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Exploiting the Cointegration Properties of U.S. Pork-Related Markets: The Emergence of a U.S. Demand for Pork as an Input

Ronald A. Babula and Mogens Lund

Abstract

We apply methods of the cointegrated vector autoregression or VAR model to quarterly U.S. pork-related markets, from the farm gate upstream, to the downstream markets for processed pork and sausage. This study extends prior VAR econometric work by concentrating on the upstream/downstream relationships between the U.S. farm market for pork and markets for processed pork and sausage. Results include a U.S. long run demand for pork, as well as empirical estimates of specific market events on these three pork-related markets.

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INTRODUCTION, MODELED MARKETS, AND PRIOR RESEARCH

There has been increasing research interest in empirically illuminating inter-relationships that connect systems of markets related to a particular meat, from the upstream farm market to the downstream manufacturing/processor markets that use the meat as an input. Lloyd et. al. (2006, 2001) and Sanjuan and Dawson (2003) attribute the increased interest to heightened public awareness from food scares, particularly the 1996 outbreak of bovine spongiform encelopathy (BSE) and its human analog, Creutzfeldt-Jakub disease that claimed 10 British lives. As a result, herds were liquidated; UK beef exports all but ceased; and as well, demand and farm prices were disastrously affected. Andersen et. al. (2007) suggest that three other food scares have likely led to heightened interest in such meat-related market transmissions:

- Subsequent BSE outbreaks in Canada and the United States that led to severe declines in demand, trade, and price of beef, with burdensome impacts on host countries’ livestock producers.

- A 1985 U.S: salmonella outbreak that elicited some 16,000 cases of illness, and

- Asian influenza outbreaks that have claimed more than 140 lives since 2003.

The second reason is a growing awareness of the value that policy makers, agribusiness strategists, and researchers (hereafter, interested agents) place on reliable empirical estimation of commodity-based inter-relationships among markets related to a meat, as noted for U.S. chicken products by Goodwin, McKenzie, and Djunaidi (2003). Because the United States is a “large economic country,” changes in one market’s price or quantity, say the price of a pork-using product, have policy-relevant effects for pork farmers upstream. The profession consequently needs models that capture not only the U.S. farm gate market relationships, but relationships in related downstream markets that use the farm pork production as well. We contend that building such an econometric model to capture a balanced array of upstream and downstream relationships is often a tailor-fit task for Johansen’s (1988) and Johansen and Juselius’ (1990, 1992) methods of the cointegrated vector autoregression (cointegrated VAR) model.

Despite an explosion of cointegrated VAR modeling applications to upstream and downstream market systems related to many U.S. farm commodities, there have been
few applications to U.S. pork-related markets. To mitigate this analytical gap, we propose modeling the quarterly system of U.S. markets for pork at the farm gate and for as many pork-related products as available data resources permit. Resulting demand/supply elasticity estimates and/or empirical estimates of event-specific market events that emerge from such a cointegrated VAR model of a pork-related market system are likely of interest to the above-noted agents.

We located three VAR econometric studies of U.S. pork-related market transmissions of direct relevance to our study. Wang and Bessler (2006) focused on causal price/quantity meat market transmission relationships and modeled the following quarterly system of nonstationary and logged variables as a cointegrated VAR where the information inherent in contemporaneously correlated residuals was utilized by an application of directed acyclic graph or DAG analysis: nominal U.S. retail prices of pork, beef and poultry; quantities of these three meats; two price indexes of non-meat food and other non-food good prices; and nominal food consumption expenditures. Their pork-related results suggested evidence of U.S. pork consumption quantity’s relative predeterminedness relative to retail price, a finding that they then related to appropriate pork demand model specification.

Wang and Bessler (2003) combined cointegrated VAR modeling with DAG analysis to build a quarterly, logged model similar to that modeled by Wang and Bessler (2006) system. They then compared the cointegrated VAR’s out-of-sample forecast accuracy levels with such forecast performances of four alternative model frameworks of the same system, and found that cointegrated VAR model forecasted most accurately. Among their pork-related results was an own-price elasticity of per capita U.S. pork consumption of -0.50 that emerged from their error correction space estimates.

Bessler and Akleman (1998) combined Bernanke’s methods of the levels-based structural VAR with DAG analysis to utilize information from contemporaneously correlated residuals, and applied the framework to a U.S. monthly, logged system of the general U.S. price level, as well as retail pork, beef, and chicken prices; pork and beef farm values; manufacturing wages; and gasoline prices. By focusing on dynamic farm/retail dynamic relationships, their analysis of selected patterns of FEV decompositions suggested that farm-level shocks are primary explanators of retail pork price behavior, and revealed patterns U.S. pork/beef market causality.
Five sections follow. First is a discussion of cointegrated VAR econometrics and data used in this study. The second provides our specification efforts and diagnostic evidence that demonstrate the statistical adequacy of the model before cointegration tests were implemented and prior to the exploitation of any cointegration properties. The third section provides a discussion of the choice of reduced rank supported by a number of evidence sources. The fourth section discusses the application of Johansen and Juselius’ (1990, 1992) hypothesis tests and inference on the error correction space parameters, in order to provide empirical estimates of own- and cross-market parameters for U.S. pork-based markets. We then provide a summary and conclusions.

TIME SERIES ECONOMETRICS AND DATA

It is well-known that econometric time series often fail to meet the conditions of stationarity and ergodicity required of valid inference (Wang and Bessler 2006, p. 510; Hendry 1986). And while data series are often individually nonstationary, they can form vectors with stationary linear combinations, such that the inter-related series are “cointegrated” and move in tandem as an error-correcting system (Johansen and Juselius 1990, 1992).

We propose modeling the following endogenous variables (denoted throughout by the parenthetical labels) defined and sourced as follows as a cointegrated VAR model:

- U.S. market-clearing quantity of pork meat (QUSPORK) defined as the sum of beginning stocks, production, and imports from the U.S. Department of Agriculture, Economic Research Service (USDA, ERS 2006a, b, c).

- U.S. farm price of pork (PUSFARM). This is the U.S. producer price index or PPI, for slaughter hogs, farm products group, series number WPU0132. This is available from the U.S. Department of Labor, Bureau of Labor Statistics (Labor, BLS 2006).

- U.S. wholesale price for processed pork meat, reflected by the U.S. PPI for fresh and frozen processed pork, not canned or made into sausage, animal (non-poultry) slaughtering group, series no. PCU311611311611G from Labor, BLS (2006).
• U.S. wholesale price for sausage reflected by the U.S. PPI for sausage and similar products (not canned), animal (non-poultry) slaughtering group, series no. PCU311611311611J, from Labor, BLS (2006).

All modeled data are quarterly, not seasonally adjusted, and shown below to be non-stationary and integrated of order-1 \([I(1)]\). Due missing values in some series, and because of a desire to craft a sample completely within the liberalized U.S./Canadian trade regime established by the January, 1989 implementation of the Canada/U.S. Free Trade Agreement (CUSTA), our sample is 1989:01-2006:03. PUSPROC and PUSSAUS were the only U.S. price variables for downstream pork-based products located with adequate observations with which to implement this study. As repeatedly noted in recent time series studies of U.S. systems of commodity-related markets, quantities of the commodity-using downstream products are usually business proprietary and un-available publicly. This necessitated the modeling of the downstream markets in the same manner as the prior research as reduced form price equations [See Bessler and Akleman 1998; Lloyd et. al. (2006, 2001), and Goodwin, McKenzie, and Djunaidi (2003).]

Following Juselius and Toro (2005) and Juselius (1998; 2006, chapters 1-4), we examined the modeled data’s logged levels and differences to assess the data’s stationarity properties. Such examinations permitted the formulation of specification implications of these properties that will maximally harness or utilize their inherent stores of information in order to avoid potentially adverse econometric consequences of ignoring such information: compromised inference and spurious regressions (see also Granger and Newbold 1986, pp. 1-5, and Hendry 1986). In turn, using the statistically supported specification implications that utilize such nonstationarity-based information results in a statistically adequate VAR model with which the modeled system’s cointegration properties may be exploited.

A stationary and ergodic series has a constant and finite mean and variance, time-independent observations, and generates regression time-invariant estimates (Juselius 2006, chapters 1-3; Granger and Newbold 1986). Such data series frequently cycle and mean-revert. Due to space limitations, we do not include the plotted logged data and their differences. The following are the highlights of the specification implications needed to capture the information inherent in the modeled data’s nonstationary elements:
• Pork quantity exhibits trending and seasonal effects throughout the sample, with protracted subsample changes in plotted behavior, probably from a number of policy, institutional, and market events (hereafter, important market events). Specification implications include a linear trend, three quarterly centered seasonal binary variables or binaries, and a number of binaries to capture effects of important market events.

• PUSFARM exhibits a number of subsample episodes of altered plotted behavior, likely from important market events introduced below. PUSFARM differences suggest a number of potentially extraordinary and observation-specific transitory effects – hereafter denoted as outlier events/effects and addressed with outlier binaries. In particular, PUSFARM underwent an extraordinary and rapid decline of 43 percent in a single quarter to all-time 1998:04 lows because of supply-glut conditions and panicky herd slaughter and sell-offs, and then rebounded by an extraordinarily large and rapid 81 percent in the in the ensuing two quarters. Specification implications include a number of permanent shift binaries and a transitory binary for the pork market event just mentioned.

• Processed pork and sausage prices exhibit trending; subsample episodes of changing plotted behavior from important market events introduced later; and possible effects. Specifications include a time trend; permanent shift binaries; and transitory outlier binaries.

THE STATISTICAL MODEL: THE LEVELS VAR AND UNRESTRICTED VEC EQUIVALENT

To avoid confusing several different and sometimes overlapping definitions in the literature, we propose a number of definitions used throughout: (1) the unrestricted levels VAR denotes a VAR model in logged levels; (2) the unrestricted vector error correction or VEC denotes the algebraic equivalent of the unrestricted levels VAR in error correction form before the cointegration space is restricted for reduced rank or for assorted statistically supported hypothesis test conditions/relations; (3) the cointegrated VEC is the unrestricted VEC where the error correction space has been re-

3 This section draws heavily on the seminal articles by Johansen and Juselius (1990, 1992), and the recent book by Juselius (2006).
stricted for whatever reduced rank that evidence suggests: and (4) the *fully restricted cointegrated VEC* or cointegrated VAR* is the unrestricted VEC after the cointegration space’s restriction for reduced rank and for statistically supported conditions from the hypothesis tests. The “p” denotes the number (four) of endogenous variables; “p1” denotes the number of variables in the cointegration space (four endogenous and various deterministic variables introduced later); and “r” represents the cointegration space rank (and number of cointegrating relationships).

**The Levels VAR and Unrestricted VEC of the U.S. Pork-Based Markets.**

Sims (1980) and Bessler (1984) note that a VAR model posits each endogenous variable as a function of k lags of itself, and of each of the remaining endogenous variables. Tiao and Box’s (1978) lag search procedure suggested a one-lag structure. The above pork-related variables comprise the following unrestricted, 4-equation model:

\[
(1) \quad X(t) = a(1)*QUSPORK(t-1) + a(2)*PUSFARM(t-1) + a(3)*PUSPROC(t-1) \\
+ a(4)*PUSSAUS(t-1) + a(c)*CONSTANT + a(T)*TREND + \\
+ a(s)*SEASONALS + \epsilon(t)
\]

Above, the equation-specific notation is suppressed for simplicity since a VAR has each endogenous equation as a function of the same regressors, and \( X(t) = QUSPORK(t), PUSFARM(t), PUSPROC(t), \) and \( PUSSAUS(t). \) The asterisk denotes the multiplication operator. The \( a \)-coefficients are ordinary least squares (OLS) estimates with the parenthetical digit denoting the current or lagged values (t, t-1). Equation 1 also includes three quarterly seasonal variables and other permanent shift and outlier binaries not shown notationally.

Johansen and Juselius (1990, 1992) and Juselius (2006, pp. 59-63) demonstrate that a levels VAR of a lag order-k can be rewritten more compactly as an unrestricted VEC:

\[
(2a) \quad \Delta x(t) = \Gamma(1)\Delta x(t-1) + \ldots + \Gamma(k-1)\Delta x(t-k+1) + \Pi x(t-1) + \Phi D(t) + \epsilon(t)
\]

And for the case where \( k=1, \) equation 2a simplifies to equation 2b:

\[
(2b) \quad \Delta x(t) = \Pi x(t-1) + \Phi D(t) + \epsilon(t)
\]

The \( \epsilon(t) \) are white noise residuals. The \( x(t) \) and \( x(t-1) \) are p by 1 vectors of current and lagged endogenous variables in natural logarithms. The \( \Pi \) is a p by p long run er-
error correction term to account for endogenous variable levels. The $\Phi^*D(t)$ is a set of deterministic variables that includes three centered seasonal binaries, a trend, and other binary variables that are added to address issues as they arise as the analysis unfolds. It is well known that the one-order lag structure simplifies without equation 2a’s $\Gamma$-terms (Patterson 2000, p. 600). The $\Pi$ matrix is decomposed as follows:

$$\Pi = \alpha*\beta'$$

(3)

The $\alpha$ is a $p$ by $r$ matrix of adjustment speed coefficients and $\beta$ is a $p$ by $r$ vector of error correction coefficients, where here $p=4$ and $r$ is the reduced rank of equation 3 to be determined below.

The $\Pi = \alpha*\beta'$ is interchangeably denoted as the levels-based long run component, error correction term, or the cointegration space of the model. The $\Pi$-term retains the levels-based information, and includes non-differenced linear combinations of the individually I(1) endogenous variables; permanent shift binaries that reflect enduring event-specific effects (presented below); and a linear trend.

It is well known that the unrestricted VAR model framework developed by Sims (1980) and introduced early on to the U.S. hog market by Bessler (1984) is a reduced form one, where estimated relations reflect a mix of demand- and supply-side elements, typically without clear structural interpretations (Hamilton 1994, chapter 11). The extension to this framework by Johansen (1988) and Johansen and Juselius (1990, 1992) enables one to identify structural error correction relationships from what was once exclusively the reduced form approach of Sims (1980) and Bessler (1984) by separating-out the long run error correction term from the short run complement; by injecting economic theory and statistical inference through well-known Johansen-Juselius hypothesis test methods; and by applying reduced-rank estimation with statistically-supported restrictions from such hypothesis tests imposed.

In addition to a trend, we considered restricting non-differenced binary variables to the levels-based cointegration space to account for the implementations of the North America Free Trade agreement (NAFTA) in 1994 and the Uruguay Round in 1995. The starting point for the unrestricted VEC was equation 2b with no trend or binary variables. An adequately specified unrestricted VEC was ultimately achieved in a series of sequential estimations using packages by Estima (2006) and Dennis (2006). We estimate the model after having added the set of seasonal binaries, and then added a linear trend, and a number of quarter-specific outlier binaries – generally one vari-
able per each estimation. Variables were retained if diagnostic test values moved in patterns indicative of improved specification. Juselius (2006, chapters 4, 7, 9) recommends the following battery of diagnostic tests: (a) trace test correlation as an overall goodness of fit indicator, (b) likelihood ratio tests of autocorrelation, (c) Doornik-Hansen test for equation residual normality, and (d) indicators of skewness and kurtosis. The estimations were stopped when the array of diagnostic values failed to further improve with inclusion of additional variables.

The above sequential estimation procedure resulted in the inclusion of three quarterly centered seasonal binaries, the above-defined permanent shift binaries for U.S. implementations of the NAFTA and Uruguay Round agreements in non-differenced form within the error correction space, as well as in differenced form to account for short run effects. A trend was restricted to the cointegration space.

We also followed Juselius’ (2006, chapter 6) method of identifying and including extraordinarily influential effects of quarter-specific “outlier” events through specification of “outlier” binaries. When a potentially included outlier was identified with a “large” standardized residual, an appropriately specified variable was included in equation 2b in differenced form to capture the transitory effect, and retained if the battery of diagnostic variables moved favorably to suggest enhanced specification.\(^4\) Two outlier binaries were ultimately included.\(^5\)

\(^4\) We followed Juselius’ (2006, chapter 6) analysis outliers. An observation-specific event was judged as “extraordinary” if its standardized residual was 3.0 or more in absolute value. Such a rule for outliers was designed based on the 70-observation effective sample size using the Bonferroni criterion: \(\text{INVNORMAL}(1-1.025)^T\) where \(T=70\) and \(\text{INVNORMAL}\) is a function for the inverse of the normal distribution that returns the variable for the \(c\)-density function of a standard normal distribution (Estima 2006). The Bonferroni variate had a 3.4 absolute value. Having realized that there were some quarter-specific events with potentially extraordinary effects with absolute standardized residual values of about 3.0, we opted to choose a conservative Bonferroni absolute value criteria of 3.0 rather than 3.4. Observations with absolute standardized residual values of 3.0 or more were considered potential outliers, and we specified an appropriately defined variable for the relevant observation for the sequential estimation procedure.

\(^5\) The first outlier binary arose with QUSPORK that generated a standardized residual value of 3.0 for 1989:02 suggesting possible extraordinary adjustment effects of the CUSTA agreement, the onset of which served as the starting point for this study’s sample. The QUSPORK levels rose during 1989:02 and then receded, suggesting the following values for the transitory binary defined as DFDT89:02: 1.0 for 1989:02, -1.0 for 1989:03, and zero otherwise. The second outlier binary was defined when we noticed that PUSFARM levels plummeted extraordinarily to all-time lows: by about 43 percent in a single quarter to 31.2 due to conditions of over-supply ensued by panicky herd sell-offs and slaughter. PUSFARM then recovered about as extraordinarily as it had previously plummeted over the ensuing two quarters by having increased by about 81 percent to 56.47 in 199:02. The differenced binary variable DFSHK98 was defined as follows: 1.0 for 1998:04, 0.0 for
Table 1’s battery of diagnostic values for the levels VAR and its unrestricted cointegrated equivalent before and after efforts at specification improvement through inclusion of the statistically supported specification implications suggests clear benefits to Juselius’ (2006) recommended efforts. Such efforts bolstered the model’s ability to explain variation of the data, as reflected from the trace correlation, a goodness of fit indicator, having risen 235 percent from 0.173 to 0.58.

The Doornik-Hansen (D-H) test for system normality suggests that specification efforts on the initially non-normal system resulted in the ultimate achievement of an approximately normal system of residuals, as the system D-H value fell from 39.7 that strongly suggested non-normal residual behavior to 8.4 (p-value of 0.39) that strongly suggests normal behavior. Further D-H evidence at the individual equation level suggests that initially, two of the system’s four equations generated non-normal residual behavior, while specification efforts ultimately generated D-H values indicative that all four equations generated normally behaving residuals.

Table 1 also suggests that skewness and kurtosis values were within acceptable ranges for all equations after specification efforts. Table 1’s results suggest that we achieved a reasonably specified unrestricted VEC, which we can use to exploit any of the system’s cointegration properties as exist.

COINTEGRATION: CHOOSING AND IMPOSING REDUCED RANK ON THE ERROR CORRECTION SPACE

The four endogenous variables are shown below to be I(1), and so their differences are I(0). Cointegrated variables are driven by common trends, and stationary linear combinations called cointegrating vectors or CVs (Juselius 2006, p. 80). The $\Pi$-matrix in equations 2a, 2b, and 3 is a 4 by 4 matrix equal to the product of two p by r matrices: $\beta$ of error correction coefficient estimates that under cointegration combine into r<p stationary CVs of the four individually non-stationary pork-related endogenous variables, and $\alpha$ of adjustment speed coefficient estimates (beta and alpha estimates, respectively). And so the reduced rank, $\beta^\prime x(t)$ is I(0) even though x(t)’s four series are nonstationary.

Determination of cointegration rank is a three-tiered process (Juselius 2006, p. 139). First, one conducts trace tests of Johansen and Juselius (1990, 1992). Second, one ex-

1999:01, -1.0 for 1999:02, and zero otherwise. As noted by market experts and by remarkably rapid recovery of PUSFARM from its late-1998 decline, this episode was likely a transitory one (see also Labor, BLS 2007).
amines patterns of characteristic roots generated under relevant assumptions of reduced rank. And third, one examines statistical significance patterns of adjustment speed coefficient estimates for CVs potentially considered for the error correction space.

Table 2’s nested trace tests reject the first null hypothesis that \( r \) is zero, but fails to reject the second that \( r \) is one of less, suggesting that the reduced rank of \( \Pi \) is 1, and that there is one CV error-correcting the system. We follow Juselius’ (2006, p. 142) recommendation and consider other evidence when choosing reduced rank.

If \( r=1 \) is appropriate, there should be \( p-r \) characteristic unit roots in the companion matrix under \( r=1 \), and the fourth being be sub-unity. The first four characteristic roots of the companion matrix under \( r=1 \) suggest that \( r \) is indeed one: 1.0, 1.0, 1.0, 0.84.

If \( r \) is an appropriate choice, then some of the \( \alpha \)-estimates generated for the unrestricted VEC’s CVs in the \( \Pi \)-matrix should have absolute values equal to or above 2.6, with this absolute critical value generally higher than a Student-\( t \) critical value. If all of a CV’s alpha estimates are insignificant, Juselius (2006, p. 142) contends that the CV contributes little to the error correction process and should perhaps be excluded from the cointegration space. Because of space considerations, we do not report all of the unrestricted VEC’s \( \Pi \)-matrix estimates. The third CV’s generated no strongly statistically significant alpha estimates, suggesting that these two should not be included in the error correction space. The second only has a single alpha estimate for \( \Delta \text{PUSPROC} \), and with a \( t \)-value of -3.2. Finally, the first CV1 has strong patterns of alpha estimate significance, with three of the four being significant: \( \Delta \text{QUSPORT} \) with a \( t \)-value of 3.6, \( \Delta \text{PUSPROC} \) with a \( t \)-value of -4.5, and \( \Delta \text{PUSSAUS} \) with a \( t \)-value of -5.8. We conclude that the sharp drop in statistically significance \( \alpha \)-estimate patterns from the CV1 set to the CV2 set may suggest more support for a reduced rank of 1 rather than 2. We conclude that evidence from trace tests, patterns of characteristic root values from relevant companion matrices, and the unrestricted VEC’s patterns of alpha estimate significance suggest that reduced rank of \( \Pi \) is likely \( r=1 \).

Equation 4 or the cointegrated VEC is the single cointegrating relationship or CV that emerged after imposing rank \( r=1 \) and re-estimation of the adequately specified unrestricted VEC with the reduced rank estimator developed and applied by Johansen (1988) and Johansen and Juselius (1990, 1992), and normalized on QUSPORK. The parenthetical terms are \( t \)-values.
QUSPORK = \(-0.84\)PUSFARM + 0.43PUSPROC + 1.86PUSSAUS
\(-0.26\)URUGUAY + 0.21NAFTA + 0.002TREND
(4)

\(-5.7\) (0.93) (2.9)
\(-2.9\) (-2.1) (0.66)

**HYPOTHESIS TESTS AND INference ON THE ECONOMIC CONTENT OF THE COINTEGRATING RELATION CV1**

We begin with equation 4, the unrestricted CV, conduct a series of tests on \(\Pi = \alpha\beta'\), and then re-estimate the system with the statistically-supported restrictions imposed. We use the reduced estimator noted earlier and programmed in Dennis (2006). Hypothesis tests on the beta coefficients take the form:

\[\beta = H\varphi\]  

(5)

The \(\beta\) is a p1 by r=1 vector of \(\beta\)-coefficients on variables in the cointegration space; \(H\) a p1 by s design matrix, with “s” being the number of unrestricted or free beta coefficients; and \(\varphi\) is an s by r=1 matrix of the unrestricted beta coefficients. Johansen and Juselius’ (1990, 1992) well-known hypothesis test value is provided in equation for the case where r is 1.

\[-2\ln(Q) = T^*[(1-\lambda^*)(1-\lambda)] \]  

(6)

The asterisked (non-asterisked) eigenvalue is generated by the model estimated with (without) the tested restriction(s) imposed.

The \(\alpha\) or adjustment speed coefficients characterize the relative speeds of error-correcting adjustment with which endogenous variables respond to a given shock (Johansen and Juselius 1990, 1992). The null hypothesis or H(0) is:

\[H(0): \alpha = A\psi\]  

(7)

A is a p by s design matrix; s is the number of unrestricted coefficients in the r=1 column of the \(\alpha\) matrix; and \(\psi\) is the s by r=1 matrix of the non-restricted or “free” ad-

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6 This methods section closely follows procedures provided in Johansen and Juselius (1990, 1992) and Juselius and Toro (2005), and thoroughly detailed in Juselius (2006, chapters 10 and 11).
justment speed coefficients (Juselius 2004, chapter 11). Equation 6's test statistic also applies here. Hypothesis tests on the $\beta$ and $\alpha$-estimates follow.

**Hypothesis Tests on the Beta Estimates.**

There are two sets of hypothesis tests on the beta coefficients. The first set of 4 tests examines if each endogenous variable is stationary under the imposed rank of 1. And the second set is performed on individual $\beta$-estimates in equation 4 after any statistically supported stationarity test conditions are imposed.

We follow Juselius (1998; 2006, p. 297), Juselius and Franchi (2007), and Juselius and Toro (2005) and used a multivariate likelihood ratio test of each endogenous variable’s stationarity within a system setting and given the imposed rank. This test for stationarity utilizes equation 5 that is rewritten as equation 8:

$$\beta^c = [b, \varphi]$$

(8)

For each of the four tests of stationarity for each endogenous variable, the $\beta^c$ is the $p_1$ by $r$, here 7 by 1, beta matrix with one of the endogenous variables being tested for stationarity restricted to a unity value (Juselius 2006, p. 183). The $b$ vector is a $p_1$ by 1 (here 7 by 1) vector: there is a unity value corresponding to the variable being tested for stationarity, zeros for the other three non-tested endogenous variables, and unity for the three deterministic components restricted to the error correction space (URUGUAY, NAFTA, and TREND). Under $r=1$, equation 8 substantially simplifies since with a single CV, the $\varphi$, with dimensions $p_1$ by $r-1$, is the null vector with dimensions of $p_1$ by $r-1$ or 7 by zero. Evidence was sufficient to reject the four null hypotheses that each of the endogenous variables is stationary.\(^\text{14}\)

The second set of hypothesis tests on individual beta coefficients are now conducted directly on equation 4, insofar as we impose no statistically supported restrictions of

\(^{14}\) With three deterministic components retained and the imposed rank of $r=1$, then equation 8's test value is distributed under the null hypothesis of stationarity as a chi-squared variable with 3 degrees of freedom. Test values with parenthetical $p$-values are as follows, with the null of stationarity rejected for $p$-values less than 0.05: 18.3 ($p=0.000$) for QUSPORK; 21.7 ($p=0.000$) for PUSFARM; 18.6 ($p=0.000$) for PUSPROC); and 20.4 ($p=0.000$) for PUSSAUS.
stationarity as justified by the first set of hypothesis tests. This second set of hypothesis tests arise from the application of economic and statistical theory, market knowledge, prior research, and/or suggestions implied by coefficient estimates and are tested using equations 5 and 6. If statistically supported at the 5-percent level of statistical significance (hereafter, the 5-percent level), the restriction is retained, and the Johansen-Juselius reduced rank estimator is then used to re-estimate the cointegration space with the statistically supported restriction imposed. The second set begins directly on equation 4, and this second set consists of two hypothesis tests:

- **Given equation 4’s negligibly valued and highly insignificant trend coefficient** (t-value of -0.66), we tested that $\beta(\text{trend}) = 0$. The chi-squared test value of 0.019 (1 degree of freedom) strongly accepted the restriction insofar as the p-value of 0.90 was far above the 0.05 decision rule. We re-estimated with the zero restriction on the beta coefficient for TREND.

- **This subsequent estimation generated a set of results that included an insignificant t-value on processed pork price**, and we tested the hypothesis that $\beta(\text{PUSPROC}) = 0$. With a chi-square test value of 0.33 (2 degrees of freedom) and a p-value of 0.85, the restriction was strongly accepted, re-imposed, and then the equation was re-estimated final time with the reduced rank estimator. Equation 9 emerged as the finally restricted cointegrating relation or vector (CV):

\[
QUSPORK = -0.68*PUSFARM + 1.84*PUSSAUS - 0.22*URUGUAY + 0.19*NAFTA
\]

\[
(9) \quad (-7.4) \quad (-7.1) \quad (-3.1) \quad (2.5)
\]

The $\alpha$-estimates (adjustment speed coefficient estimates) and the parenthetical t-values were calculated as follows: $\alpha(QUSPORK) = -0.81 \ (t = -3.3)$, $\alpha(PUSFARM) = 0.27 \ (t = 2.0)$; $\alpha(\text{PUSPROC}) = 0.19 \ (t = 3.7)$; $\alpha(\text{PUSSAUS}) = 0.14 \ (t = 5.9)$.

Equation 9 appears to be a U.S. demand for pork as an input. Own price has a strongly significant negative coefficient, that is translated as an own-price elasticity of demand of -0.68. This estimate falls within the range of alternative price elasticities of U.S. pork demand of from -0.51 to -0.84 reported by Wang and Bessler (2003, pp. 510-511) and approaches the lower limit of the range of -0.8 to -1.2 reported by Eales and Unnevehr (1993, p. 264). As well, normalizing on pork quantity conceptually places it in positive form above the equilibrium on the relation’s left side, whereby
the $\alpha$-estimate is a significant -0.81 (t-value of -3.3), and suggestive of a demand’s downward adjustment toward the long run equilibrium or “attractor set.”

The demand for pork appears positively sensitive to output prices of such downstream pork-based products, as proxied by PUSSAUS. The coefficient of +1.8 suggests an elasticity relationship above unity and may, at first glance, appear rather elastic. However, the modeled QUSPORK series, a commodity quantity, experiences a degree of variability far in excess of that of PUSSAUS, a U.S. producer price index for a manufactured meat product. Having placed all data in natural logarithms, the standard error is interpreted as a proportional change, and as a percentage change when multiplied by 100. For our sample, QUSPORK’s standard error is 21 percent, far exceeds the 7-percent standard error of PUSSAUS, and may thereby give rise to the rather elastic PUSSAUS beta estimate in equation 9.

As well, the above results suggest that the Uruguay Round and NAFTA (and concurrent events) collectively augmented QUSPORK by about 7 percent: the Uruguay Round had a modest positive effect of about 25 percent; NAFTA had a negative effect of about 18 percent; and had a net positive effect on such demand. This net positive effect is likely explained by McMahon’s (1998) finding that U.S. implementation of both agreements likely augmented U.S. pork exports by more than imports declined – both of which are included as components in our QUSPORK definition. More specifically, Morrison (1996) noted that during 1990-1995, a period inclusive of the U.S. Uruguay Round and NAFTA implementations, the U.S. transformed from a major pork importer to a major pork exporter, as such U.S. pork-importing trading partners as Korea and Japan rendered concessions under both agreements.

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7 We used Halvorsen and Palmquist’s (1980) well-known method of interpretation of binary variable coefficient estimates when regressions are implemented with data converted to natural logarithms. One takes “$e$,” the base of the natural logarithm and raises it to the power of the value of the coefficient estimates, subtracts 1.0, and multiplies the result by 100. What results is an average percentage change effect on the dependent variable of the event (and collectively of other concurrent events) for which the binary variable was defined. Clearly, the Halvorsen-Palmquist calculation has a typical binary variable limitation of imprecision, in that the percent change effect cannot be attributed solely to the event for which the binary was defined (Uruguay Round implementation for URUGUAY in January 1995 or NAFTA’s January 1994 implementation), but also to other influential concurrent events.

8 We acknowledge that the negative impact on QUSPORK from NAFTA in equation 9 is not easily explained. Perhaps because of the close coincidence of the two U.S. implementations – January 1994 for NAFTA and January 1995 for the Uruguay Round – the two variables are collinear, and one emerges as wrongly signed. Nonetheless, the two variables do generate a collectively positive effect on QUSPORK, presumably because of reasons for the McMahon (1998) finding of positive effects of the agreements on U.S. pork exports.
The above $\beta$ and $\alpha$ estimates above in equation 9 suggest that the quantity demanded suggests that farm price for pork is weakly endogenous to the system, given its solidly significant own-price coefficient estimate with a t-value of -7, while its $\alpha$-estimate is insignificant with a t-value of 2.0. Juselius (2006, pp. 193-194) suggests that in such cases, own-price’s significant $\beta$-estimate suggests an influence on demanded pork quantity and the other endogenous variables, while the insignificant adjustment speed coefficient estimate suggests no reaction or adjustment price to the other endogenous variables’ error-correction behavior in equation 9.

SUMMARY AND CONCLUSIONS

There have been few applications of the cointegrated VAR modeling methods applied to capture the upstream/downstream parameters and relationships that characterize quarterly U.S. pork-based markets. And interest in such estimates and results has been increasing given a number of food scares mentioned above, and given a rising awareness by researchers, policy makers, agribusiness agents, farmers and consumers of the policy-relevant value of such empirical estimates and results. We applied such methods, and some of the more refined procedures outlined in Juselius (1998; 2006, chapters 1-8) and Juselius and Toro (2005), to the U.S market for pork at the farm gate, and to wholesale markets for fresh/frozen pork meat and for sausage downstream. The resulting model and error correction parameter estimates achieved notable statistical strength as reflected by table 1 and equation 9.

Evidence suggested that the three U.S. markets represented by the four modeled endogenous variables are bound together by one stationary long run cointegrating relationship that appears to be a U.S. demand for pork as a productive input. There appears to be a long run own-price elasticity for U.S. pork of -0.68 that is in line with literature estimates. Evidence suggests that U.S. pork demand is sensitive to downstream conditions in markets that utilize pork as a productive input. In line with Wang and Bessler’s (2006, p. 92-93) findings, price appears predetermined relative to quantity, suggesting that a quantity-dependent, rather than price-dependent, demand function is a viable specification for the U.S. pork market. Collectively, NAFTA and the Uruguay Round implementations had a modest positive effect on U.S. pork demand, perhaps as U.S. exports rose more than imports decreased after implementation of the Uruguay Round and NAFTA agreements.
References


### Table 1  Mis-specification Tests for the Unrestricted VEC: Before and After Specification Efforts

<table>
<thead>
<tr>
<th>Test and/or equation</th>
<th>Null hypothesis and/or test explanation</th>
<th>Prior efforts at specification adequacy</th>
<th>After efforts at specification adequacy</th>
</tr>
</thead>
<tbody>
<tr>
<td>Trace correlation</td>
<td>system-wide goodness of fit: large proportion desirable</td>
<td>0.173</td>
<td>0.58</td>
</tr>
<tr>
<td>ARCH tests for heteroscedasticity (lags 1, 4)</td>
<td>Ho: no heteroscedasticity by 1st, 4th lag for system. Reject with p-values less 0.05</td>
<td>lag 1: 135 (p=0.01)</td>
<td>lag 1: 97.9 (p=0.54)</td>
</tr>
<tr>
<td>Doornik-Hansen test, system-wide normality</td>
<td>Ho: modeled system behaves normally. Reject for p-values below 0.05.</td>
<td>39.7 (p=0.01)</td>
<td>8.3 (p=0.63)</td>
</tr>
<tr>
<td>Doornik-Hansen test for normal residuals (univariate)</td>
<td>Ho: equation residuals are normal. Reject for values above 9.2 critical value</td>
<td></td>
<td></td>
</tr>
<tr>
<td>∆QUSPORK</td>
<td>10.6</td>
<td>6.4 (*)</td>
<td></td>
</tr>
<tr>
<td>∆PUSFARM</td>
<td>16.4</td>
<td>3.7 (*)</td>
<td></td>
</tr>
<tr>
<td>∆PUSPROC</td>
<td>2.3</td>
<td>0.08</td>
<td></td>
</tr>
<tr>
<td>∆PUSSAUS</td>
<td>0.82</td>
<td>0.08</td>
<td></td>
</tr>
<tr>
<td>Skewness (kurtosis)</td>
<td>skewness: ideal is zero; “small” absolute value acceptable kurtosis: ideal is 3.0; acceptable is 2.5-4.0.</td>
<td>0.06 (2.5)</td>
<td>0.4 (2.3)</td>
</tr>
<tr>
<td>∆QUSPORK</td>
<td>-1.3 (7.0)</td>
<td>-0.5 (3.5)</td>
<td></td>
</tr>
<tr>
<td>∆PUSFARM</td>
<td>-0.42 (3.0)</td>
<td>-0.07 (2.7)</td>
<td></td>
</tr>
<tr>
<td>∆PUSPROC</td>
<td>-0.13 (3.1)</td>
<td>0.06 (2.80)</td>
<td></td>
</tr>
<tr>
<td>∆PUSSAUS</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes.– An asterisk (*) denotes a favorable movement (decline) in the relevant test/diagnostic value into the range of statistical normality.

### Table 2  Trace test statistics and related information for nested tests for rank determination.

<table>
<thead>
<tr>
<th>Null Hypothesis</th>
<th>Trace Value</th>
<th>95% Fractile (critical value)</th>
<th>Result</th>
</tr>
</thead>
<tbody>
<tr>
<td>rank or r ≤ 0</td>
<td>78.3</td>
<td>67.3</td>
<td>Reject null that rank is zero</td>
</tr>
<tr>
<td>rank or r ≤ 1</td>
<td>45.9</td>
<td>46.4</td>
<td>Fail to reject null that rank r ≤ 1</td>
</tr>
<tr>
<td>rank or r ≤ 2</td>
<td>23.4</td>
<td>29.3</td>
<td>Reject null that rank r ≤ 2</td>
</tr>
<tr>
<td>rank or r ≤ 3</td>
<td>7.4</td>
<td>16.0</td>
<td>Reject null that rank r ≤ 3</td>
</tr>
</tbody>
</table>

Notes.–As recommended by Juselius (2004, p. 171), CATS2-generated fractiles are increased by 2*1.8 or 3.6 to account for the 2 permanent shift binary variables restricted to lie in the cointegration relations. Trace values are corrected with Bartlett’s small sample adjustment programmed by Dennis (2006).
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Institute of Food and Resource Economics

<table>
<thead>
<tr>
<th>Date</th>
<th>Authors</th>
<th>Title</th>
</tr>
</thead>
<tbody>
<tr>
<td>09/07</td>
<td>Ronald Babula, Mogens Lund</td>
<td>Exploiting the Cointegration Properties of US Pork related Markets: The Emergence of a U.S. Demand for Pork as an Input</td>
</tr>
<tr>
<td>08/07</td>
<td>Jørgen Dejgård Jensen, Anja Skadkær Møller</td>
<td>Vertical price transmission in the Danish food marketing chain</td>
</tr>
<tr>
<td>07/07</td>
<td>Derek Baker, Karen Hamann</td>
<td>Innovation and the policy environment Findings from a workshop with meat industry firms in Skive</td>
</tr>
<tr>
<td>06/07</td>
<td>Derek Baker, Jens Abildtrup, Anders Hedetoft, René Kusier</td>
<td>Role of regional and rural development policy in supporting small-scale agribusiness in remote areas</td>
</tr>
<tr>
<td>05/07</td>
<td>Jørgen Dejgård Jensen</td>
<td>Analyse af tre forskellige scenarier for afgiftsændringer på fødeværer</td>
</tr>
<tr>
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<td>Hans Grinsted Jensen, Kenneth Baltzer, Ronald A. Babula, Søren E. Frandsen</td>
<td>The Economy-Wide Impact of Multilateral NAMA Tariff Reductions: A Global and Danish Perspective</td>
</tr>
<tr>
<td>03/07</td>
<td>Svend Rasmussen</td>
<td>Optimising Production using the State-Contingent Approach versus the EV Approach</td>
</tr>
<tr>
<td>02/07</td>
<td>Kenneth Baltzer, Søren E. Frandsen, Hans G. Jensen</td>
<td>European Free Trade Areas as an alternative to Doha - Impacts of US, Russian and Chinese FTAs</td>
</tr>
<tr>
<td>01/07</td>
<td>Lill Andersen, Ronald A. Babula, Helene Hartmann, Martin M. Rasmussen</td>
<td>A Vector Autoregression Model of Danish Markets for Pork, Chicken, and Beef</td>
</tr>
<tr>
<td>Date</td>
<td>Authors</td>
<td>Title</td>
</tr>
<tr>
<td>---------</td>
<td>----------------------------------------------</td>
<td>-----------------------------------------------------------------------</td>
</tr>
<tr>
<td>11/06</td>
<td>Lars Otto</td>
<td>GRO modellen: Grise, Risiko og Økonomi. Datagrundlag</td>
</tr>
<tr>
<td>10/06</td>
<td>Lars Otto</td>
<td>GRO modellen: Grise, Risiko og Økonomi. Teoretiske grundlag</td>
</tr>
<tr>
<td>09/06</td>
<td>Johannes Sauer, Arisbe Mendoza-Escalante</td>
<td>Schultz’s Hypothesis Revisited – Small Scale Joint - Production in the Eastern Amazon</td>
</tr>
<tr>
<td>08/06</td>
<td>Johannes Sauer, Jesper Graversen, Tim Park, Solange Sotelo, Niels Tvedegaard</td>
<td>Recent Productivity Developments and Technical Change in Danish Organic Farming – Stagnation?</td>
</tr>
<tr>
<td>07/06</td>
<td>Johannes Sauer</td>
<td>Prices and Species Diversity – Stochastic Modelling of Environmental Efficiency</td>
</tr>
<tr>
<td>06/06</td>
<td>Jacob Ladenburg, Søren Bøye Olsen</td>
<td>Starting Point Anchoring Effects in Choice Experiments</td>
</tr>
<tr>
<td>05/06</td>
<td>Svend Rasmussen</td>
<td>Optimizing Production under Uncertainty. Generalization of the State- Contingent Approach and Comparison with the EV Model</td>
</tr>
<tr>
<td>04/06</td>
<td>Red. Johannes Christensen</td>
<td>Fremtidens biogasfællesanlæg. Nye anlægskoncepter og økonomisk potentiale</td>
</tr>
</tbody>
</table>