A Dynamic Model of U.S. Sugar-Related Markets: A Cointegrated Vector Autoregression Approach

Ronald A. Babula, Douglas Newman, and Robert A. Rogowsky

The methods of the cointegrated vector autoregression (VAR) model are applied to monthly U.S. markets for sugar and for sugar-using markets for confectionary, soft drink, and bakery products. Primarily a methods paper, we apply Johansen and Juselius' advanced procedures to these markets for perhaps the first time, with focus on achievement of a statistically adequate model through analysis of a battery of advanced statistical diagnostic tests and on exploitation of the system's cointegration properties through rank restrictions, statistically supported hypothesis test restrictions, and inference. The VEC model results illuminate the estimates of crucial policy-relevant market parameters that drive these markets, as well as the dynamic nature of the relationships linking these sugar-based markets.

Because the United States is a "large economic country," changes in market conditions, domestic/international policies, decisions from trade-remedy cases, trading-rule changes from free-trade agreements or FTAs, and other economic and institutional changes have effects with often opposing benefits on upstream U.S. farmer-producers of a commodity and on downstream product manufacturers that use the commodity. One only need consult the recent U.S. antidumping and countervailing duty cases on certain imports of Canadian wheat as an example (U.S. International Trade Commission 2003). The profession consequently needs models that capture not only the U.S. farmer costs of, say, increased U.S. import-market access of a farm commodity but import-induced benefits for downstream manufacturer-users of the commodity as well. We contend that building an econometric model to capture such a balanced array of often opposing upstream and downstream effects is a tailor-fit task for Johansen and Juselius' (1990, 1992) methods of the cointegrated vector autoregression, or VAR, model.

A varied array of events have recently influenced or even disrupted U.S. markets for raw sugar and sugar-related confectionary, bakery, and soft drink markets:

- The U.S. Department of Agriculture's Economic Research Service (USDA, ERS 2006, pp. 5–6) notes that U.S. 2005–2006 raw sugar supply was disrupted by three recent hurricanes that led to sugar cane production declines of 20 percent in Louisiana from Hurricanes Katrina and Rita and 12 percent in Florida from Hurricane Wilma.

- Because of such hurricane-induced costs on both U.S. sugar producers and downstream sugar-product manufacturers, the USDA acted in December 2005 to raise the in-quota access of the U.S. sugar tariff rate quota (TRQ) by 38 percent, or 450,000 short tons (ST)—300,000 ST for raw sugar and 150,000 ST for refined sugar—to about 1.65 million ST in order to mitigate supply shortfalls (USDA ERS 2006, pp. 25–31). This increased TRQ access likely benefits sugar-using agents more than sugar producers (USDA ERS 2006).

- There is uncertainty over future increased foreign access to the U.S. sugar-import market for Thailand and a number of Central American

Babula is Director of Research and associate professor, Division of International Economics and Policy, Food and Resource Economics Institute, The Royal Veterinary and Agricultural University, Denmark. Newman is international trade analyst in the Office of Industries and Rogowsky is Director of the Office of Operations, United States International Trade Commission.

The authors are particularly grateful to the following econometricians for their thorough reviews, their substantial advice in econometric procedures and interpretation of results, and for help in some of this study's more challenging software programming: David A. Bessler, Texas A&M University; Heiko Hesse, Oxford University; and Jonas Stulz, Swiss National Bank. We are particularly indebted to the insightful comments of an anonymous reviewer.

nations with which FTAs and/or FTA implementation legislation are pending (see USDA ERS 2006, pp. 2–3; U.S.-Thailand Free Trade Agreement Business Coalition 2004). Any increased access tends to favour the interests of U.S. sugar-using manufacturers and may work against the interests of U.S. sugar producers.

- A disparate non-systematic array of U.S. sugar policies since the late 1990s are examined and shown to be mutually conflicting in terms of price-support and stock-management goals, with no clear indications of market effects for U.S. sugar, confectionary, and soft drink markets.

- During 2001 and 2002 there was an array of confectionary production input price increases, such as sustained 2001 cocoa price rises from the Ivory Coast civil war, that have likely had substantial impacts on U.S. sugar-producing and -using markets. Such impacts are not yet chronicled by the literature, and can be illuminated with a cointegrated VAR of such markets.

We propose a monthly cointegrated VAR model of U.S. sugar, confectionary, bakery, and soft drink markets that provides a balanced array of estimated or implied upstream and downstream market effects of such events. Such estimated effects arise from the model's estimates of market-propelling elasticities, price-transmission parameters, dynamic patterns of market responses and interactions, and empirical estimates of actually observed events (hereafter, the model's empirical/dynamic results). While some of the events occurred after our estimation period, our results are generalized and sufficiently long-run to provide guidance in discerning effects of some post-sample events.

We located no studies that apply methods of the cointegrated VAR model developed by Johansen (1988) and Johansen and Juselius (1990, 1992) to U.S. sugar, confectionary, bakery, and soft drink markets. We did locate three price-only cointegration-model applications to U.S. sugar and high fructose corn syrup (HFCS) prices, although the rather specific sugar/HFCS focus of these studies rendered them of indirect relevance to our multi-market empirical estimates and findings: Williams and Bessler (1997) and Moss and Schmitz (2002, 2004). Consequently, we address these articles when relevant to our specification and results, and do not fully review them here.

Our primary purpose is consequently methods-based: to address this lack of empirical inquiry on U.S. sugar, confectionary, bakery, and soft drink markets by applying Johansen and Juselius' (1990, 1992) cointegrated VAR model methods to these U.S. sugar-related markets.[1] The principal hypothesis to be tested is whether, as theory generally suggests, U.S. market-clearing prices and quantities of sugar and sugar-using confectionary, soft drink, and bakery products form a cointegrated econometric system. In the process, we interlace these well-known procedures with a number of new tests and methods for perhaps the first time in the agricultural economics literature. Of equal importance, our second purpose is to take the model's empirical and dynamic results and illuminate how the markets interact to such important economic/trade/institutional events summarized above and examined below.

Seven sections follow. The first discusses time-series econometrics and the cointegrated VAR as a way of empirically examining five chosen monthly U.S. sugar and sugar-based markets. Data and data sources are also presented. A second section analyses data-behavior patterns to obtain model-specification implications. The third section summarizes efforts expended on achieving a statistically adequate levels-VAR model (and its algebraic equivalent, an unrestricted VEC model). This unrestricted VEC is not yet restricted for cointegration space rank or for restrictions emerging from theory-based statistical hypothesis tests discussed later. We provide a rigorous analysis of the model's statistical adequacy based on results from a battery of diagnostic mis-specification tests suggested in Juselius (2004, pp. 72–82). Fourth,

[1] We are very grateful to Dr. Katarina Juselius of the Institute of Economics, University of Copenhagen, for her generous allocation of time and attention to this study, which began as a class project for a course taught by Drs. Katarina Juselius, Soren Johansen, Heino Bohn Nielsen, and Anders Rahbek in Copenhagen, August 2–22, 2004. An author was fortunate enough to have taken the course, and received the benefits of the direct attention and skilful instruction of Dr. Juselius and the other faculty members in all phases of specification, estimation, and result interpretation of the cointegrated VAR model of U.S. sugar-related markets in this paper. Needless to say, all errors and shortcomings of this paper are solely our own.

evidence from Johansen and Juselius' (1990, 1992) well-known trace tests and from other sources is used to determine the rank of the cointegration space and the number of long-run cointegrating relationships. The cointegration space is then restricted for this reduced rank. Fifth, Johansen and Juselius' well-known hypothesis-test procedures are applied to the rank-restricted cointegration space to illuminate the long-run relationships tying together the upstream and downstream sugar-based markets. A sixth section provides an economic interpretation of the cointegrating relations that are fully restricted for rank and for statistically-supported restrictions that emerged from the hypothesis tests. A summary and set of conclusions then follow.

Time-Series Econometric Considerations, Modeled Markets, and Data Sources

Engle and Granger (1987) established that economic time series often fail to meet the conditions of weak stationarity—that is, conditions of stationarity and ergodicity summarized by Granger and Newbold (1986, pp. 1–5). And while data series are often individually nonstationary, they can form vectors with stationary linear combinations, such that the vector of interrelated series are "cointegrated" and move in tandem as an error-correcting system (Johansen 1990, 1992). As shown later, readers should note that all endogenous variables are nonstationary: integrated of order-1, or I(1), with the first differences being stationary, or integrated of order-zero or I(0).

A search of monthly data resources on U.S. sugar-based variables provided the six variables below for the January 1992–March 2004 period. Data are unavailable prior to January 1992. Given the number of endogenous variables and deterministic variables to be added from later analysis, there are potential problems with small samples. The following six variables (denoted by the parenthetical labels) are proposed as an unrestricted VAR and ultimately a cointegrated VEC model:

1. The market-clearing quantity of available sugar in the United States (QSUGAR), which includes all shipments, ending stocks, and imports of sugar in raw-sugar equivalents from the United States Department of Agriculture, Economic Research Service (USDA ERS 2004a, 2004b).

2. The market-clearing price of raw sugar (PSUGAR) published by the USDA ERS (2004).

3. The producer price index (PPI) or price of U.S. chocolate-based confectionary products (PCHOC), PPI series number PCU311320311320, from the U.S. Department of Labor, Bureau of Labor Statistics (2004).

4. The price of non-chocolate confectionary products (PNOCHOC), PPI series number PCU3113403113401 from USDL BLS (2004).

5. The price of U.S. soft drink products (PSOFT), PPI series number PCU312111312111, from USDL BLS (2004).

6. The U.S. price of commercial bakery products (PBAKERY), PPI series number PCU311812311812, from USDL BLS (2004).

In rationalizing the above vector of endogenous sugar-based variables, we were primarily guided by theoretical and practical considerations established by five time-series econometric studies of U.S. systems of markets commonly related through a farm commodity: the seminal Bessler (1984) work on U.S. pork-related markets; Goodwin, McKenzie, and Djunaidi (2003) for chicken products; Babula, Bessler, and Payne (2004) and Babula and Rich (2001) for wheat-related products; and Babula et. al. (2004) for soy-based products. Our specification was also guided by the three cointegrated analyses on U.S. sugar/HFCS price issues. First, we followed the six latter studies, particularly Bessler (1984), and selected an endogenous vector based loosely on neoclassical microeconomic theory of the firm and consumer. Our sugar-based markets are clearly related, and are shown below to be a cointegrated system, because agents from these markets participate and compete in common-input markets (particularly raw sugar) on the production side, and the markets are related through consumer-preference forces (product substitutability and/or complementarity) and through consumer allocation of scarce income on the retail-demand side. Bessler (1984) established the benefits of using theory loosely in delineating modeled VAR systems to permit the regularities in the data to reveal themselves. Second, previous work clearly establishes that monthly and quarterly quantity data from downstream commodity-based products are typically considered business proprietary information, beyond the public domain, and unavailable. So following prior work, our downstream markets are limited to reduced-form price equations, without modeled quantities. Third, we excluded the HFCS market, since prior cointegration modeling research on monthly U.S. sugar and HFCS prices suggested that the two U.S. markets have for some time ceased to be cointegrated and constitute separate and non-competing sweetener markets given their samples' grids of U.S. policies (Moss and Schmitz 2002, 2004). Fourth, given our already sizeable endogenous vector and large array of ultimately included deterministic variables introduced later, we resisted inclusion of other downstream value-added markets with weaker or perhaps debated links through raw sugar use. We justified our decision for a parsimonious model on Juselius and Toro's (2005, p. 511) recent and important observation on cointegration methods. They noted that limiting endogenous vector size is often justified because cointegrated VAR procedures are very powerful with small systems, are "almost unmanageable" with large systems, and generate cointegration properties in limited variables sets that remain invariant to future information-set expansions through new data or enhanced sample size. [2]

Analysis of the Sugar-Based Time-Series Data

Following Juselius (2004) and Bessler (1984), all data are in natural logarithms and differences of logged levels. Figures 1–6 provide the plotted data: logged levels in the upper panels and differences in the bottom panels. A weakly stationary data series has a constant and finite mean and variance, time-independent observations, and generates regression-coefficient estimates that are time-invariant, i.e. not subject to statistical structural change. (Juselius 2004, chapters 3 and 4). Weakly stationary data typically cycle and mean-revert frequently. Failure to utilize information inherent in data-nonstationarity elements may encounter serious and well-known problems with compromised inference, spurious regressions, and in some cases, biased estimates (Granger and Newbold 1986, pp. 1–5; Hendry 1986, pp. 201–204). A number of VAR model specification implications arise from examining the data's nonstationary information sources in order to obtain valid regression estimates.

QSUGAR, or the market-clearing sugar quantity, is saddled with obvious seasonal effects in Figure 1, which likely mask an upward trend that is only minimally evident from the plotted levels. Figure 1 suggests the need for centered seasonal dummy or binary variables and possibly a linear trend.

The price of raw sugar (PRAW), the own-price of a highly regulated and policy-driven market, reflects several serious issues in Figure 2. The following four major changes (described in three bullets) in U.S. sugar policy clearly destabilized the U.S. sugar market, as particularly evident from Figure 2's plotted levels and differences in PRAW during the sample period's latter half[3]:

- Two separate periods of domestic marketing allotments during October 1990–September 1996 and October 2000–present. These allotment programs were designed to restrict U.S. raw sugar supply, and presumably augmented U.S. raw sugar price.

- Two separate budget years of payment-in-kind (PIK) during October 2001–September 2003 when farmers were paid program benefits in-kind with raw sugar, which they sold on the open market. The PIK program increased marketed QSUGAR and dampened PRAW.

- The May 2002 implementation of a U.S. Government import-quota swap program that restricted imported raw sugar supplies in response to increased domestic production. This

[2] In modelling the Spanish monetary-policy mechanisms before and after entry into the European Monetary System, Juselius and Toro (2005, p. 511) opted not to include variables on the exchange-rate transmission channel of monetary policy and left including such channels to future research. They justified their decision by appealing to the above points. They noted that cointegration results found within a limited set of variables will carry-over to extended analyses with expanded information sets, with appreciable differences emerging only in the short-run dynamics.

[3] Price stability shown in Figure 2 was attributed to U.S. sugar-policy changes by the sugar market analyst with the Agriculture and Fisheries Division, Office of Industries, U.S. International Trade Commission in emails received by one of the authors on Aug. 18 and 20, 2004.

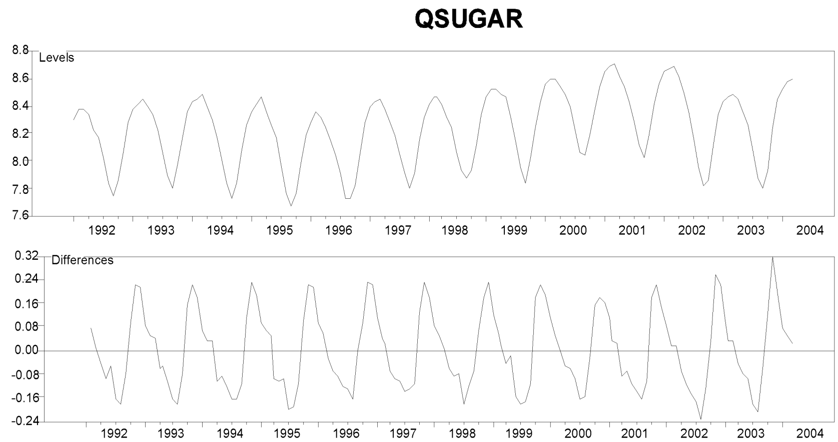

Figure 1. Plots of Logged Levels and Differences: QSUGAR.

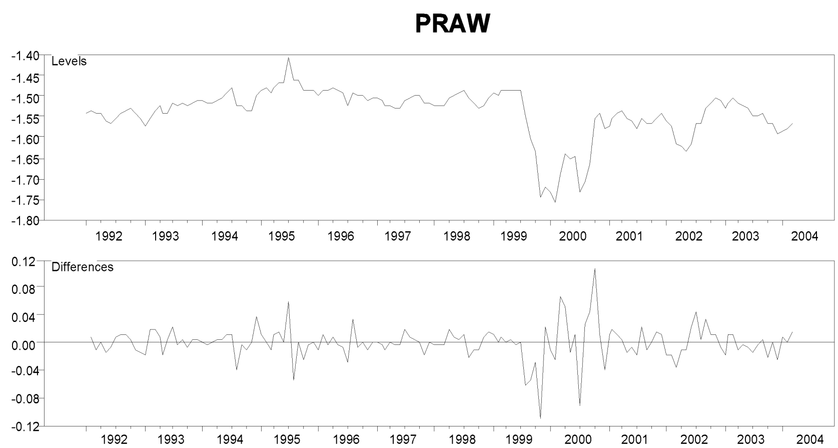

Figure 2. Plots of Logged Levels and Differences: PRAW.

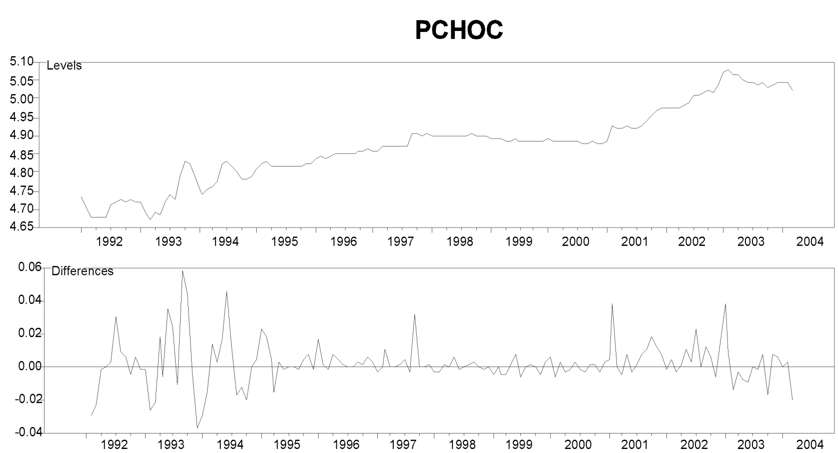

Figure 3. Plots of Logged Levels and Differences: PCHOC.

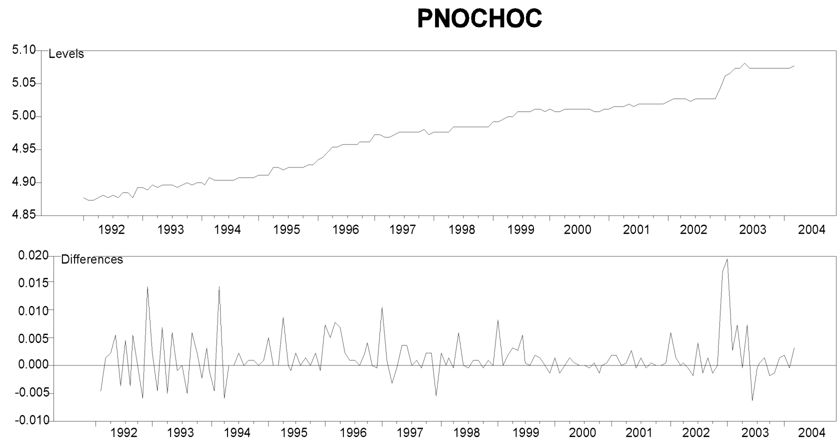

Figure 4. Plots of Logged Levels and Differences: PNOCHOC.

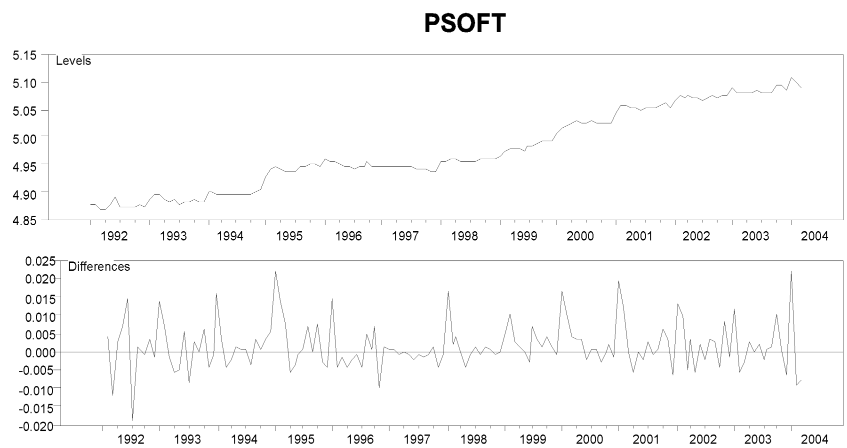

Figure 5. Plots of Logged Levels and Differences: PSOFT.

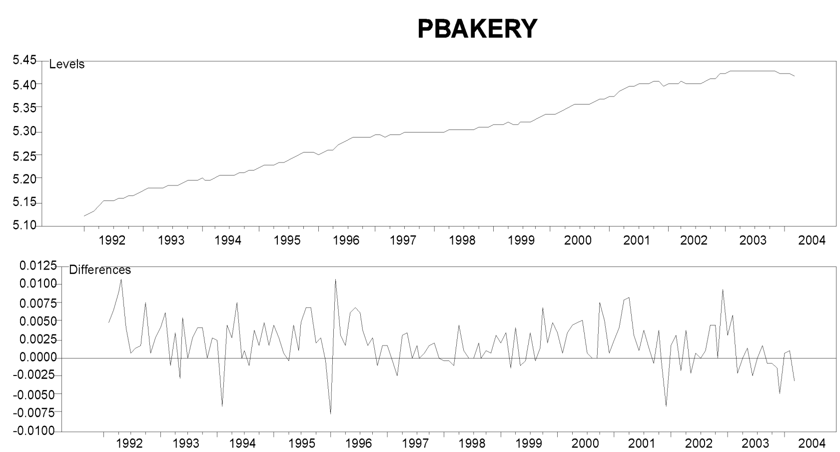

Figure 6. Plots of Logged Levels and Differences: PBAKERY.

program tended to augment PRAW through restricted raw sugar supplies in the U.S. market.

PRAW differences also suggest that these policy changes may have elicited a rise in the level of variation (i.e., heteroscedasticity or ARCH effects) toward the sample's end, as well as a number of observation-specific non-normal "outlier" effects throughout the sample. Specification considerations include a permanent shift binary after mid-1999 to account for the collective market effects of these four policies,[4] observation-specific binaries to account for effects of extraordinary events, ARCH effects, and a trend.

The prices of chocolate-based and non-chocolate confectionary products (PCHOC, PNOCHOC) in Figures 3 and 4 display several behavioral aspects inconsistent with weak stationarity: persistent trending, little or no cycling of levels data, and no mean-reversive behavior. Two dramatic permanent shifts appear, with approximate resumption of the pre-shift slopes: in early 2001 for PCHOC and in late-2002 for PNOCHOC. These slope shifts are likely caused by two events: an extraordinary and sustained run-up in cocoa prices in 2001 from the Ivory Coast civil war, and a pronounced late-2002 rise in non-cocoa confectionary input costs.[5] Specification considerations include two permanent shift binaries to account for the 2001 rise in cocoa price and the late-2002 rise in the prices of non-cocoa confectionary inputs, a series of observation-specific binaries to account for intermittent outlier and ARCH effects throughout the sample (see plotted differences), and a linear trend.

Plots and differences of the soft drink and commercial bakery price series in Figures 5 and 6 reflect clear trending, a number of outlier events, and virtually no cycling or mean-reversion. Specification considerations include a linear trend and various outlier binaries.

The Statistical Model: The Unrestricted VAR and VEC Equivalent[6]

Throughout this section, a number of terms are used. The *unrestricted levels* VAR denotes a VAR model in logged levels. An *unrestricted VEC* denotes the algebraic equivalent of the unrestricted levels VAR in error correction form, before the levels component (cointegration space) is restricted for reduced rank and/or for statistically supported restrictions. The *cointegrated VEC* is the unrestricted VEC where the cointegration space has been restricted for reduced rank. The *fully restricted cointegrated VEC* is the unrestricted VEC after the cointegration space's restriction for reduced rank and for the statistically supported restrictions on cointegration space coefficients from the hypothesis tests. The "p" denotes the number of endogenous variables (here, 6); "p1" denotes the number of variables in the cointegration space (endogenous and various deterministic variables introduced later); and "r" represents the reduced rank of the cointegration space (and the number of cointegrating relations).

We now specify and estimate a statistically adequate unrestricted levels VAR of the system and its unrestricted VEC equivalent by utilizing statistically supported specification implications that arose from examination of the data in Figures 1–6, in order to capture elements of nonstationary behavior needed to obtain uncompromised inference, and non-spurious and valid regressions (Juselius 2004, chapter 5). Adequate specification generates residual estimates that behave with approximate multi-variate normality (see Juselius 2004, pp. 42, 49, and 50).

The Levels VAR and Unrestricted VEC of the Sugar-Based Market System

Bessler (1984) notes that a VAR model posits each endogenous variable as a function of k lags of itself and of each of the remaining endogenous variables

[4] An anonymous reviewer raised the valid issue of estimating policy-specific effects of each of the four policies with separate binaries. Given our limited sample size and number of degrees of freedom, our ultimately sizeable array of endogenous and deterministic variables, and our potential problems with small samples, we followed Dr. Katarina Juselius' recommendations and settled for a single binary to illuminate the net effect of the collective influences of all four policy interventions during the 1999:07–2004:03 subsample. Private communication of Dr. Katarina Juselius with an author, Copenhagen, Denmark, August 2004.

[5] These two events were offered as possible explanation for these permanent shifts in PCHOC and PNOCHOC slopes by the sugar-market analyst with the Agriculture and Forest Products Division of the U.S. International Trade Commission, in emails received by one of the authors on August 18 and 20, 2004.

[6] This section draws heavily on the work of Johansen and Juselius (1990, 1992) and Juselius (2004).

in the system. The above six variables render the following unrestricted VAR model in logged levels:

(1) X(t)=a(1,1)*QSUGAR(t–1)+…+a(1,k)*QSUGAR(t–k)+
a(2,1)*PRAW(t–1)+ . . .+a(2,k)*PRAW(t–k)+
a(3,1)*PCHOC(t–1) + ...+a(3,k)*PCHOC(t–k)+
a(4,1)*PNOCHOC(t–1)+…+a(4,k)*PNOCHOC(t–k)+
a(5,1)*PSOFT(t–1)+ . . . +a(5,k)*PSOFT(t–k) +
a(6,1)*PBAKERY(t–1)+…+a(6,k)*PBAKERY(t–k)+
a(c)*CONSTANT + a(D)* DETERMINISTIC + ε(t).

The asterisk denotes the multiplication operator. The residual vector ε(t) is distributed as white noise. X(t) = QSUGAR(t), PRAW(t), PCHOC(t), PNOCHOC(t), PSOFT(t), and PBAKERY(t). The a-coefficients are ordinary least-squares (OLS) estimates with the first parenthetical digit denoting the six endogenous variables as ordered above and with the second denoting lags 1, 2, . . ., k. The a(c) term is the coefficient on the intercept. The a(D) term denotes a vector of regression estimates on the various seasonal, binary, and trend deterministic variables (DETERMINISTIC) suggested by the study of Figures 1–6, and the inclusion of which is statistically supported in the analyses below. The parenthetical terms on the endogenous variables refer to the lag: t to the current value and t–k to the kth lag.

We applied Tiao and Box's (1978) lag-selection procedure that uses a likelihood-ratio test method, and which corrects for small samples. Results suggested a two-order lag structure (k=2).[7]

Johansen and Juselius (1990) and Juselius (2004, p. 66) demonstrated that the above VAR model in Equation 1 is now re-written more compactly into equation 2's unrestricted VEC format:

$$(2)\quad \Delta x(t) = \Gamma(1) * \Delta x(t-1) + \Pi * x(t-1) + \Phi * D(t) + \varepsilon(t) .$$

The ε(t) is distributed as white noise. The x(t) and x(t–1) are p × 1 vectors of the six sugar-related variables in current and lagged levels, Γ(1) is a p × p matrix of short run regression coefficients on the lagged differences, and Π is a (p × p) error-correction term on the six lagged levels variables. The Φ*D(t) is a set of statistically supported deterministic components added to utilize nonstationarity-related information as the analysis unfolds. The rank-unrestricted error-correction term is decomposed:

$$(3)\quad \Pi = \alpha * \beta' .$$

The α is a p × r matrix of adjustment-speed coefficients and β is a p × r vector of error-correction coefficients. The Π = α*β′ term is interchangeably denoted as the levels-based long-run component, error-correction term, or cointegration space of the model. The [Δx*(t–1), Φ*D(t)] is collectively considered the short-run/deterministic-model component.

Data analyses above suggested possible inclusion of a linear trend (TREND) and three permanent shift dummies (presented below) in Equation 2's long-run cointegration-space component. These same variables in differenced form, along with a set of centered seasonal variables, were considered for the short-run/deterministic component. Analysis will also lead to consideration, where appropriate, of various outlier binaries in the short-run/deterministic component.

The data analysis gave rise to the definition of three permanent shift dummies:

- a DS9907 shift binary was defined as unity for 1999:07–2004:03 (and zero otherwise) to account for the four above-cited U.S. sugar policy changes/interventions.

- a DS0101 shift binary was defined as unity for the 2001:01–2004:03 period (and zero otherwise) to account for the apparent shift in PCHOC trends from a 2001 increase surge in cocoa price arising from the Ivory Coast civil war and a late-2002 rise in prices of other confectionary materials prices.[8]

- a DS0212 shift binary was defined as unity for 2002:12–2004:03 (zero otherwise) to account for a permanent shift in PNOCHOC from a late-2002 rise in prices of non-cocoa of confectionary production inputs.[9]

[7] Given the 1992:01–2004:03 data sample located and the two-order lag, the estimation period for this study is 1992:03 through 2004:03.

[8] The source of this information was two emails from a U.S. International Trade Commission industry analyst responsible for monitoring markets for sugar and sugar-containing products. The emails were received August 18 and 19, 2004.

[9] This information was sourced from two emails from a U.S. International Trade Commission industry analyst responsible

The initial estimation starting point for the unrestricted VEC model was Equation 2 with only the seasonal variables included in the short-run/deterministic component of the model. A well-specified unrestricted VEC was achieved in a series of sequential estimations. These estimations added a linear trend and permanent shift and outlier binaries, one variable for each estimation, whereby the added variable was retained if a battery of specification diagnostic test values moved in favorable patterns indicative of improved specification. Juselius (2004, chapters 4, 7, and 9) recommends the following battery of diagnostics: (a) trace correlation as an overall goodness of fit indicator, (b) likelihood-ratio test of autocorrelation for the system, (c) univariate Doornik-Hansen tests for normality, (d) indicators for skewness and kurtosis, and (e) univariate tests for ARCH effects (heteroscedasticity). The successive estimations were stopped when the array of diagnostic values failed to improve with inclusions of additional binary variables. After achievement of an adequately specified unrestricted VEC, tests for parameter constancy and the existence of I(2) trends were performed.

The sequential estimations comprised two sets. The first included, one by one, the following in both the levels-based, long-run component (Π*X(t–1)) and in differenced form in the short-run/deterministic component of Equation 2: DS9907, DS0101, DS0212, and TREND. All four were ultimately included.

The second set of sequential estimations aimed to further improve specification of the unrestricted VEC just obtained above with the three permanent shift dummy variables and trend. More specifically, the second set of estimations captured the extraordinary effects of observation-specific events through the use of transitory "outlier" binaries. Each time a potential outlier was deemed potentially disruptive because of a "large" standardized residual,[10] an appropriately specified dummy variable was included in Equation 2's short-run/deterministic component and retained if the array of diagnostic values moved favorably to indicate improved specification. Ultimately, 13 observation-specific dummy variables were included.[11]

Equation 2 achieved adequate statistical specification with a trend and three permanent shift dummies included in the levels-based and short-run/deterministic component, and with seasonal vari-

for monitoring markets for sugar and sugar-containing product, and were received August 18 and 19, 2004. Note that DS0101 includes the events for which DS0212 was defined. This is because PCHOC was likely directly influenced by the increases in prices of both cocoa and non-cocoa confectionary materials, while PNOCHOC was likely affected directly by the rise in the costs of non-cocoa confectionary materials.

[10] We followed Juselius' (2004, chapter 6) procedure for examination and analysis of potential outlier dummy variables based on the Bonferoni criterion. An observation was judged as an "extraordinary outlier" if it generated a standardized residual within the 3.0–3.5 range. Such a criterion for outliers was designed based on the sample size using the Bonferoni criterion: INVNORMAL[$1.0 - 0.025)^{(1/T)}$], where INVNORMAL is the RATS instruction for the inverse of the normal distribution function that returns the variable for the c-density function of a standard normal distribution (Estima 2004, p. 503). Here, the Bonferoni value equals 3.5 for sample size T=147. Observations with standardized residuals with absolute values from 3.0 to 3.5 or more were considered outliers and observation-specific transitory binaries were specified, placed into the short-run/deterministic component of the model, and used for the above process of sequential and diagnostic monitoring. In several instances, clearly influential standardized residuals were nearer the value of 3.0 than 3.5, leading to our use of leeway in defining the criterion range as 3.0–3.5.

[11] To conserve space, we have not included the extensive variable-by-variable analysis from numerous estimations which resulted in inclusion of the 13 outlier binaries. All included outlier dummy variables were of the transitory "blip" format described in Juselius (2004, chapter 6). This form is appropriate for the short-run/deterministic component of the model and take on a value of 1.0 for the outlier observation, followed by a value of –1.0, and a value of zero otherwise. The 13 included observation-specific transitory blip dummy variables were so-defined for the following observations: 1992:07, 1992:12, 1993:09, 1994:03, 1994:06, 1995:08, 1996:01, 1999:11, 1996:08, 2000:03, 2000:07, 2000:10, 2004:01. The extraordinary effects captured by the outliers associated with the 1992:07, 1992:12, and 1993:09 observations likely arose from large expected and realized increases in sugar and sugar beet production from increases in planted area and yields during 1992/93 (USDA ERS 1993, p. 9). The outliers for 1994:03, 1994:06, and 1995:08 were likely associated with observation-specific and extraordinary influences elicited after the January 1994 and January 1995 implementations of the NAFTA agreement and the Uruguay Round of the WTO trade negotiations. Major U.S. farm-policy changes of the 1996 farm bill (FAIR ACT) likely generated extraordinary and month-specific effects captured by the outlier binaries defined for 1996:01 and 1998:08. The five outlier binaries defined for 1999:11, 2000:03, 2000:07, 2000:10, and 2004:01 may reflect additional month-specific influences over and beyond those generated by the four U.S. Government sugar policy changes that inspired DS9907's definition, and the rises in confectionary material costs that inspired the definitions of DS0101 and DS0212.

ables and 13-observation-specific binaries included in the short-run/deterministic component. Table 1 provides evidence of adequate specification.

Table 1 provides an array of diagnostic values focused on the statistical adequacy of two model estimations: the initially estimated unrestricted VEC before the sequential estimations and the unrestricted VEC model judged as adequately specified after inclusion of the various dummy variables (and TREND) suggested by analysis of the sequential estimations. Table 1's results demonstrate a marked improvement to specification and the benefits of utilizing information inherent in the data's nonstationary elements that Juselius feels are often not adequately considered.[12] Evidence suggests that the efforts to improve specification increased the model's ability to explain variation by more than 50 percent, since the trace correlation, a system-wide goodness of fit indicator, rose from 0.43 to 0.66. Additionally, ARCH effects did not appear to be a serious problem.

There were notable shifts in most equations' residual behavior toward statistically normal behavior, as evidenced by the univariate Doornik-Hansen (D-

[12] Dr. Katarina Juselius made this point during the econometrics course on cointegrated VAR methods. Her point is that many applications of time-series econometric techniques are done without getting an adequate handle on the properties of the data being modeled. One of our primary aims was to present the diagnostic results in Table 1 and show the value of considering the data-series properties when specifying VAR and VEC models. The course "Econometric Methodology and Macroeconomic Applications, the Copenhagen University 2004 Econometrics Summer School," taught by Drs. Katarina Juselius, Soren Johansen, Anders Rahbek, and Heino Bohn Nielsen, August 2–22, 2004, Institute of Economics, Copenhagen University, Denmark.

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

Test and/or equation	Null hypothesis and/or test explanation	Prior efforts at specification adequacy	After efforts at specification adequacy
Trace correlation	System-wide goodness of fit: large proportion desirable	0.43	0.66
ARCH(4) test for heteroscedasticity (lag 4)	Ho: no heteroscedasticity by 4th lag for system. Reject with p-values less 0.05	40.40 (p=0.28)	33.09 (p=0.61)
Doornik-Hansen test for normal residuals (univariate)	Ho: equation residuals are normal. Reject for values above 9.2 critical value		
ΔQSUGAR		4.9	7.2
ΔPRAW		66.8	6.8
ΔPCHOC		26.6	22.6
ΔPNOCHOC		19.6	1.7
ΔPSOFT		7.5	4.9
ΔPBAKERY		12.9	2.7
Skewness(kurtosis) univariate values	Skewness: ideal is zero; "small" absolute value acceptable Kurtosis: ideal is 3.0; acceptable is 3–5.		
ΔQSUGAR		0.26 (3.7)	0.44 (3.97)
ΔPRAW		–0.75 (8.73)	–0.008 (3.9)
ΔPCHOC		1.24 (6.78)	0.62 (5.61)
ΔPNOCHOC		1.00 (5.58)	0.25 (3.13)
ΔPSOFT		–0.13 (3.98)	–0.16 (3.72)
ΔPBAKERY		–0.22 (4.44)	–0.16 (3.44)

H) test values, which test the null hypothesis of an equation's residuals being statistically normal (Juselius 2004, chapter 4). D-H values suggest that four of the six equations generated non-normal residuals prior the specification effort, while five of the six apparently generated normally behaving residuals after such specification effort. The PRAW equation benefited most from the specification efforts, with the D-H value having fallen from a highly untenable 66.8 value initially to 6.8, well below the critical 9.2 value. D-H values for the PNOCHOC and PBAKERY equations also dropped noticeably.

Table 1 also provides indications on skewness and kurtosis of each equation's residuals. Normal behavior generates skewness values that are as near to zero as possible with acceptable (absolute) values being about unity or less; and kurtosis values that are as near to 3.0 as possible with acceptable values within the 3.0–4.0 range (see Juselius 2004, chapter 4). Specification efforts achieved clear improvements: five of the six skewness and kurtosis values declined noticeably, and final value ranges for both sets were considered acceptable for all equations except PCHOC.

Following Juselius (2004), we deemed these diagnostics generally acceptable, despite one of the six equations (chocolate-based confectionary product price) having failed to achieve statistically normal residual behavior, probably because of Figure 3's 2001 marked incline in PCHOC slope. The efforts did appreciably improve PCHOC specification, but not to a degree needed to suggest approximately normal residual behavior. Despite PCHOC's results, Table 1's results demonstrate that an adequately specified unrestricted VEC model has been achieved with which to exploit the system's cointegration properties.

Cointegration: Determining and Imposing Reduced Rank on the Cointegration Space

This paper's six endogenous variables are (later) shown to be nonstationary, or I(1); the first differences are stationary, or I(0); and cointegration implies that there are from one to five possible linear combinations which are stationary. Cointegrated variables are driven by common trends, and stationary linear combinations (or cointegration) arise when the nonstationarity of one variable corresponds to the nonstationarity in another (Juselius 2004, p. 86).

The Π-matrix (Equations 2 and 3) is a 6×6 matrix equal to the product of a $p \times r$ matrix, β, of long-run error-correction coefficients which combine into r stationary linear combinations of the six sugar-based variables (Johansen and Juselius 1990, 1992). As a result of a reduced rank for Π, $\beta'^{*}x(t)$ is stationary even though each of the six sugar-based series is shown to be nonstationary.

Determination of cointegration rank is a four-tiered process (Juselius 2004, chapter 8). First, we conduct Johansen and Juselius' (1990, 1992) trace tests. Second, we examine patterns of the characteristic unit roots in the companion matrix under chosen rank. Third, we consider average cycle durations of the monthly data. Fourth, we analyze patterns of relevant α-estimates. Finally, plotted cointegrating relations are examined for stationarity properties.

Nested Trace Tests and Other Evidence for Rank Determination of the Error-Correction Matrix, Π

Table 2 provides trace test evidence for rank determination. Trace tests are corrected (increased) with Bartlett's adjustment factor for small samples, while the 95-percent fractile values were adjusted for restriction of three permanent shift dummies into the cointegration space (see Table 2's notes and Juselius 2004, chapter 8). Evidence at the 95-percent significance level is sufficient to reject the first three nested null hypotheses, and fails to reject the fourth hypothesis, that rank is three or less. Taken alone, this evidence suggests that there are three long-run cointegrating relations which error-correct the system of six individually nonstationary series to move tandemly through time. Evidence generated by the fourth test (rank = 3 or less) may be of borderline strength, as the 45.3 trace value approaches the corrected fractile of 47.9. Sole reliance on these trace-test results to determine rank is not recommended (Juselius 2004, chapter 8).

Table 3 provides information on the companion matrix's roots for the estimated VAR model. The number of unit characteristic roots in Table 3 corresponds to the number of common stochastic trends ($p - r = 3$) when a rank of 3 is imposed: there appear to be $r = 3$ relations toward which the process is adjusting and p-r or 3 common trends driving or pushing the system (Juselius 2004, p. 173). If $r = 3$ is an appropriate choice, then there should be $p - r = 3$ unity-valued characteristic roots with the

Table 2. Trace-Test Statistics and Related Information for Nested Tests for Rank Determination.

Null Hypothesis	Eigen value	Bartlett-corrected trace value	95% Fractile (critical value)	Result
rank or r ≤ 0	0.3678	198.3	122.9	Reject null that rank is zero.
rank or r ≤ 1	0.344	140.5	94	Reject null that rank ≤ 1
rank or r ≤ 2	0.2375	84.4	69.1	Reject null that rank ≤ 2
rank or r ≤ 3	0.1565	45.3	47.9	Fail to reject null that rank ≤ 3
rank or r ≤ 4	0.092	17.9	31.1	Fail to reject null that rank ≤ 4
rank or r ≤ 5	0.0764	9	17.9	Fail to reject that rank ≤ 5

Notes: The trace-test values are corrected for small samples with the Bartlett adjustment generated by the CATS2 program. As recommended by Juselius (2004, p. 171), CATS2-generated fractiles are increased by 3*1.8 (5.4) to account for the 3 shift dummy variables restricted to lie in the cointegration relations.

Table 3. Roots of the Companion Matrix, Rank = 3.

	Real	Imaginary	Modulus	Argument
Root1	1.0000	0.0000	1.0000	0.0000
Root2	1.0000	0.0000	1.0000	0.0000
Root3	1.0000	0.0000	1.0000	0.0000
Root4	0.8488	0.0000	0.8488	0.0000
Root5	0.6787	0.3023	0.7429	0.4191
Root6	0.6787	–0.3023	0.7429	0.4191
Root7	0.5110	–0.4832	0.7033	–0.7575
Root8	0.5110	0.4832	0.7033	0.7575
Root9	0.3209	–0.0000	0.3209	–0.0000
Root10	0.1629	–0.0000	0.1629	–0.0000
Root11	–0.1281	–0.0000	0.1281	–3.1416
Root12	0.0695	–0.0000	0.0695	–0.0000

fourth being statistically sub-unity (Juselius 2004, pp. 173–174). Table 3's fourth root, 0.85, is likely statistically sub-unity and reinforces the appropriateness of having chosen rank as r = 3 (Juselius 2004, chapter 8).

We followed Juselius (2004, p. 275) and noted that because some of CV3's α-estimates are statistically significant, CV's inclusion results in contributions to Π's error-correction process. In turn, this supports the choice of r = 3 as the appropriate Π-rank and number of CVs.[13]

Further support of r = 3 as the appropriate rank arises if the model's average cycling duration rea-

[13] To conserve space, we do not provide all of the cointegration results for the pre-final unrestricted VEC model. Note that the third cointegration relation's alpha vector is as follows, with t-values: α(QSUGAR) = –0.0053, t = –1.73; α(PRAW) = –0.0041, t = –2.30; α(PCHOC) = 0.0004, t = 0.29; α(PNOCHOC) = 0.001, t = 3.72; α(PSOFT) = 0.001, t = 2.50; and α(DPBAKERY) = 0.0006, t = 1.82.

sonably short. Cycles of mean-reverting behavior average six or seven months, which is deemed reasonable when monthly data are involved.[14]

The plots of the three cointegrating relationships in Figures 7, 8, and 9 suggest that the CVs behave in patterns that approximate stationarity: plots frequently revert to the zero mean in non-enduring cycles (see Juselius 2004, pp. 172–176). Juselius (2004, chapter 8) recommends focusing on the β′*R models corrected for the short run components. Generally, all three relationships exhibit stationary behavior, despite a CV1 outlier in 1999 (Figure 7); some CV2 ARCH effects in the early third of the sample (Figure 8); and some early-sample CV3 cycling that ultimately winds down (Figure 9).

We conclude that a rank of three is likely appropriate, after having considered evidence from the rank tests, patterns of the companion matrix's characteristic roots, significance of relevant α-estimates, average data-cycling durations, and a visual inspection of the plotted cointegrating relationships.

Two Final Diagnostic Tests: Parameter Constancy and I(2) Trends

Two final diagnostic tests were applied to the cointegrated VEC, and results further suggested that the model achieved statistical adequacy. The tests are for constancy of error-correction parameter estimates and for the presence of trends that are integrated of order 2, or I(2).

The "known beta" test of parameter constancy detailed in Juselius (2004, pp. 186–190) tests if there is constancy of cointegration parameters. The method tests if the full-sample "baseline" model's cointegration relations could have been accepted as those of each recursively estimated model over the 1999:07–2004:03, a period chosen because of its inclusion of a number of previously cited events that greatly influenced our modeled markets. Normalized by the 5-percent critical value, values are ideally unity or less. Since all values were below unity, evidence suggests that time-variance of parameters is likely not an issue. Because of space limitations, we do not provide the plotted known-beta values.

Nielsen (2002) and Juselius (2004, chapter 16) note that imposing reduced rank restrictions on an unrestricted VEC's Π-matrix when there are I(2) trends in the data encounters well-known adverse statistical consequences because the modeled data is still nonstationary. Evidence from Nielson's (2002) series of tests for I(2) trends suggest that I(2) trends are likely not an issue in our model.[15]

Equations 4–6 below are the three cointegrating relationships which emerged from imposing rank using Johansen and Juselius' (1990, 1992) reduced-rank estimator. These estimates are not yet restricted for evidentially supported economic restrictions which emerge from the next section's hypothesis tests.

(4) QSUGAR = –4.6*PRAW – 4.7*PCHOC – 22.8*PNOCHOC – 9.45*PSOFT + 16.8*PBAKERY – 0.85*DS9907 + 0.10*DS0101 + 0.87*DS0101 + 0.03*TREND

(5) PCHOC = –0.26*QSUGAR + 0.26*PRAW – 0.32*PNOCHOC – 1.3*PSOFT + 0.47*PBAKERY + 0.07*DS9907 + 0.02*DS0101 + 0.07*DS0212 + 0.004*TREND

(6) PRAW = 0.42*QSUGAR – 0.10*PCHOC + 9.62*PNOCHOC + 4.84*PSOFT – 2.96*PBAKERY – 0.26*DS9907 + 0.12*DS0101 –0.13*DS0212 – 0.01*TREND.

Hypothesis Tests and Inference on the Economic Content of the Three Cointegrating Relations

Determination of the reduced rank for the cointegrating space separates the system's six eigenval-

[14]According to Juselius, 1.0 less the near-unity root of 0.85 equals 0.15, which, when divided into 1.0 provides 6.7, a value reflecting average cycling lengths of 6–7 months. Private communication with K. Juselius, University of Copenhagen, Denmark, August, 2004. Also see Juselius (2004, pp. 172–174).

[15] Nielsen (2002) and Juselius (2004, chapter 16) provide a series of values that each test a null hypothesis of the model restricted for I(2) trends against the model without such restrictions. Given that p = 6 and r = 3, 27 tests were calculated; they are not provided here due to space considerations, but are available from the authors on request. Evidence was sufficient at the 5-percent significance level to reject the null of I(2) trends in all cases except one, with the excepted case rejecting the null at a very marginal 7-percent significance level. Since evidence was sufficient to reject the null at the 10-percent level, at the 7-percent level, and marginally at the 5-percent level, we conclude that I(2) trends were likely not an issue.

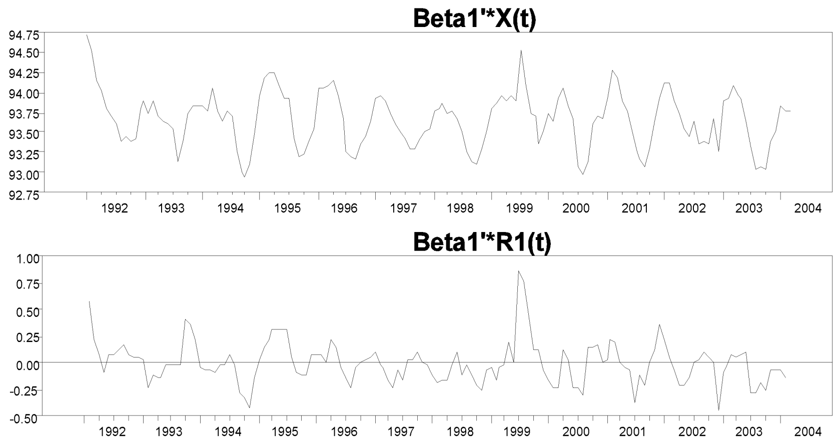

Figure 7. Plotted Cointegrating Relation 1, Unrestricted VEC: Versions With and Without Correction for Short Run Effects.

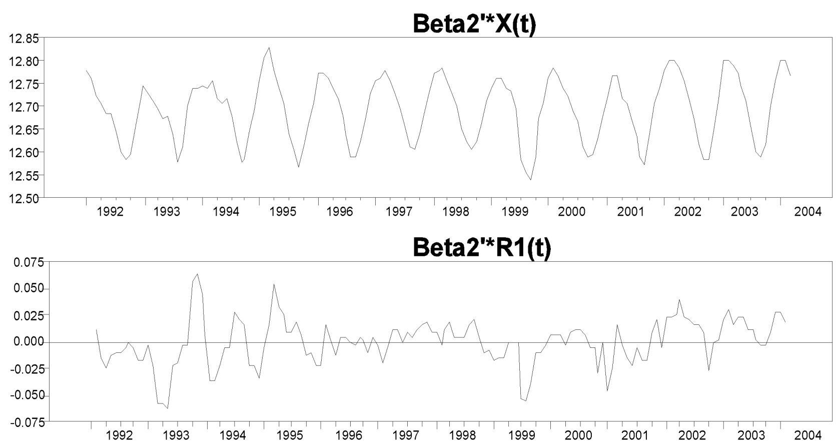

Figure 8. Plotted Cointegrating Relation 2, Unrestricted VEC: Versions With and Without Correction for Short Run Effects.

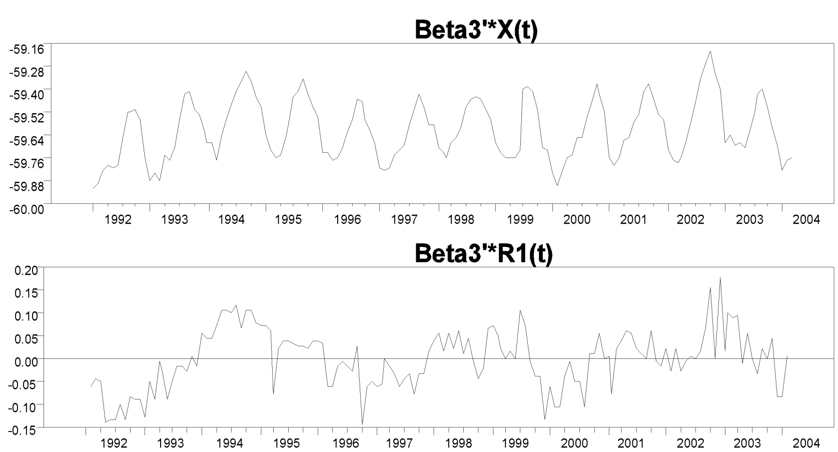

Figure 9. Plotted Cointegrating Relation 3, Unrestricted VEC: Versions with and Without Correction.

ues into r = 3 stationary eigenvalues and (p – r) = 3 nonstationary vectors (Juselius, 2004, p. 205). One begins with equations 4–6 above, the three unrestricted cointegrating vectors, and conducts a series of hypothesis tests on the $\alpha\beta' = \Pi$, using well-known procedures detailed in Johansen and Juselius (1990, pp 194–206) and Juselius (2004, chapter 10). Hypothesis tests on the β-coefficients take the form

(7) $\beta = H^*\varphi$,

where β is a p1 × r vector of coefficients on variables included in the coingetration space; H is a p1 × s design matrix, with "s" being the number of non-restricted or "free" beta coefficients; and φ is an s × r matrix of the unrestricted beta coefficients (Juselius 2004, pp. 245–248). Johansen and Juselius' (1990, 1992) well-known hypothesis test value is provided in Equation 8:

(8) $-2\ln(Q) = T^*\sum\ln[(1-\lambda_i^*)/(1-\lambda_i)]$,
for i = 1, 2, and r = 3.

The asterisked (nonasterisked) eigenvalues λ_i are generated by the model estimated with (without) the tested restriction(s) imposed.

The α, or adjustment-speed coefficients, characterize the relative speeds of error-correcting adjustment with which the system responds to a given shock (Johansen and Juselius 1990, 1992). The null hypothesis on the αs take the form

(9) H(0): $\alpha = A^*\psi$,

where A is a p × s design matrix, s is the number of unrestricted coefficients in each of the r = 3 columns of the p × r α-matrix, and ψ is the s × r matrix of the non-restricted or "free" adjustment-speed coefficients (Juselius 2004, chapter 11). Equation 8's test statistic applies here, and is distributed asymptotically as chi-squared distribution with degrees of freedom equal to the number of imposed coefficient restrictions (Juselius 2004, pp. 206–207 and 211–213)

Hypothesis Tests on the Betas

Three groups of hypothesis tests are performed on the β-estimates of the rank-reduced Equations 4–6. The first group consists of six stationarity tests for the endogenous variables and the second consists of 10 exclusion tests on all endogenous and deterministic variables in the cointegration space. The third group is a set of sequential hypothesis tests on selected β-estimates in Equations 4–5.

Tests of Stationarity. Juselius (2004, pp. 220–222) recommends a likelihood-ratio test of each individual endogenous variable's stationarity within a system setting and given the imposed rank (here r = 3), over univariate stationarity tests (e.g., Dickey-Fuller tests) which are independent of the cointegrated system and its chosen rank. Such tests examine if each endogenous variable itself constitutes a separate stationary cointegrating relation, using a unity value for the tested variable's betas. Equation 7 then takes the following form (Juselius 2004, p. 221):

(10) $\beta^c = [b,\varphi]$,

where β^c is the p1 × r (10 × 3) beta matrix with one of the variable's levels restricted to a unit vector and b is a p1 (or 10) × 1 vector with a unity value corresponding to the relevant variable whose stationarity is being tested and for the four deterministic components restricted to the cointegration space and retained as non-restricted,[16] and φ is a p1 × (r – 1) or 10 × 2 matrix of the remaining two unrestricted cointegrating vectors (Juselius 2004, 221). Evidence was sufficient to reject that each of the six endogenous variables was stationary.[17]

[16] This test can be conducted in CATS2 (beta version) in two settings: with and without inclusion of these four deterministic variables restricted to the cointegration space: DS9907, DS0101, DS0212, and TREND. Given the above-cited evidence on the importance of these variables and associated events, we included them in the test.

[17] Given the rank of 3, then Equation 8's test values and parenthetical p-values are as follows, with the null of stationarity rejected for "small" p-values below 0.05: 23.0 (0.0000) for QSUGAR, 7.4 (0.06) for PRAW, 12.4 (0.006) for PCHOC, 11.5 (0.009) for PNOCHOC, 22.6 (0.0000) for PSOFT, and 15.9 (0.001) for PBAKERY. PRAW's result is borderline, and we deemed the variable stationary because of the clear elements of nonstationarity which emerged from the data analysis (Figure 2), and because the stationarity test without the four deterministic components suggested stationarity, with a test variable of 28.6 (0.0002). Treating PRAW as nonstationary seemed appropriate given the test value's closeness to the critical value and because of this other evidence.

Tests of Beta Exclusions. Exclusion tests examine whether each of the 10 variables restricted to the cointegration space have zero coefficients in the model's three CV's using the hypothesis test in Equation 7 and test value in Equation 8.[18] Failure to reject the null of a variable's zero-valued β's suggests that the variable should be excluded from the cointegration space. In all cases except the β(DS0101)'s, the evidence rejected the null of zero-valued β's in the r = 3 cointegrating relations and supported retention of these variables in the cointegration space.[19] Additional evidence from the unrestricted VEC's Π estimates and market knowledge led to the retention of DS0101 β-estimates in the cointegration space, despite the low test exclusion value.[20]

Set of Sequential Hypothesis Tests on Individual Beta Coefficients. After imposing reduced rank (r = 3) on the cointegration space's Π-matrix, and after a stationary long-run structure emerges as Equations 4–6, these cointegrating relationships must be identified (Juselius 2004, pp. 245–246). Under the rank condition of identification, the three CV's are identified by imposing at least r – 1 or 2 restrictions on each of Equations 4, 5, and 6 (Juselius 2004, pp. 245–246). Equations 7 and 8 are used to test (and possibly impose) hypotheses chosen based on market expertise, theory, patterns of regression estimates, and/or conditions needed to meet the rank condition of identification.[21] The relations are re-estimated with the reduced-rank estimator while imposing the last-accepted and statistically supported restriction(s). This process is repeated to obtain a set of finally-restricted cointegrating relationships.

This multi-iterative process is summarized in Table 4. Following Juselius' (2004, chapters 10–12) procedures, we first tested two hypotheses of β-equality which were very evident from the unrestricted estimates in Equations 4–6: tests on whether the betas on PRAW and PCHOC are equal in the first relation, or Equation 4 (along with restrictions necessary to identify the system), comprising Test Set 1, and tests on whether the betas on QSUGAR and PRAW in the second relation (Equation 5) are negatively related and equal in absolute value (along with conditions necessary to identify the system), comprising Test Set 2. Evidence in Table 4 was sufficient to reject Test Set 2 (p-value = f 0.049) and insufficient to reject Test Set 1 (p-value = 0.128). We retained these set-1 restrictions, re-estimated the system with the Johansen-Juselius reduced-rank estimator, and conducted a stream of sequential tests on the three re-estimated cointegrating relations. Each sequential test focused on additional testable restrictions that became evident (Juselius 2004, chapters 10–12). Space considerations led to only a summary of the sequential hypothesis tests and reduced rank estimations in Table 4, without having reported results of the various estimations. The finally restricted cointegrating relations are in Equations 11, 12, and 13.

Because of an insignificant β(DS0212) t-value, the first and most obvious was that β(DS0212) = 0, which was added to form Test Set 3 (Table 4).

[18] Equation 7 includes a p1 × r β-vector; a p1 (= 10) × 9 design matrix H with 9 being the nine unrestricted beta coefficients in each relation; and a 9 × 3 matrix φ of the nine unrestricted coefficients in each of the three cointegrating relationships. Basically, the φ-matrix is the β-matrix without the betas of the variable being tested for exclusion. See Juselius (2004, chapter 10).

[19] The test values, each distributed as a chi-square distribution with three degrees of freedom, (and parenthetical p-values) were as follows, with the null of a variable's three beta coefficients being rejected for p-values of 0.05 or less (that is, for 5%): 20.6 (0.0001) for QSUGAR; 30.7 (0.0000) for PRAW; 37.2 (0.0000) for PCHOC; 21.5 (0.0001) for PNOCHOC; 22.7 (0.0000) for PSOFT; 23.4 (0.0000) for PBAKERY; 26.5 (0.0000) for DS9907; 23.0(0.0000) for DS0212; and 10.1 (0.0003) for TREND. The test value for DS0101's coefficients equaling zero was 4.8 (0.19), which suggests that evidence at the five percent level was insufficient to reject the null hypothesis that β(DS0101) had betas in the three cointegrating relations that were zero.

[20] Of the six DS0101-related elements with the Π matrix restricted for a reduced rank of r = 3, four were statistically nonzero at the 5-percent level, and five were statistically nonzero at the 10-percent level, suggesting that the DS0101 dummy variable defined to capture the 2001 sustained rise in cocoa prices and the late-2002 rise in prices of other confectionary product inputs was important to the error-correction process. Also, market experts suggested that inclusion of this DS0101 was probably important to this cointegrated system (two emails from a U.S. International Trade Commission industry analyst responsible for monitoring markets for sugar and sugar-containing products, received August 18 and 20, 2004.

[21] This third set of beta hypothesis tests should be conducted while retaining any statistically supported restrictions from the stationarity and exclusion hypothesis tests done above. Since evidence rejected all stationarity and exclusion restrictions, no restrictions carried over to this third set of hypothesis tests. In effect, the third set of hypothesis tests on individual betas were applied directly to Equations 4, 5, and 6.

Table 4. Four Sets of Sequential Hypothesis Tests on Specific Beta Estimates or Beta Estimate Subsets.

Tested restrictions, restriction numbers in each cointegrating vector (CV)	Explanation/reasons	Test value, parenthetical p-value and test results.
Test Set 1: β(PRAW) = β(PCHOC) in CV1 plus identifying restrictions in the 3 cointegrating vectors or CVs		
4 in CV1: β(PRAW) = β(PCHOC); β(PNOCHOC) = β (PSOFT) = β(PBAKERY) = 0	Near-equal coefficients; identifying zero restrictions.	Chi-squared test value = 4.12 (0.128). Evidence was insufficient to reject the tested restrictions since the p-value exceeded 0.05.
2 in CV2: β(QSUGAR) = β(PRAW) = 0	Identifying zero restrictions.	
2 in CV3: β(QSUGAR) = β(PCHOC) = 0	Identifying zero restrictions.	
Test Set 2: β(QSUGAR) = – β(PRAW) plus identifying restrictions in the 3 CVs.		
3 in CV1: β (PNOCHOC) = β (PSOFT) = β(PBAKERY) = 0	Identifying zero restrictions.	Chi-squared test value = 3.86 (0.049). The p-value of 0.049 suggests that evidence at the 5 percent level was sufficient to reject the tested restrictions.
2 in CV2: β(QSUGAR) = – β(PRAW);	Nearly proportional, negatively related coefficients;	
β(DS9907) = 0;	Shift dummy for sugar market perhaps not significant and/or relevant to chocolate-based confectionary price, PCHOC;	
2 in CV3: β(QSUGAR) = β(PRAW) = 0	Identifying zero restrictions	
Test Set 3: β(DS0212) = 0 in CV1 plus the accepted restrictions above on 3 CVs in Test Set 1.		
5 in CV1: β(DS(0212) = 0; β(PRAW) = β(PCHOC) = 0; β (PNOCHOC) = β(PSOFT) = β(PBAKERY) = 0;	Insignificant t-value (-1.3) in Test Set 1; Previously accepted from Test Set 1; Previously accepted identifying restrictions from Test Set 1.	Chi-squared test value of 4.42 (0.219). Evidence was insufficient to reject the tested restrictions since the p-value exceeded 0.05.
2 in CV2: β(QSUGAR) = β(PRAW) = 0;	Previously accepted identifying restrictions.	
2 in CV3: β(QSUGAR) = β(PCHOC) = 0;	Previously accepted identifying restrictions.	
Test Set 4: β(DS9907) = 0 in CV2 plus accepted restrictions accepted in Test Sets 1 and 3 in all 3 CVs.		
5 in CV1: β(DS(0212) = 0; β(PRAW) = β(PCHOC) = 0; β (PNOCHOC) = β(PSOFT) = β(PBAKERY) = 0;	Previously accepted restrictions accepted in Test Sets 1 and 3.	Chi-squared value of 4.4 (0.35). Evidence was insufficient to reject the tested restrictions as the p-vale exceeded 0.05.
3 in CV2: β(DS9907) = 0; β(QSUGAR) = β(PRAW) = 0;	Insignificant t-value (-0.02) in Test Set 2; Previously accepted identifying restrictions.	
2 in CV3: β(QSUGAR) = β(PCHOC) = 0;	Previously accepted identifying restrictions.	
Test set 5: β(DS0101) = 0 in CV3 with all of Test Set 4's retained restrictions.		
5 in CV1: β(DS(0212) = 0; β(PRAW) = β(PCHOC) = 0; β (PNOCHOC) = β(PSOFT) = β(PBAKERY) = 0;	Previously accepted restrictions from Test Sets 1, 2, and 3.	Chi-aquared value of 4.4 (0.489). Evidence was insufficient to reject the tested restrictions as the p-vale exceeded 0.05.
3 in CV2: β(DS9907) = 0; β(QSUGAR) = β(PRAW) = 0;	Previously accepted restrictions from Test Sets 1, 2, 3, and 4.	
3 in CV3: β(DS0101) = 0;	Insignificant t-value (-0.7) on β(DS0101) in Test Set 4.	
β(QSUGAR) = β(PCHOC) = 0	Previously accepted identifying restrictions.	

Evidence was insufficient to reject Test Set 3 (test statistic p-value of 0.22), and this led to the retention of a zero restriction on β(DS0212) in CV1 (Equation 4) and the re-estimation with the reduced-rank estimator. This process was repeated on all three CV's in succession, and led to Table 4's Test Sets 4 and 5, which evidence failed to reject.

Three rationales for having chosen additional hypotheses leading to Table 4's Test Sets 3, 4, and 5 emerged from consideration of economic theory, market knowledge, policy considerations, and the finally restricted CV's in equations 11–13:

- While DS0212's significance in CV3 (equation 13) suggests that the non-cocoa input-price surges appreciably influenced prices of chocolate-based confectionary products, these cocoa-price increases were likely not sufficient to ultimately affect sugar demand (Equation 11), since the data accepted Test Set 3's restriction that β(DS0212) = 0 in CV1.

- While DS9907's observed significance in sugar demand and price (Equations 11 and 13) was expected, the four U.S. policy interventions apparently did not directly influence prices of chocolate-based confectionary products, since the evidence accepted Test Set 4's added restriction of β(DS9907) = 0 in the finally restricted PCHOC relation (Equation 12).

- While DS0101's significance in CV1 and CV2 suggest that the 2001–2002 confectionary input-cost increases did affect sugar demand and PCHOC, the evidence's acceptance of the Test Set 5 restriction of β(DS0101) = 0 in CV3 suggests that these cost increases did not directly influence the price of raw sugar.

Individual signs and values of the finally restricted CV estimates in Equations 11–13 are provided below in the section on economic analysis of the finally restricted cointegrating relations.

Hypothesis Tests on the Adjustment-Speed Coefficients

A principal hypothesis test on the estimated adjustment-speed coefficients is determining if each of the variables is weakly exogenous using Equations 8 and 9. The test for weak exogeneity tests whether a variable influences the others in the error-correction system without itself adjusting or responding to the process, thereby implying a one-way causal relation from said variable to the equilibrating relation. The hypothesis is equivalent to testing if, given the statistical significance of at least some of a variable's β-estimates, the variable's (here, r = 3) adjustment-speed coefficients are all zero (Juselius 2004, pp. 231–232). Evidence in all cases was sufficient to reject the null of weak exogeneity.[22]

Economic Analysis of the Three Cointegrating Relationships for U.S. Sugar-Related Markets

We economically interpret the identified and fully restricted set of cointegrating relationships in Equations 11, 12, and 13, which are followed by the adjustment-speed (or α) coefficients. Parenthetical values below the β- and α-estimates are t-values. Having modeled in natural logarithms, the β-estimates are interpreted as long-run elasticities in Equations 11–13.

$$(11)\quad QSUGAR = \underset{(-24.12)}{-5.02}{*}PRAW - \underset{(-24.12)}{5.02}{*}PCHOC - \underset{(-8.2)}{0.91}{*}DS9907 + \underset{(4.4)}{0.50}{*}DS0101 + \underset{(9.6)}{0.014}{*}TREND$$

$$(12)\quad PCHOC = \underset{(7.36)}{20.32}{*}PNOCHOC + \underset{(4.32)}{7.47}{*}PSOFT - \underset{(-6.2)}{12.28}{*}PBAKERY - \underset{(7.5)}{0.66}{*}DS0212 + \underset{(4.0)}{0.30}{*}DS0101 - \underset{(-2.68)}{0.014}{*}TREND$$

[22] In Equation 9, and given that r = 3 and p = 6, α is a 6 × 3 matrix of alphas with a row of zeros for the variable being tested for weak exogeneity, A is a 6 × 5 design matrix (with 5 being the number of nonzero alphas in each of the three columns of alphas), and ψ is a 5 × 3 matrix of nonzero alphas. Equation 8's weak exogeneity test values are each distributed as a chi-square variable (three degrees of freedom) and were calculated as follows with (parenthetical) p-values: 14.7 (0.002) for QSUGAR, 22.2 (0.000) for PRAW, 28.6 (0.000) for PCHOC, 9.0 (0.03) for PNOCHOC, 11.2 (0.01) for PSOFT, and 12.6 (0.005) for PBAKERY. Evidence was sufficient to reject the null hypothesis of a variable's zero-valued α-coefficients when the test value p-values fell below 0.05 (coinciding with the 5-percent significance level). See Juselius (2004, pp. 231–232).

$$\text{(13) PRAW} = \underset{(-6.12)}{-4.53}\text{*PNOCHOC} - \underset{(-1.92)}{0.92}\text{*PSOFT} + \underset{(4.9)}{2.66}\text{*PBAKERY} - \underset{(-10.7)}{0.16}\text{*DS9907} + \underset{(6.86)}{0.16}\text{*DS0212} + \underset{(2.33)}{0.003}\text{*TREND}$$

	Alpha1	Alpha2	Alpha3
ΔQSUGAR	–0.1069	0.1398	0.4831
	(–3.7630)	(3.0133)	(2.9391)
ΔPRAW	0.0027	–0.0404	–0.2582
	(0.1713)	(–1.5769)	(–2.8466)
ΔPCHOC	–0.0721	0.0996	0.3603
	(–6.4254)	(5.4320)	(5.5463)
ΔPNOCHOC	–0.0043	0.0118	0.0311
	(–1.4341)	(2.4076)	(1.7866)
ΔPSOFT	–0.0165	0.0275	0.0921
	(–3.6132)	(3.6981)	(3.4913)
ΔPBAKERY	–0.0031	–0.0009	0.0229
	(–1.0997)	(–0.1998)	(1.4164)

Our long-run relationships among U.S. sugar and sugar-using products in Equations 11 and 13 somewhat coincide with Moss and Schmitz's (2004) result that U.S. monthly and refined sugar prices currently appear stably cointegrated. We realize, however, that comparing our results from a 6-series vector of price and quantity variables with Moss and Schmitz's bivariate and trivariate work on U.S. raw sugar, refined sugar, and HFCS prices should be done carefully.

We note initially that the first CV, which focuses on the raw sugar market, has a more precise structural interpretation as a sugar demand, probably because that market's available information set included both prices and quantities. CV1 thus provided a reasonably clear structural-demand relationship based on a relatively more complete sugar market information set not available for the downstream markets. Such monthly data are generally business-proprietary data and unavailable for downstream commodity-based manufactured products (see Babula, Bessler, and Payne 2004). Lack of equally complete information for the downstream markets likely rendered the two price transmission CVs in Equations 12 and 13 with less precise, non-structural, and reduced-form interpretations that included elements of both demand and supply. Our less precise reduced-form interpretations for Equations 12–13 is common in cointegration analysis, as recently noted by Juselius:

> It is important to note that the cointegration rank is not in general equivalent to the number of theoretical equilibrium relations derived from an economic model. . . . Thus, cointegration between variables is a statistical property of the data and only exceptionally can be given a direct interpretation as an economic equilibrium relation. The reason for this is that a theoretically meaningful relation can be (and often is) a linear combination of several "irreducible" cointegration relations. (2004, p. 175)

As such, the CVs in Equations 12 and 13 each may be a reduced-form combination of several irreducible relations, without a clear structural demand-side or supply-side interpretation, but inclusive of interpreted elements of both. Yet the literature establishes that such error-correcting transmission relations are of economic interest and have policy-relevance, as seen by Bessler, Yang, and Wongcharupon's (2002) work on world wheat prices. Additionally, Williams and Bessler (1997) and Moss and Schmitz (2002, 2004) found bivariate and trivariate CVs of monthly U.S. sugar and HFCS prices of great use in illuminating how related or potentially related sweetener markets interact. Our six-series data set and economic analysis makes first-cut progress in illuminating one long-run structural and two reduced-form relationships among important U.S. upstream and downstream sugar-based markets. We must relegate more complete economic structural interpretations for Equations 12 and 13 to future research, when more comprehensive downstream information sets (i.e., quantities) may become available.

Cointegrating Relation 1: An Economy-Wide Demand for Sugar as an Input

Equation 11 appears to be a U.S. economy-wide demand for raw sugar as an input. Own-price (PRAW) has a negative and strongly significant β-coefficient. Normalizing on sugar quantity conceptually places it in positive form "above the equilibrium" on the relation's left side, whereby the α-estimate corresponding to QSUGAR in the first relation is a strongly significant –0.1067, and suggestive of a demand's downward adjustment toward the long-run equilibrium or "attractor set."

The own-price elasticity estimate of raw sugar demand is –5.0, and may initially seem a rather elastic estimate for a raw farm commodity. However, the modeled QSUGAR series has a standard error (SE) of 26 percent, compared with only 9.9 percent for PCHOC and an even smaller 5.9 percent for PRAW. There appear to be far higher QSUGAR variation levels relative to levels of PRAW (and PCHOC) variation over time, and hence a high elasticity level. Perhaps this –5.0 own-price elasticity of demand is, as discussed below, inflated by mutually offsetting U.S. sugar policies such as those discussed above. Future research would do well to determine if a more inelastic parameter would emerge in the absence of such exogenous policy influences.

More specifically, perhaps PRAW's ability to clear the raw sugar market has been stymied by the four exogenous seemingly nonsystematic and mutually offsetting U.S. sugar-policy interventions. We offer three arguments. First, when one considers that the SE of PCHOC, an aggregate price index of more than 50 U.S. sugar-related manufactured prices, is nearly double the SE of PRAW, a one- or mono-commodity price, one can intuitively question whether PRAW's behavior is artificially constrained.[23] Our intuition and extended experience with food-related prices certainly suggests the opposite situation, where the SE of a 50-product industrial-price aggregate should be less than the SE for mono-commodity raw sugar price near the U.S. farmgate. Second, PRAW behaves in what Juselius (2004) would characterize as a weakly exogenous manner in CV1: the statistically significant β(PRAW) and the statistically insignificant and near-zero α(PRAW) suggest that raw price influences but does not fully respond to CV1's error-correcting demand relationship.[24] Williams and Bessler's (1997, pp. 227–229) earlier work with bivariate models of U.S. sugar and HFCS prices also uncovered β- and α-estimate patterns that suggested that U.S. sugar price's weak exogeneity. And third, β(DS9907) was designed to capture net effects of the four noted U.S. sugar-policy interventions, and has a very significant and negative effect on PRAW in CV3, suggesting that PRAW was indeed highly influenced by the interventions as intuitively suggested Figure 2. Given PRAW's seemingly unnaturally sluggish variation levels, PRAW's apparently less than fully endogenous participation in the demand relation, and the strong significance of β(DS9907) on PRAW in CV3, we conclude that PRAW has likely been impeded in fully servicing the U.S. sugar market by exogenous U.S. sugar-policy interventions.

Evidence also strongly supports the hypothesis that PCHOC equally and negatively influences QSUGAR as own-price: the large t-values and the unambiguous hypothesis test results for Table 4's test on PRAW/PHCOC β-estimate equality in the demand relation. Even more interestingly, PCHOC has α- and β-estimates that are both significant in CV1, suggesting that PCHOC influences and adjusts to the CV1 error-correcting mechanism in a more fully endogenous manner than even own-price. When confronted with a partially adjusting PRAW or own-price in CV1, perhaps market agents need to turn elsewhere for signals to aid in market-clearing decisions. It seems that sugar-demanding agents monitor both equally: a policy-constrained own-price and a fully endogenous PCHOC. Apparently, agents consider output prices of major sugar-based manufactured products represented by PCHOC as a "co-own-price," with the latter's increases/decreases signaling ultimately decreased/increased needs for sugar as a production input. The validity of this relationship should be examined by future researchers, particularly if more expansive downstream information sets emerge.

A number of implications arise from Equation 11's results for research and for policy makers reliant on such research. First, the USDA's published raw sugar price may have behavior that is policy-impeded to such a degree that it may not serve well as a fully participating and endogenous own-price with which to empirically estimate valid U.S. sugar demands and own-price demand elasticities. The implication is that perhaps our elasticity estimate is artificially high from policy interventions, and the parameter may become more elastic in the

[23] According to a U.S. Department of Labor staff analyst, PCHOC used here is an industry price aggregate incorporating prices of products for the entire U.S. chocolate-based confectionary products subsector. And while Department of Labor staff would not divulge sampling information, the analyst stated that PCHOC was aggregated over more than 50 separate industrial-product prices.

[24] One should note that PRAW was not weakly exogenous *to the system*, as test evidence above on zero-valued α-estimates for *all CVs* rejected the hypothesis of PRAW's weak exogeneity to the system. Nonetheless, in CV1 alone, PRAW does have an insignificant α-estimate and a significant β-estimate that would suggest that PRAW is not fully endogenous, weakly exogenous, and less than fully participatory in CV1or U.S. sugar input demand. In CV3, PRAW has statistically significant α- and β-estimates that preclude PRAW's weak exogeneity to the system. See Juselius (2004).

absence of such policies. Future research needs to focus more on whether using this widely monitored USDA price indeed renders valid own-price elasticities of U.S. raw sugar demand. Second, the –5.0 own-price elasticity estimate and our comparative analysis of our logged variables' SEs suggests that U.S. sugar demand has likely responded with far more elasticity than previously thought in those cases when policy-encumbered own-price did manage to move appreciably. Third, policy formulators should note that demand for sugar is equally dictated by price-relevant conditions in major markets for sugar-using industrial markets as by own-price, and that efforts to formulate effective sugar policy need to carefully account for price-influencing forces in sugar-using product markets that may inadvertently inflate or cancel-out goals of proposed sugar policies in the U.S. raw sugar market. And fourth, policy makers need to note that events in non-sugar markets—say the 2001–2002 increases in various confectionary input prices—importantly influenced raw sugar price.

Cointegrating Relation 2: A Price Transmission among Sugar-Based Products

The second cointegrating relationship (Equation 12) appears to be a generally positive long-run price transmission between PCHOC normalized on the left side and PNOCHOC, PSOFT, and PBAKERY. Recent research has uncovered repeated instances of similar price-transmission relationships in other markets.[25] Equation 12 suggests a positive relationship among these prices.[26]

Theory suggests two potential explanations for this reduced-form relationship based on both demand and supply: common influences from consumer-based substitutability among the three product groups, and common supply-side production-cost influences from a shared dependence on sugar as a production-input cost. A rise in PNOCHOC may induce consumers to switch from non-chocolate confectionary products, and may therefore elicit a PCHOC increase in CV2. CV2 may also arise from common patterns of sugar-based production costs. Should sugar prices fall (or rise), production costs of the confectionary and soft drink industries would fall (or rise). Relative absolute β-values suggest that PCHOC and PNOCHOC sensitively react to each other, while PSOFT has more-subdued confectionary price effects, probably due to its aggregation across prices of sugar- and HFCS-sweetened products.

CV2's α-estimates (alpha2) are consistent with the β-related results in Equation 12 above. The α-estimates on PCHOC, PNOCHOC, and PSOFT are statistically significant and positive, suggesting common directions of error-correcting adjustment toward the attractor set from market shocks The relative speeds of such adjustment are reflected by the relative absolute α-values and suggest that soft drink prices adjust more sluggishly than confectionary prices, perhaps again due to PSOFT's aggregation across both sugar- and HFCS-based products. CV2's positive and significant β(DS0101) suggests that the 2001–2002 increases in confectionary input costs importantly influenced PCHOC over time.[27]

Cointegrating Relation 3: A Production-Cost Link among Sugar and Sugar-Using Prices

The third CV negatively relates raw sugar price to non-chocolate confectionary and soft drink prices, also in a reduced-form context with both elements of supply and demand. CV3's explanation may be tied to a common input-demand relationship with sugar as a productive input. Should PNOCHOC rise (fall), consumers may demand less (more) of such products, leading to a drop (rise) in production, a drop(rise) in the use of sugar as an input, and hence a fall (rise) in PRAW. This relationship among raw

[25] See Bessler, Yang, and Wongcharupan (2002) for cointegration relationships among different world wheat prices; Goodwin, McKenzie, and Djunaidi (2003) for relationships among U.S. chicken-product prices; Babula, Bessler, and Payne (2004) and Babula and Rich (2001) for relationships among U.S. wheat-based product prices.

[26] Coefficients on PBAKERY in this and the third relation suggest opposing patterns of influence and are discussed below separately.

[27] The negative and significant beta on the shift dummy DS0212, restricted into the cointegration space to account for the late-2002 run-up in prices of U.S. non-chocolate confectionary products, is not easily explained. However, DS0112 and DS0101 do overlap: DS0101 was designed to capture both the early 2001 rise in cocoa prices and the late-2002 rise in other confectionary input prices, while DS0212 was designed to capture only the latter late-2002 event to service the non-cocoa confectionary price. Collinearity between DS0101and DS0212 may explain β(DS0212)'s unexpectedly negative sign in equation 13.

sugar and processed sugar-based final-product prices is somewhat reminiscent of Moss and Schmitz's (2002, pp. 1279–81) finding of a stable, long-run or cointegrating relationship among monthly U.S. raw and refined sugar prices, though this comparison of results should be done cautionsly.

The adjustment coefficients (alpha3) appear to reinforce CV3's relationship. Its α(PRAW) is significant and negative, while its α(PSOFT) and and α(PNOCHOC) are positive and significant (though α(PNOCHOC) marginally at the 10-percent level). These α-estimates suggest that for a PRAW above equilibrium, error correction would elicit a fall in PRAW and a rise in PNOCHOC and PSOFT. That β(PSOFT) is nearly three times β(PNOCHOC)'s value suggests soft drink price responds more swiftly, which coincides with Williams and Bessler (1997) and Moss and Schmitz's (2002, 2004) analysis that over the long run, HFCS and sugar compete as sweeteners, although with the then-current grid of U.S. sugar policies, HFCS has held sway with major soda drink producers since 1996.

The strongly significant and negative β(DS9907) estimate highlights the previously noted importance that policy interventions have had on PRAW: that PRAW is lower than it would be without these policies. Perhaps the two years of payment-in-kind sugar policy benefits placed enough raw sugar on the market undermined U.S. sugar-price support policies. CV3's significant β(DS0212) estimate suggests the importance for sugar policymakers to account for relevant cross-market input market effects.

Comments on the PBAKERY's Role in the Two Price Transmissions

PBAKERY appears to engage in the error-correction process in Equations 12 and 13 differently than other downstream prices. Equation 12 positively relates PCHOC, PNOCHOC, and PSOFT, while PBAKERY is negatively related to these three prices. Equation 13 negatively relates PRAW to PNOCHOC and PSOFT, while the PRAW relationship with PBAKERY is a positive one. The significant β(BAKERY) and insignificant α(PBAKERY) estimates in CV2 and CV3 suggests that PBAKERY influences but does not respond to both error-correcting relations, suggesting a one-way influence. Perhaps PBAKERY serves as a closely-monitored informational variable of market conditions for the confectionary, soft drink, and sugar-related industries of a similar sort that recent research has uncovered for bread price within a system of U.S. wheat-based markets (Babula, Bessler, and Payne 2004). Exact structural interpretations for PBAKERY's role are likely clouded by the reduced-form nature of CV2 and CV3, and their likely combinations of two or more irreducible economic relations. Better interpretations are left to future research with expanded information sets.

Result-Suggested Market Implications of Recent Influential and Disruptive Events

Equations 11–13 suggest several upstream/downstream effects on the modeled U.S. sugar-related markets from the array of recent/current events introduced at this paper's onset:

- Under the sample's grid of current U.S. sugar policies, U.S. sugar input demand appears elastic and PRAW appears un-naturally sluggish. The final net effect from the 2005 hurricane-induced raw sugar supply declines and the USDA's immediately offsetting increase in the in-quota sugar-import allotment under the TRQ will likely have little or no effects on PRAW and QSUGAR.

- Non-sugar industrial input market events such as the 2001–2002 increases in confectionary input costs highly influenced the markets: they augmented QSUGAR and PRAW, and appeared to reduce some prices of sugar-based industrial products.

- The four U.S. sugar-policy interventions collectively had clear and important effects: demand and prices for sugar are lower than they would be without the policies. The policies seem to be disrupting the widely used and monitored USDA raw sugar price as own price.

- In CV3, relative α-values suggest that PSOFT responds more quickly than PNOCHOC to PRAW shocks. This latter result coincides with previous research suggesting that over time, soft drink producers consider sugar and HFCS as highly substitutable sweeteners (see Williams and Bessler 1997; Moss and Schmitz 2002, 2004).

Summary and Conclusions

Our primary goal is methodological: for perhaps the first time, to apply cointegrated VAR methods to test if the five U.S. sugar-related market systems comprise a cointegrated system. Ours may be the first monthly cointegrated VEC model of the U.S. markets for raw sugar and for confectionary, soft drink, and bakery products, and the model that emerged displayed notable statistical strength. Given the repeated and influential effects of U.S. sugar-policy changes and relevant events, an array of long-run and short-term deterministic components were worked into the fully restricted cointegrated VEC.

To our knowledge, some of our specific methods provided by Juselius (2004) may be first-time applications in the refereed agricultural economics literature: reduced-rank determination based on multiple sources of evidence; a thorough use of a battery of statistical-adequacy diagnostics (i.e., Doornik-Hansen values) to achieve a statistically adequate model; use of the Bonferoni criterion on standardized residuals in specifying outlier binaries; use of a systems-based test for series stationarity based on chosen reduced rank; use of the "known beta" test for parameter constancy; and application of Nielsen's test for the presence of I(2) trends.

Three error-correcting CVs emerged. The first focuses on the upstream raw-sugar market endowed with the most adequate information set and appears structurally interpretable as the beginnings of a sugar-demand curve. The second and third comprise price-transmission relationships that focus primarily on the downstream sugar-based markets with less adequate price-only endogenous vectors devoid of market quantities. CV2 and CV3 appear to have emerged as reduced-form price transmissions and are possibly each a combination of two or more irreducible economic relationships that may not be identified until future research uncovers more adequate and expanded downstream-market information sets (see Juselius and Toro 2005, p. 511).

U.S. sugar demand as an input appears tied to a linear combination of the raw sugar price and the chocolate-based confectionary price. Own-price appears to be an important factor, although evidence suggests that it is not fully responsive, may display an unnaturally sluggish behavior due to perhaps a series of exogenous U.S. sugar-policy interventions since the late 1990s, and may not be serving as an effective own-price.

A second surprising result was that U.S. sugar demand, perhaps due to own-price's less than fully endogenous participation, is equally determined by price-impacting events in major markets for sugar-using industrial products: evidence strongly suggests that PCHOC seems to act as a "co-own-price." In fact, results suggests that PCHOC more fully and endogenously participates in the CV1 demand relation than does the USDA's widely monitored own-price. A question arose as to whether this possibly policy-encumbered PRAW can be used to empirically estimate a valid own-price elasticity estimate. Policy makers should note that market conditions for sugar-using products in PCHOC must be considered when formulating sugar policy, to avoid unintended effects that augment or offset intended sugar-policy goals.

The cointegration results suggest that it is very important that researchers, agribusiness agents, and policy makers be aware of cross-market effects in input markets. For example, surges in cocoa and other input prices in 2001–2002 had important effects on sugar demand, raw price, and sugar-using product prices. Results also suggest that interested agents must be cognizant of the influential effects, whether intended or not, of U.S. policy mixes, e.g. having PIK concurrent with price-support policies.

The model results have some implications for market effects on U.S. sugar-related markets of the array of important/disruptive events introduced at this study's onset. Considering all events, including the USDA's increase in TRQ access for sugar imports, the 2005 hurricane-induced raw sugar shortfalls are likely not to have large increases on the U.S. price of raw sugar and sugar-using products. Cross-market impacts for confectionary inputs have importantly influenced U.S. markets for sugar and sugar-using products. The array of U.S. sugar policy interventions since the late 1990s, especially when Figure 2 is considered, seem to have reduced price and demand for sugar. And under certain conditions, results suggest that any FTA-induced U.S. access to sugar imports will likely have moderate impacts on U.S. sugar-using manufacturers through reduced production costs.

A number of recommendations for future research emerge. Future efforts need to delineate more precisely the role that price-influencing conditions in markets for sugar-related manufactured products have on U.S. sugar demand. U.S. sugar-policy makers should better analyze policy mixes for the optimal mix and need to thoroughly account for major events occurring in sugar-using input and output markets. More inquiry is needed on the effects of each of the four U.S. sugar policies captured collectively by DS9907, when expanded samples are available. While seemingly "far-flung" perhaps to some, sugar-policy makers must be of aware of the important effects that such events as the Ivory Coast civil war have had on U.S. sugar-related markets. More conclusive research is needed to determine whether the widely used and monitored USDA raw sugar price is indeed unnaturally sluggish due to exogenous policy interventions, if it can render valid sugar-demand parameter estimates with the current grid of policies, and what would happen to our own-price elasticity of demand estimate and those of others in the absence of such interventions. More research is needed on expanding the structural interpretability of the price-transmission relationships (CV2, CV3) as information sets and samples expand in the future. More research on the exact role of PBAKERY in the modeled system is needed.

References

Babula, R., D. Bessler, and W. Payne. 2004. "Dynamic Relationships Among U.S. Wheat-related Markets: Applying Directed Acyclic Graphs to a Time Series Model." *Journal of Agricultural and Applied Economics* 36:1–22.

Babula, R., D. Bessler, J. Reeder, and A. Somwaru. 2004. "Modeling U.S. Soy-Based Markets with Directed Acyclic Graphs and Bernanke Structural VAR Methods: The Impacts of High Soy Meal and Soybean Prices. *Journal of Food Distribution Research* 35:29–52.

Babula, R. and K. Rich. 2001. "A Time Series Analysis of the U.S. Durum Wheat and Pasta Markets." *Journal of Food Distribution Research* 32(2):1–19.

Bessler, D. 1984. "An Analysis of Dynamic Economic Relationships: An Application to the U.S. Hog Market." *Canadian Journal of Agricultural Economics* 32:109–24.

Bessler, D., J. Yang, and M. Wongcharupan. 2002. "Price Dynamics in the International Wheat Market: Modelling with Error Correction and Directed Acyclic Graphs." *Journal of Regional Science* 42:793–825.

Engle, R. and C. W. J. Granger. 1987. "Cointegration and Error Correction: Representation, Estimation, and Testing." *Econometrica* 55: 251–256.

Estima. 2004. *Regression Analysis of Time Series (RATS), Version 6*. Evanston, IL: Estima.

Goodwin, H. S., Jr., A. McKenzie, and H. Djunaidi. 2003. "Which Broiler Part is the Best Part?" *Journal of Agricultural and Applied Economics* 35:483–95.

Granger, C. W. J. and P. Newbold. 1986. *Forecasting Economic Time Series*. New York: Academic Press.

Hendry, D. F. 1986. "Econometric Modelling with Cointegrated Variables: An Overview." *Oxford Bulletin of Economics and Statistics* 48: 201–212.

Johansen, S. 1988. "Statistical Analysis of Cointegration Vectors." *Journal of Economic and Dynamic Control* 12:231–253.

Johansen, S. and K. Juselius. 1990. "Maximum Likelihood and Inferences on Cointegration: With Applications to the Demand for Money." *Oxford Bulletin of Economics and Statistics* 52: 169–210.

Johansen, S. and K. Juselius. 1992. "Testing Structural Hypotheses in a Multivariate Cointegration Analysis of the PPP and UIP for U.K." *Journal of Econometrics* 53: 211–244.

Juselius, K. 2004. *The Cointegrated VAR Approach: Methodology and Applications*, draft. Economics Institute, University of Copenhagen.

Juselius, K. and J. Toro. 2005. "Monetary Transmission Mechanisms in Spain: The Effect of Monetization, Financial Deregulation, and the EMS." *Journal of International Money and Finance* 24:509–31.

Moss, C. B. and A. Schmitz. 2004. "Delineating the Relevant U.S. Sweetener Markets." *Journal of Agricultural and Food Industrial Organization* 2(1):article 1.

Moss, C. B. and A. Schmitz. 2002. "Price Behavior in the U.S. Sweetener Market: A Cointegration Approach." *Applied Economics* 34:1273–1281. [Note that the journal mis-spelled Dr. Schmitz's name as Dr. Schmits].

Nielsen, H. B. 2002. "An I(2) Cointegration Analysis of Price and Quantity Formation in Danish Manufactured Exports." *Oxford Bulletin of Economics and Statistics* 65:449–472.

Tiao, G. and G. Box. 1978. "Modelling Multiple

Time Series: With Applications." *Journal of the American Statistical Association* 76: 802–16.

United States Department of Agriculture Economic Research Service (USDA ERS). 1996. *Sugar and Sweetener Situation and Outlook Yearbook,* Publication No SSS-245, January 31.

United States Department of Agriculture Economic Research Service (USDA ERS). 2004a. *Sugar and Sweeteners Outlook,* Publication No. SSS-240, May.

United States Department of Agriculture Economic Research Service (USDA ERS). 1993–2004b. *Sugar and Sweetener Situation and Outlook Yearbook.* various annual issues.

United States Department of Labor, Bureau of Labor Statistics. 2004. Producer Price Index Databases." www.bls.gov, accessed July 2004.

United States International Trade Commission (USITC). 2003. Durum and Hard Red Spring Wheat from Canada, Investigations No.s 701-TA-430A and 731-TA-1019A and 1019B (Final). Washington DC: USITC, pub. No. 3639, October.

U.S.–Thailand Free Trade Agreement Business Coalition. 2004. "Written Testimony" delivered to the public records of Investigation Numbers TA-131-29 and TA-2104-12: U.S.-Thailand Free Trade Agreement Advice Concerning the Probable Economic Effect of Duty-Free Treatment for Imports, May 14. www.us-asean.org/us-thai-fta, accessed December 3, 2004.

Williams, O. and D. A. Bessler. 1997. "Cointegration: Implications for the Market Efficiencies of the High Fructose Corn Syrup and Refined Sugar Markets." *Applied Economics* 29:225–32.