Price Transmission in Pakistan's Poultry Supply Chain

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Increasing levels of concentration in the upstream sector of poultry supply chains have led to concerns about the influence of producers on farm-gate prices. Against this background, we employ threshold cointegration models to study price transmission mechanisms in Pakistan's poultry sector. We do not find evidence of asymmetric price transmission in the chicken supply chain, but evidence of asymmetric price transmission in the egg supply chain points towards the exercise of market power by egg producers. Differences in the short-run price transmission mechanisms of eggs and chicken can be traced to the underlying production and marketing activities, particularly degree of product storability.

Key words: asymmetric price transmission, chicken and egg prices, error-correction model, threshold cointegration

Introduction

Economists have a keen interest in studying how prices adjust to the forces of supply and demand under different institutional settings because prices coordinate commercial activities in a market economy. Radical changes in the farm, processing and retail sectors of the food industry during the twenty-first century spurred research on price transmission mechanisms between upstream and downstream markets of agricultural supply chains. Meyer and von Cramon-Taubadel (2004), Frey and Manera (2007), and Lloyd (2017) summarize this vast literature.

The goal of vertical price transmission research is to describe the mechanisms of price adjustment between upstream and downstream markets. To this end, different econometric models are employed to estimate the magnitude, direction, and speed of price adjustments along various stages of a supply chain. More precisely, economists want to determine whether there is a long-run relationship between upstream and downstream prices. Of equal importance are questions about the "leading" price in a supply chain (i.e., are prices set in upstream or downstream markets?). Economists are also interested in the direction of price transmission in supply chains and the nature of the equilibrium-adjustment path (i.e., the speed with which upstream and downstream prices adjust to deviations from the long-run equilibrium). In particular, economists are concerned about the extent to which price transmission mechanisms are symmetric (i.e., whether negative and positive deviations from the long-run equilibrium are corrected in an identical manner).

Answers to these questions shed light on the efficacy of market mechanisms in a supply chain. Cointegration of upstream and downstream prices in the long run provides evidence of well-

Review coordinated by Darren Hudson.

functioning markets, whereby economic forces ensure that upstream and downstream markets of a supply chain are interlinked in the long run. The causal role of farm supply or retail demand in the long-run dynamics of prices can be inferred from the "leading" price in a supply chain. While, analysis of the direction and speed of price transmission in upstream and downstream markets sheds light on short-run price dynamics. Above all, evidence of asymmetries—either in the speed of adjustment toward the long-run equilibrium or in the magnitude of short-run adjustments to positive and negative price shocks—is often interpreted as a sign of market failures.

The empirical price transmission literature has provided important insights on the functioning of agricultural markets. Peltzman (2000) shaped the early literature by observing that downstream prices of consumer goods responded faster to increases in upstream prices than to decreases in upstream prices due to the market power of U.S. retailers. Subsequent research on vertical price transmission was largely motivated by increases in the market power of food retailers worldwide (Lloyd, 2017). For example, Abdulai (2002) show that increases in farm-gate pork prices in Switzerland are passed on to retail prices faster than equivalent reductions in farm-gate pork prices. Alam et al. (2016) and Korale Gedara, Ratnasiri, and Bandara (2016) find similar types of asymmetric price transmission from upstream to downstream markets in the rice sectors of Bangladesh and Sri Lanka. However, asymmetric price transmission from upstream to downstream markets is not a rule. Bakucs, Fałkowski, and Fertő (2014) note that asymmetric price transmission is rejected in roughly half of all price transmission studies on agricultural markets. For instance, Serra and Goodwin (2003) and Tekgüç (2013) do not find evidence of asymmetric price transmission in the milk supply chains of Spain and Turkey, respectively, despite suspicions of retailer market power.

Interestingly, and contrary to popular perception, empirical research shows that price transmission does not necessarily "flow" from upstream to downstream prices (i.e., input prices do not always drive output prices) (Meyer and von Cramon-Taubadel, 2004). Bakucs, Fałkowski, and Fertő (2014) cite several studies in which the dynamics of upstream prices are driven by changes in downstream prices. For example, Bakucs and Fertő (2005) find that price transmission in Hungary's pork supply chain is symmetric with short-run and long-run causality running from retail to producer prices. Likewise, Reziti and Panagopoulos (2008) document symmetric transmission from retail to producer prices in Greek fruit markets but asymmetric price transmission in the opposite direction in vegetable markets. Answers to questions about price transmission are also influenced by attributes of the underlying commodities (degree of storability, production dynamics, etc.). Ahn and Lee (2015) find that price transmission patterns are closely related to product perishability, while, Simioni et al. (2013) study the effects of uncertainties in farm supply on price transmission. Finally, Frey and Manera (2007) point out that findings of vertical price transmission literature are sensitive to methodological differences like type of econometric model and frequency of price data.

A review of the literature thus suggests that the nature of price transmission in a given supply chain is an empirical question that depends on the underlying commodity, institutional settings, and research methodology. Against this background, our study of price transmission mechanisms in Pakistan's egg and chicken supply chains extends the existing literature in several directions. First, price transmission research on the poultry supply chain of Pakistan—an agrarian economy plagued by poor marketing infrastructure, fragmented supply chains, and slow growth in the retail sector— promises to offer key insights on the functioning of agricultural markets in developing countries.[1]

Second, the empirical literature is almost exclusively based on studies motivated by retailers' market power. A major focus of these studies is to determine whether estimates from price transmission models are consistent with retailer market power. However, the transformation of poultry production from small-scale subsistence farming to large-scale commercial operations has led to concerns about market power at the upstream level of poultry supply chains.[2] Yet we do not find a single case study in the literature that attempts to study mechanisms of price transmission

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[1] For sake of brevity, the phrase "poultry supply chains" in our paper refers to both chicken and egg supply chains.

[2] Poultry producers in the United States were also hit by a series of antitrust lawsuits between 2016 and 2018 (Dewey, 2018).

under conditions of farmers' alleged market power. In this context, lawsuits filed by the Competition Commission of Pakistan (CCP) against the Pakistan Poultry Association (PPA) for manipulating farm-gate prices provide the ideal setting to address an important gap in the existing literature. Specifically, our paper represents the first attempt in the empirical literature that utilizes econometric models of price transmission to investigate the existence and form of market power abuse by producers in an agricultural supply chain. The empirical analysis in our paper is informed by extensive fieldwork on Pakistan's poultry supply chains.[3] Thus, our paper also serves as an example of the elusive "smoking gun" case study described by Meyer and von Cramon-Taubadel (2004), whereby we examine antitrust litigation in the light of estimates from price transmission models and knowledge of local institutional settings.

Third, from a methodological standpoint, we utilize advances in nonlinear cointegration analysis to study price transmission in Pakistan's poultry supply chains. In light of arguments made by Frey and Manera (2007), we examine the frequency of the underlying poultry price adjustment process in Pakistan before specifying an appropriate data frequency for econometric analysis. Accordingly, in contrast to the pervasive use of monthly price data in the existing literature, our empirical analysis is based on weekly prices. Last, by employing identical research methods to study price transmission mechanisms of two different commodities (eggs and chicken) in similar institutional settings, we are able to better understand the effect of product characteristics on the patterns of price adjustment, particularly how product storability affects short-run price transmission mechanisms under conditions of producer market power. In summary, the combination of unique institutional settings, rigorous econometric methodology and high-frequency data characterize the novel contributions of our paper to the empirical price transmission literature.

Background

Pakistan is the sixth most populous country in the world and is categorized as a lower-middle-income economy with per capita GDP of $1,548 (World Bank, 2017). The poultry sector plays a pivotal role in Pakistan's economy since it accounts for 1.3% of GDP, provides jobs to 1.5 million people, and contributes 35% toward total meat production. Pakistan ranked among the top 20 poultry-producing countries worldwide, with an estimated annual production of 1.3 million metric tons of chicken and 12.9 billion eggs in 2017 (Government of Pakistan, 2017). Poultry production in Pakistan is on par with international standards; in fact, Pakistan's breeding stock is considered to be among the best in the world (U.S. Department of Agriculture, 2010).

Organization of Supply Chain and Mechanics of Price Determination

Private-sector investment and favorable government policies have transformed Pakistan's poultry production sector from small-scale, subsistence farming to a large-scale, technologically advanced industry. Barriers to entry created by the fixed costs of specialized machinery and economies of scale have led to high levels of concentration in the upstream sector of the poultry supply chain. Pakistan's poultry retailing and processing sectors have not experienced the structural changes witnessed in the poultry production business due to poor marketing infrastructure (particularly lack of cold-storage facilities), relatively low consumer incomes, and preference for freshly slaughtered chicken.[4] Consequently, the retail sector is fragmented and consists of small shops selling eggs or slaughtered chickens to consumers without any significant value addition (no grading, packaging, or storing). The processing sector is also largely missing from Pakistan's poultry supply chain, and branded eggs/chicken account for only a small portion (< 2%) of total sales volume.

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[3] As part of our fieldwork on Pakistan's poultry supply chains, we conducted 40 structured interviews with retailers, producers, and commission agents; two focus groups with producers and regulators; and several market visits.

[4] Concerns about the *halal* status of chicken amongst the large Muslim population are a major reason for this preference.

The government plays no role in the price determination process, and international trade in poultry products is negligible. Instead, egg and chicken prices are determined by interactions between producers and retailers in informal poultry auction markets known as *mandis*, which are typically located in the major poultry production belts in northern Pakistan and remain active throughout the year (since eggs and chicken are continuously produced commodities). Commission agents, present in each *mandi*, have information on retailers' demand and stock of chicken/eggs ready for sale in adjoining areas. These agents serve as intermediaries between farmers and retailers in exchange of a fixed commission. Due to lack of product differentiation, cold-storage facilities, and contracted production, price negotiations typically revolve around the prevailing supply/demand situation in regional *mandis*. Networks of commission agents and arbitrageurs, abetted by good transportation infrastructure and mobile phone networks, ensure that information on regional demand/supply conditions is aggregated into poultry prices across *mandis*.

Market Power of Poultry Producers and Antitrust Litigation

Basic economic principles suggest that concentration leads to market power; hence, regulators have raised concerns about the influence of poultry producers on farm-gate prices in Pakistan. The Competition Commission of Pakistan (CCP) initiated antitrust proceedings against the Pakistan Poultry Association (PPA) in 2010 following reports of price manipulation in the popular press.[5] As part of their inquiry, the CCP conducted several raids and claimed to have found evidence of tacit collusion in confidential documents seized from the PPA's regional offices. Accordingly, a fine of 50 million Pakistani rupees (Rs.) ($577,861) was levied on the PPA for violating Pakistan's competition laws.[6] Poultry producers denied the CCP's charges, arguing that the PPA served as a platform for disseminating farm-gate prices emerging from supply/demand conditions and that the perishability of poultry products along with the scale of poultry production in Pakistan made fixing poultry prices practically impossible. Subsequently, poultry producers obtained a "stay order" from Pakistan's High Court and were able to avoid the CCP's stipulated penalty. In 2016, the CCP served another show-cause notice to poultry producers for allegedly engaging in anticompetitive behavior and imposed a fine of Rs. 100 million ($1,086,957) on the PPA. Once again, poultry producers challenged the CCP's decree in the High Court on technical grounds. At present, the CCP's legal proceedings against the PPA linger in Pakistan's courts, with no clear resolution in sight.

Conceptual Framework: Asymmetric Price Transmission and Market Power

Theoretical research has shown that imbalances in bargaining powers of agents in a supply chain lead to asymmetric price transmission because agents with market power are able to delay price adjustments that diminish their profits and vice versa (Meyer and von Cramon-Taubadel, 2004). But the type of asymmetries depends on the specific institutional settings (Bakucs, Fałkowski, and Fertő, 2014). For example, McCorriston, Morgan, and Rayner (1998) prove that increases in farm-gate prices are passed on to consumers more fully than decreases in farm-gate prices if retailers have market power. Assefa, Kuiper, and Meuwissen (2014) show that the nature of asymmetric price transmission in an agricultural supply chain is determined by the relative bargaining powers of retailers and farmers. Adjustment costs are another explanation for asymmetric price transmission (i.e., differences in costs of activities such as relabeling and advertising associated with increasing and decreasing prices). For instance, retail prices are "sticky" and adjust slowly to price shocks due to "menu" costs (Frey and Manera, 2007). Other plausible, but less common, explanations for asymmetric price transmission include government interventions, inventory management methods, and information asymmetries.

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[5] PPA is an umbrella organization representing the interests of poultry producers in Pakistan at a national level.

[6] Conversion of the penalties into USD is based on the average exchange rate during the mentioned year.

Market power is generally viewed as the primary cause of asymmetric price transmission in the literature. Findings of asymmetric price transmission in empirical research are typically attributed to market power, even though reduced-form econometric models employed in the existing literature cannot explicitly identify the underlying cause of asymmetric price transmission (Lloyd, 2017). Nevertheless, despite the lack of causal inference in price transmission studies, there is a broad consensus among economists about the importance of price transmission analysis in identifying strategic behavior of agents in a supply chain (Meyer and von Cramon-Taubadel, 2004). Besides, the limited availability of data (beyond market prices) means that it is often not possible to estimate elaborate structural models capable of differentiating between competing theories of asymmetric price transmission. Thus, researchers have relied on anecdotal evidence or knowledge of institutional settings to rule out alternative explanations for asymmetric price transmission.

Hypothesis Development

The structure of poultry supply chains in Pakistan points toward a widening gulf between the bargaining powers of producers and retailers. However, high levels of concentration in upstream poultry markets are not sufficient to prove price manipulation, and the CCP has not been able to find concrete evidence of manipulation in farm-gate prices due to lack of proper documentation in Pakistan's poultry markets. Under these circumstances, price transmission analysis offers a viable path toward resolving the CCP's antitrust lawsuits against the PPA. The theoretical literature has shown that agents with market power in a supply chain influence the transmission of price shocks between upstream and downstream markets to increase their profits. Hence, in the context of Pakistan's poultry supply chains, estimates of the speed with which farm-gate prices adjust to positive and negative deviations from the long-run equilibrium can provide important insights about the alleged manipulation of farm-gate prices by producers. More precisely, if poultry producers are guilty of market power abuse, then positive shocks to farm-gate prices would tend to persist, while negative shocks to farm-gate prices would tend to quickly revert to the long-run equilibrium. It is reasonable to assume that empirical evidence of this type of asymmetric price transmission is caused by the market power of poultry producers because the mechanics of poultry price adjustments in Pakistan suggest that adjustment or "menu" costs are not significant. Lack of cold-storage facilities implies that explanations based on inventory management are not relevant. Agents in *mandis* have no market power and simply work for retailers/producers on the basis of a fixed commission. And government interventions in Pakistan's poultry markets are very rare.

Data

Data frequency plays a key role in price transmission analysis because estimates from econometric models of price transmission are misleading if there is a mismatch between the frequency of the underlying commodity's price adjustment process and the frequency of available price data (Frey and Manera, 2007). Meyer and von Cramon-Taubadel (2004) note that coefficients of error-correction models estimated from monthly price data are artificially inflated if the underlying commodity's price adjustment process follows a weekly frequency. Likewise, von Cramon-Taubadel and Loy (1996) find that contemporaneous relationships gain importance over lagged relationships if the frequency of price data is lower than the frequency of the underlying commodity's price-adjustment process. Yet despite the adverse effects of an inappropriate data frequency on price transmission analysis, issues pertaining to data frequency are often overlooked in the extant literature. To address this shortcoming, we study the nature of the price adjustment process in Pakistan's poultry markets before specifying an appropriate data frequency for our price transmission analysis.

   The nature of poultry production limits the efficacy of a monthly frequency because supply-side dynamics arising from the relatively short duration of the poultry production cycle are obscured (due to aggregation) in monthly price data. Anecdotal evidence on the mechanics of price discovery

in Pakistan's poultry markets suggests that new information on prices does not arrive daily. This is reflected in the daily price data, in which we observe variations in farm-gate prices in a given week but typically no change in farm-gate prices between consecutive days.[7] Presumably, in the absence of a commodity exchange for poultry products, a few days pass before new information related to supply/demand conditions is incorporated into poultry prices in Pakistan. A closer look at the data reveals that the price adjustment process in Pakistan's poultry markets is predominantly weekly. Theoretical models of poultry price dynamics in developing countries are also based on a weekly frequency, as opposed to monthly or daily (Chaudhry and Miranda, 2018). In light of these arguments, our analysis of price transmission mechanisms in Pakistan's poultry supply chain is based on weekly prices.

Retail prices were retrieved from the Pakistan Bureau of Statistics (PBS) database. The PBS undertakes a comprehensive survey of Pakistan's retail markets and publishes average retail prices of chicken and eggs in its weekly report. The corresponding farm-gate prices were sourced from the Pakistan Poultry Association (PPA), which records the average market-clearing farm-gate prices of chicken and eggs across different *mandis* in Pakistan on a daily basis. Daily farm-gate prices were converted to a weekly frequency by taking an arithmetic average over the PBS weekly reporting cycle of retail prices. The final dataset, spanning from July 2008 to June 2015, consists of 363 weekly observations of farm-gate and retail prices of chicken and eggs.

Line plots of egg (Figure 1(a)) and chicken (Figure 1(b)) prices highlight the dynamics of poultry prices in Pakistan. First, comovement of retail prices (dotted line) and farm-gate prices (solid line) in the long run points toward the integration of upstream and downstream markets in the egg and chicken supply chains. Second, poultry prices depict cyclical short-run dynamics; fluctuations in chicken prices are visibly larger than fluctuations in egg prices. Third, retailers' margins in the egg supply chain are considerably lower than those of retailers in the chicken supply chain (i.e., average retailer margin is 20.1% for chicken compared to 6.7% for eggs).

Econometric Methodology

Our econometric analysis follows the methodology outlined in the survey articles by Meyer and von Cramon-Taubadel (2004) and Frey and Manera (2007). First, we undertake unit root analysis to ascertain the order of integration of each price series. Second, we use the Johansen (1995) cointegration test and the two-step Engle and Granger (1987) approach to determine whether farm-gate and retail prices are cointegrated. Third, we employ threshold autoregressive (TAR) and momentum-threshold autoregressive (M-TAR) models to account for nonlinearities in the long-run relationship between farm-gate and retail prices. Threshold values in the consistent TAR and M-TAR models are obtained via grid-search methods. Each cointegration model is estimated via ordinary least squares (OLS) and an information criterion is used to select an appropriate model. Next, we conduct statistical tests to examine the nature of asymmetries in the long-run relationship between farm-gate and retail prices. Fourth, we estimate the associated vector error-correction model to study short-run dynamics of egg and chicken prices in Pakistan.

Analysis of Unit Roots and Cointegration

We employ the augmented Dicky–Fuller (ADF) and the Dicky–Fuller generalized least squares (DF-GLS) tests to check for unit roots and find the order of integration of weekly prices. The DF-GLS test, developed by Elliott, Rothenberg, and Stock (1996), is a modified version of the ADF test applied to GLS-detrended data. The literature has found the DF-GLS test to possess higher power

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[7] There is no change in farm-gate prices between consecutive days in 80% of the egg price data and 40% of the chicken price data. From a technical standpoint, in this case, estimates of price transmission models based on daily data would have large standard errors due to low levels of variation in the regressors. Moreover, fitting price transmission models on daily data is also problematic due to a significant increase in the number of parameters to be estimated.
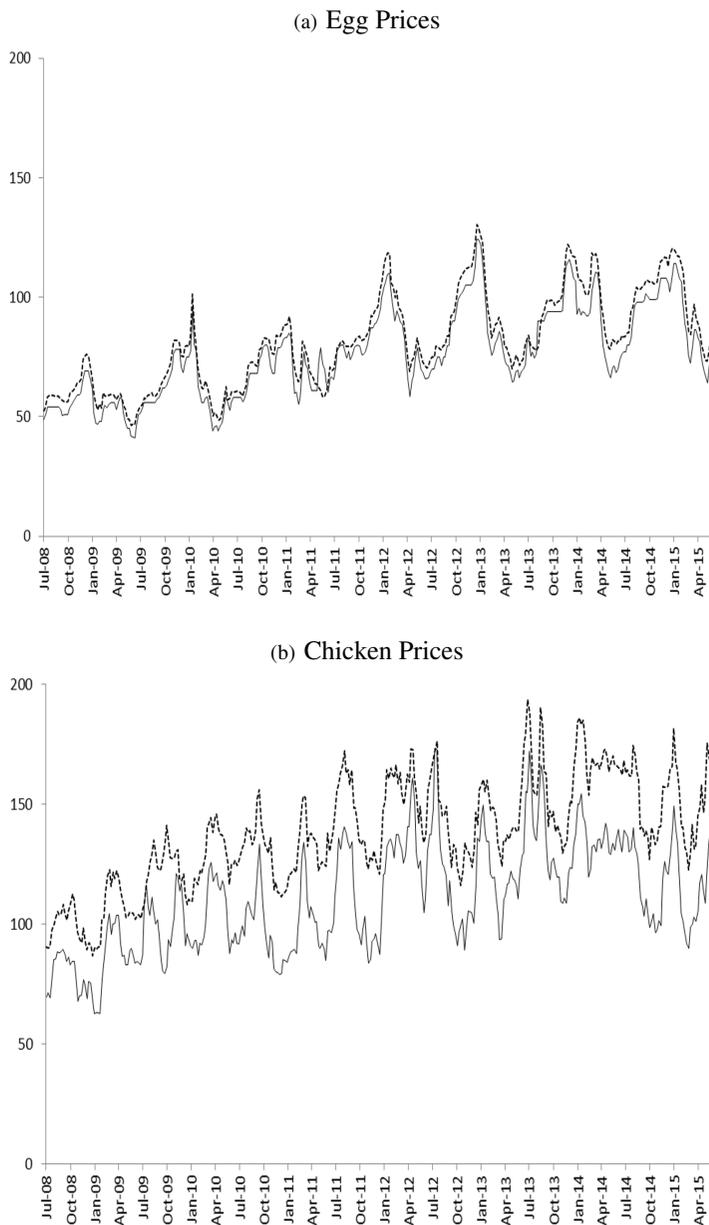
(a) Egg Prices

(b) Chicken Prices

Figure 1. Weekly Prices of Eggs and Chicken in Pakistan, July 2008 to June 2015

in detecting unit roots compared to the ADF test (Ng and Perron, 2001). For the sake of brevity, technical details of unit roots analysis are relegated to the Online Supplement (www.jareonline.org). The results of unit root tests (Table S1) reveal that each price series has a unit root and is integrated of order 1 (i.e., I(1)).

Cointegration refers to the tendency of nonstationary variables to move together over the long run, despite drifting apart in the short run, due to linkages between the stochastic data-generating process underlying each nonstationary variable (Hendry and Juselius, 2001). As a starting point, we use the Johansen (1995) cointegration test to determine whether there is a long-run relationship between farm-gate and retail prices in Pakistan's poultry supply chains. Results of the Johansen

cointegration test—provided in Panel A of Table S2—clearly show that farm-gate and retail prices in Pakistan's chicken and egg supply chains are cointegrated.

Engle and Granger (1987) developed another two-step procedure to test for cointegration between nonstationary variables. The first step involves estimating the long-run (i.e., cointegrating) relationship between two I(1) variables via OLS. In the second step, unit root tests are employed to determine whether residuals of this long-run relationship are stationary. These tests show that the underlying variables are cointegrated if the estimated residuals are stationary, and vice versa. Following their methodology, we apply unit root tests to residuals estimated from the cointegrating regression between farm-gate (P^f) and retail (P^r) prices in the chicken and egg supply chains:

(1) $$\hat{\mu}_t = P_t^f - \hat{\alpha} - \hat{\beta} P_t^r.$$

In consideration of our research question (i.e., the effect of farmers' market power on price transmission mechanisms), we normalize residuals of the cointegrating equation to farm-gate prices instead of retail prices.[8] Doing so allows us to interpret evidence of asymmetric price transmission as a sign of the exercise of market power by poultry producers (see the Conceptual Framework section). This normalization is also necessary from a time-series modeling perspective because retail prices were found to be weakly exogenous (see Table 3) and econometric theory dictates that the weakly exogenous variable be used as the independent variable in the cointegrating regression to avoid specification errors (Hendry and Juselius, 2001). Results of unit root tests applied to residuals ($\hat{\mu}_t$) of the cointegrating regression defined in equation (1) are provided in Panel B of Table S2. It is easy to see that residuals of both cointegrating regressions are stationary; hence, farm-gate and retail prices in Pakistan's egg and chicken supply chains are cointegrated according to the Engle–Granger (1987) test.[9]

Nonlinear Cointegration Analysis

Enders and Granger (1998) showed that the cointegration tests of Johansen (1995) and Engle and Granger (1987) perform poorly if the long-run relationship between the underlying variables is not symmetric (i.e., positive and negative deviations from the long-run equilibrium are not corrected in an identical manner). Enders and Siklos (2001) therefore developed an alternative approach based on TAR and M-TAR models to test for cointegration without assuming symmetric adjustment toward the equilibrium. Following the recent empirical price transmission literature (Korale Gedara, Ratnasiri, and Bandara, 2016; Simioni et al., 2013; Tekgüç, 2013), we adopt Enders and Siklos's (2001) two-regime threshold cointegration model to test for asymmetric adjustment in the long-run equilibrium relationship between farm-gate and retail prices in Pakistan's egg and chicken supply chains.

TAR and M-TAR Cointegration Models

TAR and M-TAR models are piecewise linear models that account for the asymmetric adjustment of positive and negative deviations from the long-run equilibrium. Cointegration tests based on the TAR and M-TAR models are grounded in the two-step Engle and Granger (1987) approach. Let $\hat{\mu}_t$ denote the residuals estimated from the long-run relationship between farm-gate and retail prices in the first step of Engle and Granger procedure (see equation 1). The corresponding TAR model is

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[8] Although it is a convention in the price transmission literature, the use of downstream prices as the dependent variable in the cointegrating equation lacks a sound basis. In fact, price transmission research (especially in developing countries) has shown that upstream prices are often determined by downstream prices (Bakucs, Fałkowski, and Fertő, 2014).

[9] Changing the order of variables in the cointegrating regressions did not affect the conclusions of the tests of cointegration. However, statistical properties of the underlying model are compromised (especially the associated VECM) if the weakly exogenous variable is not used as the independent variable in the cointegrating regression.

given by

$$\Delta\hat{\mu}_t = \rho_1 I_t \hat{\mu}_{t-1} + \rho_2 (1 - I_t)\hat{\mu}_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta\hat{\mu}_{t-j} + \in_t$$

(2)   such that

$$I_t = \begin{cases} 1 \text{ if } \hat{\mu}_{t-1} \geq \tau \\ 0 \text{ if } \hat{\mu}_{t-1} < \tau \end{cases}$$

The indicator variable (I_t) in the TAR model allows for asymmetric adjustment toward the long-run equilibrium depending on the "size" of deviations. Estimates of ρ_1 and ρ_2 denote the speed at which the variables adjust to the equilibrium following a positive or negative shock, respectively. Enders and Siklos (2001) proposed the M-TAR specification for modeling asymmetrically sharp movements in the data, whereby adjustment toward the equilibrium depends on the "momentum" of deviations and not the "size" of deviations. Accordingly, the indicator variable (I_t) in the M-TAR model is a function of the change in deviations from the equilibrium instead of the level of deviations from the equilibrium as in the TAR model. They showed that M-TAR models perform better than TAR models if prices display more momentum in one direction (e.g., increases tend to persist and decreases tend to revert to equilibrium). The M-TAR model is given by

$$\Delta\hat{\mu}_t = \rho_1 I_t \hat{\mu}_{t-1} + \rho_2 (1 - I_t)\hat{\mu}_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta\hat{\mu}_{t-j} + \in_t$$

(3)   such that

$$I_t = \begin{cases} 1 \text{ if } \Delta\hat{\mu}_{t-1} \geq \tau \\ 0 \text{ if } \Delta\hat{\mu}_{t-1} < \tau \end{cases}$$

TAR and M-TAR can be consistently estimated via OLS. The Akaike information criterion (AIC) is used to determine the optimal number of lags of the dependent variable in each model and the threshold value (τ) is assumed to equal 0 (i.e., mean of the residuals of the cointegrating regression). Rejection of the null hypothesis of no cointegration ($\rho_1 = \rho_2 = 0$) provides evidence of a long-run relationship between the price series. The F-statistic for the null hypothesis $\rho_1 = \rho_2 = 0$ does not follow a standard distribution under the TAR and M-TAR models; hence, adjusted critical values tabulated in Enders and Siklos (2001) are used for statistical inference. If the price series are found to be cointegrated, then the standard F-test can be used to evaluate the null hypothesis of symmetric adjustment (i.e., $\rho_1 = \rho_2$). Rejection of this null hypothesis provides evidence of asymmetries in the long-run relationship between farm-gate and retail prices.

Consistent TAR and M-TAR Cointegration Models

One limitation of TAR and M-TAR models is that the mean of the residuals of the cointegrating regression may not be an unbiased estimate of the threshold (τ) if the adjustment process is asymmetric (Enders and Siklos, 2001). Hence, the threshold value is estimated in consistent TAR and M-TAR models. Chan (1993) showed that grid-search methods yield a super-consistent estimate of the threshold (τ). Applying this approach, we sort the values of threshold variables ($\hat{\mu}_{t-1}$ in TAR and $\Delta\hat{\mu}_{t-1}$ in M-TAR) in ascending order. The smallest and largest values of the threshold variable are excluded from the set of potential threshold values to ensure an adequate number of observations in each regime of the TAR/M-TAR models. Following Balcombe, Bailey, and Brooks (2007), we use the middle 60% values of the sorted threshold variable as potential threshold values. The TAR and

M-TAR models are estimated for each potential threshold value and a grid-search algorithm is used to identify the value of the threshold variable that minimizes the sum of squared residuals ($\hat{\in}_t$) in each model. Once the optimal threshold ($\hat{\tau}$) is determined, each model is re-estimated with the optimal threshold value taken as given due to the super-convergence property of the estimator and the usual statistical analysis is carried out:

$$\Delta\hat{\mu}_t = \rho_1 I_t \hat{\mu}_{t-1} + \rho_2 (1 - I_t)\hat{\mu}_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta\hat{\mu}_{t-j} + \in_t$$

(4)   such that

$$I_t = \begin{cases} 1 \text{ if } \hat{\mu}_{t-1} \geq \hat{\tau} \\ 0 \text{ if } \hat{\mu}_{t-1} < \hat{\tau} \end{cases}$$

$$\Delta\hat{\mu}_t = \rho_1 I_t \hat{\mu}_{t-1} + \rho_2 (1 - I_t)\hat{\mu}_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta\hat{\mu}_{t-j} + \in_t$$

(5)   such that

$$I_t = \begin{cases} 1 \text{ if } \Delta\hat{\mu}_{t-1} \geq \hat{\tau} \\ 0 \text{ if } \Delta\hat{\mu}_{t-1} < \hat{\tau} \end{cases}$$

Consistent TAR (equation 4) and consistent M-TAR (equation 5) models are employed together with TAR (equation 2) and M-TAR (equation 3) models to examine the relationship between upstream and downstream prices in Pakistan. Determining which nonlinear cointegration model best captures the long-run equilibrium relationship between the underlying variables is an empirical question, and an information criterion is used to identify the model that best fits the actual price data (Enders and Siklos, 2001).

Vector Error-Correction Model

According to the Engle–Granger representation theorem, the short-run dynamics of cointegrated variables can be studied by a vector error-correction model (VECM). In a VECM, changes in the values of an I(1) variable at any point in time are determined by deviations from the long-run equilibrium in the previous period and changes in the past values of all variables in the system. Therefore, in addition to the dynamic effects of other variables in the system, a VECM captures how each variable adjusts toward the long-run equilibrium following any short-run disequilibria. Estimates from a VECM provide important insights about the short-run mechanisms of price transmission. First, weak exogeneity tests identify the "leading" price in a supply chain. Second, Granger causality tests reveal the direction of information feedback between upstream and downstream markets. Third, coefficients of error-correction terms highlight asymmetries in the speed of adjustment toward the long-run equilibrium following a positive or negative shock.

We estimate separate VECMs for the egg and chicken supply chains of Pakistan. Following the price transmission literature, we break down changes in farm-gate prices, retail prices, and deviations from the long-run equilibrium into positive and negative components in each VECM to focus on asymmetries in short-run price dynamics. Error-correction terms (i.e., *ECT*) in both VECMs are constructed from the nonlinear cointegration model that best fits the actual price data. *ECT* below the estimated threshold represent negative deviations of farm-gate from the long-run equilibrium and *ECT* above the estimated threshold denote positive deviations of farm-gate prices from the long-run equilibrium. The Akaike information criterion (AIC) is used to select the optimal number of lags and specification of deterministic components is based on the Pantula (1989) principle. The general

Table 1. Estimates of Threshold Cointegration Models in the Egg Supply Chain

	Threshold	Momentum	Threshold-Consistent	Momentum-Consistent
	1	**2**	**3**	**4**
ρ_1	−0.225***	−0.178***	−0.220***	−0.181***
	(0.045)	(0.050)	(0.044)	(0.046)
ρ_2	−0.314***	−0.337***	−0.329***	−0.367***
	(0.057)	(0.049)	(0.059)	(0.054)
τ	0.000	0.000	−1.826	−1.323
AIC	1,582	1,578	1,581	1,576
BIC	1,589	1,586	1,589	1,584
Φ^a	27.63	29.73	28.05	30.72
$\rho_1 = \rho_2$	1.48	5.14**	2.22	6.87***

Notes: These estimates are based on weekly prices from July 2008 to June 2015, for a total of 363 observations. Parameter estimates are reported adjacent to the corresponding row variables. Numbers in parentheses are standard errors. The information criterion is minimized in both TAR and M-TAR models with no lagged differences of errors in the right side of the cointegrating equation. Single, double, and triple asterisks (*, **, ***) indicate statistical significance at the 10%, 5%, and 1% level. Entries in the last row represent the values of the standard *F*-statistic for the null hypothesis of symmetric adjustment (i.e., $\rho_1 = \rho_2$).
^a Entries in this row are the values of the *F*-statistic (i.e., null hypothesis of $\rho_1 = \rho_2 = 0$) in each model specification. Associated critical values (at the 1% level) for each model specification, taken from Table 1 and Table 5 of Enders and Siklos (2001), are as follows: column 1, 8.08; column 2, 8.61; column 3, 9.15; and column 4, 8.82.

specification of a VECM for price transmission analysis is given by

$$\Delta P_t^f = c^1 + \varphi^{1+} ECT_{t-1}^+ + \varphi^{1-} ECT_{t-1}^- + \sum_{j=1}^{k} \phi_{j,1}^{1+} \Delta P_{t-j}^{f+} + \sum_{j=1}^{k} \phi_{j,1}^{1-} \Delta P_{t-j}^{f-}$$

$$+ \sum_{j=1}^{k} \phi_{j,1}^{2+} \Delta P_{t-j}^{r+} + \sum_{j=1}^{k} \phi_{j,1}^{2-} \Delta P_{t-j}^{r-} + \in_t^1;$$

(6)

$$\Delta P_t^r = c^2 + \varphi^{2+} ECT_{t-1}^+ + \varphi^{2-} ECT_{t-1}^- + \sum_{j=1}^{k} \phi_{j,2}^{1+} \Delta P_{t-j}^{f+} + \sum_{j=1}^{k} \phi_{j,2}^{1-} \Delta P_{t-j}^{f-}$$

$$+ \sum_{j=1}^{k} \phi_{j,2}^{2+} \Delta P_{t-j}^{r+} + \sum_{j=1}^{k} \phi_{j,2}^{2-} \Delta P_{t-j}^{r-} + \in_t^2.$$

Empirical Estimates

Nonlinear Cointegration Models

Estimates of the nonlinear cointegration models for egg prices are reported in Table 1. A specification with no lagged differences of residuals $(\Delta \hat{\mu}_{t-j})$ in the cointegration model best fits the egg price data according to the information criterion. The null hypothesis of no cointegration $(\rho_1 = \rho_2 = 0)$ between farm-gate and retail prices is rejected at the 1% significance level in all models. The long-run relationship between farm-gate and retail prices in the egg supply chain follows the M-TAR model as the value of the AIC is minimized in the consistent M-TAR model and subsequent statistical analysis of egg prices focuses on estimates from the consistent M-TAR model.

Estimates of ρ_2 are considerably larger than ρ_1 in all models. Differences between the speeds of adjustment to positive and negative shocks are statistically significant (i.e., the null hypothesis of

Table 2. Estimates of Threshold Cointegration Models in the Chicken Supply Chain

	Threshold	Momentum	Threshold-Consistent	Momentum-Consistent
	1	2	3	4
ρ_1	−0.231***	−0.226***	−0.211***	−0.194***
	(0.049)	(0.052)	(0.047)	(0.047)
ρ_2	−0.227***	−0.232***	−0.256***	−0.281***
	(0.052)	(0.049)	(0.056)	(0.056)
γ_1	−0.069	−0.070	−0.069	−0.070
	(0.053)	(0.053)	(0.053)	(0.053)
τ	0.000	0.000	−6.452	−2.378
AIC	2,275	2,275	2,275	2,274
BIC	2,287	2,287	2,286	2,285
Φ[a]	19.23	19.23	19.44	20.08
$\rho_1 = \rho_2$	0.00	0.01	0.39	1.56

Notes: These estimates are based on weekly prices from July 2008 to June 2015, for a total of 363 observations. Parameter estimates are reported adjacent to the corresponding row variables. Numbers in parentheses are standard errors. The information criterion is minimized in both TAR and M-TAR models with one lagged difference of errors in the right side of the cointegrating equation. Single, double, and triple asterisks (*, **, ***) indicate statistical significance at the 10%, 5%, and 1% level. Entries in the last row represent the values of the standard *F*-statistic for the null hypothesis of symmetric adjustment (i.e., $\rho_1 = \rho_2$).
[a] Entries in this row are the values of the *F*-statistic (i.e., null hypothesis of $\rho_1 = \rho_2 = 0$) in each model specification. Associated critical values (at the 1% level) for each model specification, taken from Table 1 and Table 5 of Enders and Siklos (2001), are as follows: column 1, 8.08; column 2, 8.61; column 3, 9.15; and column 4, 8.82.

symmetric adjustment ($\rho_1 = \rho_2$) is rejected in the consistent M-TAR model at the 1% significance level). This shows that positive deviations (ρ_1) of farm-gate prices from the long-run equilibrium are relatively slowly corrected compared to negative deviations (ρ_2). The asymmetries in the cointegrating relationship are consistent with the exercise of market power at the upstream level of the egg supply chain. Producers with market power are able to sustain farm-gate prices above the long-run equilibrium for a relatively longer period while ensuring that negative deviations of farm-gate prices from the long-run equilibrium are quickly eliminated. More specifically, estimates from the consistent M-TAR model imply that approximately 37% of a unit negative shock from the long-run equilibrium is eliminated within 1 week compared to only 18% in the case of a unit positive shock. Moreover, in the context of M-TAR models, the results show that the long-run price adjustment process between upstream and downstream prices displays lopsided momentum such that increases in upstream prices above the long-run equilibrium tend to persist but decreases below the long-run equilibrium are quickly reversed. Both these findings lend support to the CCP's claims regarding the exercise of market power by producers in the egg supply chain.

Table 2 reports estimates of nonlinear cointegration models for chicken prices. A specification with one lagged difference of residuals ($\Delta\hat{u}_{t-j}$) in the cointegration model best fits the chicken price data according to the information criterion. The null hypothesis of no cointegration ($\rho_1 = \rho_2 = 0$) between farm-gate and retail prices is rejected at the 1% level. The consistent M-TAR model is our preferred specification for chicken prices since it yields the minimum value of the AIC. We do not find evidence of asymmetries in the long-run relationship between upstream and downstream prices in the chicken supply chain because differences between the speeds of adjustment for positive and negative shocks are not statistically significant (i.e., the null hypothesis of symmetric adjustment cannot be rejected at conventional levels of statistical significance (*p*-value = 0.21)).

Table 3. Estimates of Vector Error-Correction Model

	Egg Supply Chain		Chicken Supply Chain	
	Farm Price (ΔP_t^f)	Retail Price (ΔP_t^r)	Farm Price (ΔP_t^f)	Retail Price (ΔP_t^r)
ECT_{t-1}^+	−0.106	0.038	−0.199***	−0.031
	(0.112)	(0.117)	(0.058)	(0.058)
ECT_{t-1}^-	−0.372***	−0.031	−0.357***	−0.143**
	(0.076)	(0.076)	(0.07)	(0.069)
$\Delta P^f{}_{t-1}^+$	0.047	−0.014	0.144	0.230**
	(0.135)	(0.135)	(0.096)	(0.095)
$\Delta P^f{}_{t-1}^-$	−0.075	0.009	0.191*	0.166
	(0.137)	(0.137)	(0.11)	(0.109)
$\Delta P^r{}_{t-1}^+$	0.318**	0.199	0.244**	−0.127
	(0.134)	(0.134)	(0.111)	(0.109)
$\Delta P^r{}_{t-1}^-$	0.815***	0.543***	0.246**	−0.026
	(0.154)	(0.155)	(0.122)	(0.12)
Q(4)-Test	7.052	7.032	1.440	1.981
	(0.133)	(0.134)	(0.838)	(0.734)
R^2	0.308	0.165	0.227	0.041

Notes: These estimates are based on weekly prices from July 2008 to June 2015, for a total of 363 observations. The estimated vector error-correction models (VECM) are based on the consistent M-TAR model. Information criterion was minimized in the VECM with one lag for both the egg and chicken prices. Parameter estimates are reported adjacent to the corresponding row variables. Numbers in parentheses are standard errors. Single, double, and triple asterisks (*, **, ***) indicate statistical significance at the 10%, 5%, and 1% level. The portmanteau (or Q) test for white noise is based on 4 lags of the residuals, where serially uncorrelated errors is the null hypothesis of the portmanteau test. The associated *p*-value is reported below the test statistic in parentheses.

Vector Error-Correction Models (VECM)

Estimates of VECMs in Table 3 are based on the consistent M-TAR model (i.e., the cointegration model that best fits the actual price data). Selection of an optimal lag length of 1 week in both VECMs strengthens arguments about the importance of high-frequency data in price transmission analysis. VECMs for egg and chicken prices are well specified, as evidenced by the negative and statistically significant coefficients of error-correction terms.[10] Results of the Q test show that residuals of both VECMs are not serially correlated.

VECM for Egg Prices

The left panel of Table 3 reports estimates of the VECM for egg prices, with dependent variables in the columns and independent variables in the rows. Coefficients of the error-correction terms (ECT_{t-1}^+ and ECT_{t-1}^-) in the retail price equation are not statistically significant (i.e., the *F*-test for joint insignificance yields a *p*-value of 0.87). Thus, retail prices are weakly exogenous in the long run and serve as the "leading" price in the egg supply chain. Granger causality tests show that retail prices drive farm-gate prices in the short run since lagged changes of farm-gate prices in the retail price equation are not statistically significant but lagged changes of retail prices in the farm-gate price equation are highly significant. Taken together, these results point toward strong exogeneity

[10] In a well-specified VECM, adjustment coefficients in at least one equation are negative and statistically significant.

of retail prices in Pakistan's egg supply chain. Hence, the dynamics of egg prices in Pakistan are essentially determined by shocks at the downstream level of the supply chain. Consistent with the results of nonlinear cointegration models, the null hypothesis of symmetric adjustment of farm-gate prices toward the long-run equilibrium is rejected in the VECM. Estimates of error-correction terms in the farm-gate equation clearly show that the speed of adjustment for deviations below the long-run equilibrium is significantly greater than the speed of adjustment for deviations above the long-run equilibrium.

VECM for Chicken Prices

The right panel of Table 3 reports estimates of the VECM for chicken prices. Retail prices enter the cointegration space as a weakly exogenous variable (i.e., F-test for joint insignificance of error-correction terms in the retail price equation yields a p-value of 0.11). Thus, we conclude that retail prices serve as the "leading" price in the chicken supply. Lagged changes of farm-gate prices are statistically significant in the retail price equation and lagged changes of retail prices are also statistically significant in the farm-gate price equation. This points toward bidirectional Granger causality between retail and farm-gate prices in the chicken supply chain. The null hypothesis of symmetric speed of adjustment of farm-gate prices toward the long-run equilibrium is rejected, albeit only marginally (p-value = 0.08). Analysis of the distributed-lag effect reveals strategic behavior of retailers vis-à-vis consumers in the short run (i.e., increases in farm-gate prices are more fully transmitted to retail prices relative to decreases in farm-gate prices).

Discussion of Results

Estimates from the long- and short-run models of price transmission provide a clear picture of the price adjustment mechanisms in Pakistan's chicken and egg supply chains. First, estimates from cointegration models clearly show that farm-gate and retail prices move together over the long-run in both supply chains.

Second, results of weak exogeneity tests from VECMs reveal that the long-run dynamics of chicken and egg prices in Pakistan are driven by shocks originating from downstream markets (i.e., upstream markets adjust to price signals emanating from downstream markets). The unidirectional causality from downstream markets to upstream markets in the long run implies that changes in retail demand affect retail prices and hence farm-gate prices.[11] However, changes in farm supply only lead to short-run disequilibria and do not affect retail prices in the long run. These results are consistent with the empirical literature on price transmission in developing countries. For example, Bakucs and Fertő (2005) document the weak exogeneity of retail prices in the Hungarian pork industry. Sapkota et al. (2015) also find that retail prices are weakly exogenous in Bangladesh's fish supply chains. Intuitively, the derived (as opposed to primary) nature of demand for agricultural products explains the importance of retail markets in the long-run dynamics of agricultural prices in developing countries.

Third, Granger causality tests reveal unidirectional causality from retail to farm-gate prices in the egg supply chain but bidirectional causality between retail and farm-gate prices in the chicken supply chain. In other words, the short-run dynamics of egg prices are driven by retail demand, while short-run dynamics of chicken prices emerge from the interaction of farm supply and retail demand. These differences can be explained by the biological processes underlying commercial production of eggs and chickens. The relatively longer life cycle of layer hens results in a stable productive stock and hence a steady short-run supply of eggs.[12] Additionally, eggs can be stored for up to 2 months. Thus, short-run fluctuations in egg prices are primarily caused by retail-demand shocks. However,

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[11] This validates our approach of using retail prices as the independent variable in the cointegrating regression.

[12] Eggs are produced by layer hens. In general, a layer hen enters egg production when it is 21 weeks old and continues to lay 3–4 eggs per week (on average) until it is culled at an age of about 72 weeks.

the relatively shorter life cycle of broilers results in an unstable productive stock and hence a volatile short-run supply of broilers.[13] Moreover, in the absence of a frozen-bird market, even small shocks to farm supply or retail demand lead to an imbalance between the supply of and demand for chicken, causing large fluctuations in chicken prices. Thus, we observe bidirectional causality between farm-gate and retail prices in the chicken supply chain.

Fourth, given that asymmetric price transmission is indicative of market power, estimates of adjustment parameters from nonlinear cointegration models and the associated VECM clearly point toward the exercise of market power by egg producers in Pakistan, whereby negative deviations of farm-gate egg prices from the long-run equilibrium are corrected more quickly than positive deviations. These findings are consistent with the industrial organization of egg supply chain in Pakistan (i.e., large, organized producers and small, fragmented retailers). However, we do not find statistically significant evidence of asymmetric price transmission in the chicken supply chain. Stylized facts of the price data and anecdotal evidence supports the conclusion that egg producers are guilty of market-power abuse. For example, concentration ratios in the upstream sector of the egg supply chain are relatively higher than concentration ratios in the upstream sector of the chicken supply chain. Likewise, the significantly lower margins of retailers (6.7% on average) in the egg supply chain compared to margins of retailers (20.1% on average) in the chicken supply chain also point toward the exercise of market power by egg producers over retailers. Further, we observe considerably less variation in egg prices over time compared to chicken prices (see Figure 1), and it is well known that sticky prices are a feature of oligopolistic markets.

Differences in the attributes of both commodities offer a plausible explanation for producers being able to exercise market power in the egg supply chain but not in the chicken supply chain, despite similarities in the industrial organization of both supply chains. The perishability of broiler chicken reduces the bargaining power of producers over retailers because producers are eager to sell their mature stocks as soon as possible due to the high risks of mortality in broilers after maturity. As a result, producers are unable to influence farm-gate prices by exercising their (perceived) market power in the chicken supply chain. On the contrary, storability of eggs provides producers in the egg supply chain latitude to favorably influence farm-gate prices by exerting their market power over retailers (e.g., by "holding" supply). The availability of close substitutes for chicken (e.g., beef, lamb, fish, etc.) also diminishes the market power of producers in the chicken supply chain, while the market power of producers in the egg supply chain is reinforced by the lack of close substitutes for eggs. One may argue that the characteristics of the underlying commodities moderate the effect of producer market power on price transmission mechanisms.

Conclusion

This paper examines the mechanisms of price transmission in Pakistan's egg and chicken supply chains. In doing so, we highlight the importance of price transmission analysis in unravelling regulatory lawsuits on the alleged abuse of market power by producers in agricultural markets. Thus, our paper serves as a prototype of the elusive "smoking gun" case study described in Meyer and von Cramon-Taubadel's (2004) influential survey of price transmission research.

We find that upstream and downstream markets in Pakistan's poultry supply chains are integrated. We find strong evidence of asymmetric price transmission in the egg supply chain, consistent with the CCP's claims regarding the exercise of market power by egg producers. We do not find sufficient statistical evidence to accept the CCP's claim that producers exercise market power in the chicken supply chain. Differences between the short-run dynamics of chicken and egg prices can be traced to the underlying production and marketing activities in each supply chain.

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[13] Chicks hatch from fertilized eggs after an incubation period of 3 weeks and are subsequently raised into broilers after another 6–7 weeks. These broilers are sold in live-bird markets as soon as they reach maturity due to lack of cold-storage facilities and consumer preferences for freshly slaughtered chicken in Pakistan.

From a modeling perspective, our paper highlights the importance of specification tests for weak-exogeneity, high-frequency data and incorporating nonlinearities in price transmission analysis.

From a policy perspective, our findings highlight important differences between the functioning of agricultural markets in developed and developing countries. For instance, contrary to traditional models of industrial organization, we find that the long-run dynamics of egg and chicken prices are largely determined by downstream markets in Pakistan. Similarly, we find evidence of market power at the upstream level of egg supply chains in Pakistan, even though price transmission research on developed countries has generally found evidence of market power at the downstream level of agricultural supply chains. Therefore, we need to revisit existing theoretical models to better understand the effects of the unique institutional features of agricultural supply chains in developing countries on price transmission mechanisms.

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Online Supplement:
Price Transmission in
Pakistan's Poultry Supply Chain

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Unit Root Analysis

Results of the ADF and DF-GLS test clearly show that egg prices are nonstationary in levels. The results are less clear-cut in case of chicken prices (i.e., the null of nonstationarity is not rejected by the DF-GLS test but is rejected by the ADF test). Conclusions of the DF-GLS test appear to be more reliable in light of the low power of ADF tests in detecting unit roots and statistical properties of chicken prices. For example, visual inspection of line plots in Figure 1 highlights changes in volatility of chicken prices over time. Additionally, analysis of the autocorrelation function of chicken prices reveals that autocorrelations do not decline with lag length. Both observations point toward a nonstationary data generating process. Accordingly, we conclude that each price series is integrated of order 1 (i.e., I(1)) since the null of nonstationarity is not rejected in levels but rejected (under both tests) after taking first differences.

Table S1. Results of Unit Root Tests

	Augmented Dicky–Fuller Unit Root Test		Dicky–Fuller (GLS) Unit Root Test	
	Levels	First Difference	Levels	First Difference
Egg farm-gate prices	−2.465	−12.464***	−1.382	−3.813***
Egg retail prices	−2.195	−12.919***	−1.226	−6.216***
Broiler farm-gate prices	−4.508***	−12.870***	−1.497	−7.189***
Broiler retail prices	−3.174**	−11.828***	−1.454	−7.298***

Notes: Existence of a unit root is the null hypothesis under the ADF and DF-GLS tests. The Aikake information criterion (AIC) was used to select the optimal number of lags in the ADF test. A modified AIC (MAIC) and Ng–Perron sequential t method were used to determine the optimal number of lags in the DF-GLS test. Experiments with fewer or more lags yielded similar results. Single, double, and triple asterisks (*, **, ***) indicate statistical significance at the 10%, 5%, and 1% level.

Linear Cointegration Analysis

Johansen (1995) showed that cointegration between two I(1) variables can be examined by looking at the characteristic roots of the underlying vector error-correction model (VECM) . He derived two log-likelihood ratio tests (i.e., trace and maximum Eigenvalue test) to study the number of cointegrating relationships (r) between a set of I(1) variables. Following his approach, we use the Pantula (1989) principle to specify the deterministic components of the underlying VECM and Akaike information criterion (AIC) to select the optimal number of lags. Values of the trace and maximum Eigenvalue statistics in Panel A of Table S2 clearly show that the null of cointegration between farm-gate and retail prices in Pakistan's chicken and egg supply chains cannot be rejected at conventional levels of statistical significance.

Table S2. Results of Linear Cointegration Tests

Panel A. Johansen (1995) Cointegration Test	Statistic ($r = 1$)	Max Statistic ($r = 1$)
Farm-gate and retail prices: egg	8.898	0.149
Farm-gate and retail prices: chicken	0.054	0.111
Panel B. Engle–Granger (1987) Cointegration Test	**ADF**	**DF-GLS**
Cointegrating equation residuals: egg	−7.318***	−4.967***
Cointegrating equation residuals: chicken	−6.201***	−3.048***

Notes: In Panel A, two lags were selected in the underlying vector error-correction models (VECM) for both egg and chicken prices based on AIC. Neither the trace or max statistic follows a χ^2 distribution; hence, critical values for the trace statistic (at the 5% level) are provided by Johansen (1995) as follows: 9.42 (egg series) and 3.84 (chicken series). In Panel B, the existence of a unit root is the null hypothesis under both the ADF and DF-GLS tests. The Akaike information criterion (AIC) was used to select the optimal number of lags in the ADF test. A modified AIC (MAIC) and Ng–Perron sequential t method were used to determine the optimal number of lags in the DF-GLS test. Single, double, and triple asterisks (*, **, ***) indicate statistical significance at the 10%, 5%, and 1% level.

[Received April 2019; final revision received October 2019.]

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