.jclepro.2017.12.113
EU (2019) European Union (EU): overview of EU feed supply. In: https://ec.europa.eu. https://www.pig333.com/latest_swine_news/eu-overview-of-eu-feed-supply_14987/. Accessed 14 Apr 2020
Commission E (2003) Regulation (EC) No 1829/2003 of the European Parliament and of the Council of 22 September 2003 on genetically modified food and feed. Off J Eur Union L 268:1–23
European Commission (2007) Economic Impact of Unapproved GMOs on EU Feed Imports and Livestock Production. Directorate-General for Agriculture and Rural Development, Brussels. https://agriculture.ec.europa.eu/system/files/2020-04/economic_impactgmos_en_0.pdf
Frey G, Manera M (2007) Econometric models of asymmetric price transmission. J Econ Surv 21:349–415. https://doi.org/10.1111/j.1467-6419.2007.00507.x
Gizaw D, Myrland Ø, Xie J (2021) Asymmetric price transmission in a changing food supply chain. Aquac Econ Manag 25(1):89–105. https://doi.org/10.1080/13657305.2020.1810172
Goodwin BK, Holt MT (1999) Price transmission and asymmetric adjustment in the U.S. Beef Sector. Am J Agric Econ 81:630–637. https://doi.org/10.2307/1244026
Grant W (2008) GM feed ban crisis | CAP Reform. In: CAP Reform.EU. http://capreform.eu/gm-feed-ban-crisis/. Accessed 16 Jan 2022
Henseler M, Piot-Lepetit I, Ferrari E, Mellado AG, Banse M, Grethe H, Parisi C, Hélaine S (2013) On the asynchronous approvals of GM crops: potential market impacts of a trade disruption of EU soy imports. Food Policy 41:166–176. https://doi.org/10.1016/j.foodpol.2013.05.005
Jordbruksaktuellt (2021) Rekordhög kostnad för GMO-fritt sojamjöl. In: Nr 13. https://www.ja.se/artikel/2228884/visa?utm_source=nyhetsbrev&utm_medium=email&utm_content=Artikel-2228884&utm_campaign=210709-2520. Accessed 12 Jul 2021
Kalaitzandonakes N (2011) The economic impacts of asynchronous authorizations and low-level presence: an overview. International Food & Agricultural Trade Policy Council (IPC) Position Paper 22
Kalaitzandonakes N, Kaufman J, Miller D (2014) Potential economic impacts of zero thresholds for unapproved GMOs: the EU case. Food Policy 45:146–157. https://doi.org/10.1016/j.foodpol.2013.06.013

Kamaruddin RM, Sofyan S, Shabri AM (2021) Asymmetric price transmission of Indonesian coffee. Cogent Econ Finance 9:1. https://doi.org/10.1080/23322039.2021.1971354
Kamyabi N, Chidmi B (2023) Asymmetric price transmission between crude oil and the US gasoline market. J Risk Financ Manag 16(7):326. https://doi.org/10.3390/jrfm16070326
Karantininis K, Katrakylidis K, Persson M (2011) Price transmission in the Swedish pork chain: asymmetric nonlinear ARDL. In: EAAE congress, change and uncertainty challenges for agriculture, food and natural resources. pp 1–12
Krautgartner R (2017) Austrial Agricultural Biotechnology Annual. GAIN Report Number AU1707
Kripfganz S, Schneider D (2019) Response surface regressions for critical value bounds and approximate p-values in equilibrium correction models1. Oxford Bull Econ Stat 82(6):1456–1481. https://doi.org/10.1111/obes.v82.6, https://doi.org/10.1111/obes.12377
Kroes H, Kuepper B (2015) Mapping the soy supply chain in Europe, Amsterdam, The Netherlands, Profundo. https://d2ouvy59p0dg6k.cloudfront.net/downloads/mapping_soy_supply_chain_europe_wwf_2015.pdf
Kuepper, B., M. Stravens (2022). Mapping the European Soy Supply Chain—Embedded Soy in Animal Products Consumed in the EU27+UK, Amsterdam, The Netherlands: Profundo. https://wwfeu.awsassets.panda.org/downloads/mapping_the_european_soy_supply_chain_e4c.pdf
Lee J, Strazicich MC (2004) Minimum LM Unit Root Test with One Structural Break. No 04–17, Working Papers. Department of Economics, Appalachian State University, Boone
Margarido MA, Turolla FA, Bueno CRF (2007) The world market for soybeans: price transmission into Brazil and effects from the timing of crop and trade. Nova Econ 17(2):241–270. https://doi.org/10.1590/S0103-63512007000200002
Meyer J, Cramon-Taubadel S (2004) Asymmetric price transmission: a survey. J Agric Econ 55:581–611. https://doi.org/10.1111/j.1477-9552.2004.tb00116.x
Nakajima T (2011) Asymmetric price transmission in the U.S. soybean exports. Int J Agric Res 6(4):368–376. https://doi.org/10.3923/ijar.2011.368.376
Peeters L, Surry Y (1997) A review of the arts of estimating price-responsiveness of feed demand in the European Union. J Agric Econ 48:379–392. https://doi.org/10.1111/j.1477-9552.1997.tb01160.x
Pesaran MH, Shin Y (1998) An autoregressive distributed-lag modelling approach to cointegration analysis. In: Strom S (ed) Econometrics and economic theory in the 20th century. Cambridge University Press, Cambridge, pp 371–413
Pesaran MH, Shin Y, Smith RJ (2001) Bounds testing approaches to the analysis of level relationships. J Appl Econom 16:289–326. https://doi.org/10.1002/jae.616
Pindyck RS, Rubinfeld DL (1998) Econometric models and economic forecasts. McGraw-Hill/Irwin, New York
Popp J, Pető K, Magda R, Lakner Z (2013) Economic Impact of GM Hysteria on EU Feed Market. Am J Plant Sci 4:1547–1553. https://doi.org/10.4236/ajps.2013.48186
Serra T, Goodwin BK (2003) Price transmission and asymmetric adjustment in the Spanish dairy sector. Appl Econ 35:1889–1899. https://doi.org/10.1080/00036840310001628774
Shin Y, Yu B, Greenwood-Nimmo M (2014) Modelling asymmetric cointegration and dynamic multipliers in a nonlinear ARDL framework. In: Festschrift in Honor of Peter Schmidt, Springer New York, pp 281–314
Smith PJ, Katovich ES (2017) Are GMO policies "trade related"? Empirical analysis of Latin America. Appl Econ Perspect Policy 39:286–312. https://doi.org/10.1093/aepp/ppw021
Taslim MA, Hossain MA (2015) Asymmetric transmission of international price of edible oil in Bangladesh. Bangladesh Dev Stud 38(1):33–54
Tillie P, Rodríguez-Cerezo E (2015) Markets for non-Genetically Modified, Identity-Preserved soybean in the EU. JRC Science and Policy Report EUR 27203. Publications Office of the European Union. Luxembourg. https://doi.org/10.2791/949110
Weldegebriel HT (2004) Imperfect price transmission: is market power really to blame? J Agric Econ 55:101–114. https://doi.org/10.1111/j.1477-9552.2004.tb00082.x
Zhou D, Koemle D (2015) Price transmission in hog and feed markets of China. J Integr Agric 14:1122–1129. https://doi.org/10.1016/S2095-3119(14)60995-3
Zivot E, Andrews DWK (2002) Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis. J Bus Econ Stat 20:25–44. https://doi.org/10.1198/073500102753410372

Publisher's Note