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30. April 2008

Online at <http://mpra.ub.uni-muenchen.de/9684/>

MPRA Paper No. 9684, posted 22. July 2008 / 22:08

The Commodity Terms of Trade, Unit Roots, and Nonlinear Alternatives: A Smooth Transition Approach*

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This Draft: April 30, 2008

Abstract

Market price dynamics for North American oriented strand board markets are examined. Specifically, the role of transactions costs are examined vis-à-vis the law of one price. Weekly data for the January 3rd, 1995 through April 14th, 2006 period are used in the analysis. Nonlinearities induced by unobservable transactions costs are modelled by estimating smooth transition autoregressions (STARs). Results indicate that nonlinearity is an important feature of these markets and that the parity relationships implied by economic theory are generally supported by the STAR models. Implications for the efficiency of spatial market linkages are examined by estimating generalized impulse response functions.

Keywords: Law of one price, Oriented strand board, Nonlinear model; Smooth transition autoregression; Unit root tests

JEL Classification Codes: E30; C22; C52; Q23

*This research has been supported by USDA-USFS Cooperative Agreement No. xxxxxx. Senior authorship is not assigned.

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1 Introduction

Over the years there has been considerable debate about the veracity of the Law of One Price (LOP) as it pertains to markets for tradeable goods. On one hand, economists take it as being nearly axiomatic that freely functioning markets for traded, homogeneous products should ensure that prices are efficiently linked across markets and thus that no persistent opportunities for spatial arbitrage profits exist. This general concept is so fundamental that it often serves as an untested assumption that forms the starting point in models of exchange rate determination. In spite of the prominent role played by the LOP, a number of qualifications to this general relationship exist. In particular, the strength of spatial linkages among commodity markets typically depends on the availability of reasonably accurate market information and the lack of significant impediments to spatial trade. Linkages that involve national borders also raise a number of issues pertaining to exchange rate pass-through and international trade policies. It is therefore common to examine commodities whose prices are denominated in a common currency, such as is the case with regional trade within national borders. An entire literature devoted to considerations of spatial market linkages within a country—the case of spatial market integration—has paralleled the Law of One Price studies.¹

The general implication underlying these concepts is that prices for homogeneous products at different geographic locations should differ by little more than transport costs and any other costs that might be related to market arbitrage. On the other hand, there is substantial evidence that the adjustment lags required to restore arbitrage equilibrium are often found to be far longer than would seem natural based upon any reasonable understanding of the mechanics of physical trade as pertains to the markets in question. Indeed, in some instances price differences have been observed to exhibit near unit root behavior, a condition that would most certainly be at odds with any typical rendering of the LOP.

¹Distinctions between tests of the “LOP” and “spatial market integration” are not especially meaningful. In both cases, the economic phenomena being evaluated—spatial market arbitrage—is identical. A survey of both strands of literature is provided by Fackler and Goodwin (2001).

What then is the empirical consensus regarding the LOP in commodity and goods markets? In a rather extensive study that used disaggregate data for traded goods, Isard (1977) found rather conclusive evidence against the LOP. These early findings by Isard (1977) were subsequently confirmed for a variety of commodities in a wide array of market settings by Richardson (1978), Thursby, Johnson, and Grennes (1986), Benninga and Protopapadakis (1988), and Giovannini (1988), among others. Goodwin, Grennes, and Wohlgenant (1990) did, however, find some support for the LOP when it is specified in terms of price expectations as opposed to observed prices. A potential shortcoming associated with all of these empirical studies is a general failure to explicitly consider the role of transactions costs and potential delivery lags. More recently, cointegration techniques have been used to rationalize the LOP as a long-run equilibrium concept. By using this view of LOP, numerous authors have found rather more compelling evidence in favor of the law, including, for example, Buongiorno and Uusivuori (1992) (U.S. pulp and paper exports), Michael, Nobay, and Peel (1994) (international wheat prices), Bessler and Fuller (1993) (U.S. regional wheat markets), and Jung and Doroodian (1994) (softwood lumber markets).

Most recently, economists have sought to explore the implications of spatial arbitrage by using nonlinear models of various forms. The underlying notion is that adjustments to equilibrium may not be linear, and that this nonlinearity may, in turn, be associated with unobservable transactions costs associated with arbitrage. The theoretical underpinnings for nonlinearity in the LOP induced by transactions costs have been put forward by, among others, Dumas (1992), though the idea dates back at least to the work of Heckscher (1916), who noted that transactions costs may define “commodity points” within which prices are not directly linked because the price differences are less than the costs of trade.

Empirical investigations of the role of nonlinearity as pertains to the LOP have been reported by Goodwin and Piggott (2001), Lo and Zivot (2001), Sephton (2003), Balcombe, Bailey, and Brooks (2007), and Park, Mjelde, and Bessler (2007). The empirical work reported in these studies has been conducted primarily by using variants of discrete threshold cointegration

models of the sort introduced by Balke and Fomby (1997). In general these studies have found in favor of threshold effects, with the path of adjustment to equilibrium depending typically on the size and the sign of the shock. Specifically, the recognition of nonlinearities typically provides much greater support for spatial market linkages of the sort implied by theory. In particular, shocks that lead to profitable arbitrage opportunities (i.e., net of transactions costs) are quickly eliminated.

The extent to which spatially distinct markets are efficiently linked may have important implications for the overall market performance.² We consider regional U.S. markets for a very prominent traded commodity—oriented strand board (OSB). OSB now accounts for a larger share of the overall panel lumber products market, exceeding the production and use of plywood by a considerable extent. Spatial linkages in this market are of particular interest because it is a good that is widely traded across considerable distances within North America. Consumption is widespread and spatially dispersed while production tends to be concentrated in particular regions such as the Southeastern U.S. and Eastern Canada. Depletion of the old-growth timber stocks that traditionally served as a source for panel lumber products has brought about tremendous growth in the use of engineered lumber products such as OSB. A burgeoning housing market (and its recent contraction) have brought about a number of significant shocks to this rapidly expanding industry. These and related factors underscore the importance of understanding and quantitatively measuring linkages among regional OSB markets. Lumber products are also bulky commodities, which are costly to transport. This simple observation suggests that transactions costs should play an important role in shaping market linkages and reactions to market shocks for this particular good.

In this paper we apply a class of nonlinear, time-series models that allow for gradual ad-

²Many different conceptual definitions of the notion of “spatial efficiency” are used in the literature and as a result the coherence of conclusions based upon empirical tests is often strained. Here we are interested in the extent to which price shocks in one location provoke market reactions in another. This is consistent with conventional views of market efficiency—that is, an absence of persistent arbitrage opportunities. However, we do not explicitly pursue other aspects of efficiency that are often considered (especially in developing economy studies), such as the underlying structure and transportation linkages associated with regional commodity trade.

justments among price linkages. Specifically, we focus on a class of smooth transition models known as QSTAR, or Quadratic Smooth Transition Autoregressions. As the name implies, QSTAR models allow for a potentially smooth transition in and out of a transactions cost band. As noted by Taylor et al. (2000), this latter feature may be important if agents operating in OSB markets are heterogenous, and each faces a slightly different and unique set of trading costs. Furthermore, and as we discuss in more detail below, nests as a special case many of the more common threshold-type models used previously to examine the role of transactions costs vis-à-vis the LOP. We present an empirical application wherein we examine market linkages on a weekly basis for six price pairs derived from four regional OSB markets in North America. Preliminary test results indicate there is considerable evidence in every case in favor of nonlinearity. In every instance the estimated nonlinear models are superior to their linear counterparts and, moreover, the results show how ignoring transactions costs can lead to erroneous conclusions regarding the nature and the strength of OSB market linkages and the economic behavior underlying them.

The plan of the paper is as follows. In the next section we briefly discuss the North American OSB market and the data used in the empirical analysis. In Section 3 we discuss the econometric methods including the specification of STAR models as well as well as a testing strategy that will help guide the final model specifications. In Section 4 we present and evaluate the estimation results. Section 5 concludes.

2 The OSB Market and Price Data

In this study we focus on spatial price relationships for an increasingly important wood product: oriented strand board (OSB). OSB is a manufactured wood product that was first introduced in 1978 (the forerunner to oriented strand board was waferboard). OSB is engineered by using waterproof and heat cured resins and waxes, and consists of rectangular shaped wood strands that are arranged in oriented layers. OSB is produced in long, continuous mats which are

then cut into panels of varying sizes. In this regard OSB is similar to plywood, although OSB is generally considered to have more uniformity than plywood and is, moreover, cheaper to produce. The Structural Board Association (SBA) reports that in 1980 OSB panel production in North American was 751 million square feet (on a $3/8^{th}$'s inch basis), but that by as early as 2005 this number had grown to 25 billion square feet. The SBA also reports that by 2000 OSB production exceeded that of plywood, and that by 2006 OSB production enjoyed a sixty-percent market share among all panel products in North America. Figure 1 illustrates the substantial growth in OSB use and the corresponding decline in use of plywood products. OSB is widely used in residential and commercial construction in North America, with the bulk of OSB produced in this region coming from the Southern United States and Canada.

Considering the above, we focus on price relationships for OSB in four important regional North American markets. Specifically, the regions examined are: (1) Eastern Canada; (2) North Central U.S.; (3) Southeastern U.S.; and (4) Southwestern U.S. The result is there are six possible pairwise spatial price relationships that may be examined. The price data are for panels of $7/16^{th}$'s inch oriented strand board, and are expressed in U.S. dollars per thousand square feet. All price data are observed on a weekly basis and were obtained from the industry source Random Lengths. The period covered is from January 3, 1995 through April 14, 2006, the result being there are 589 usable observations. Several points about this period of time are notable. First, this period of time represents the growth and eventual dominance of OSB as a building material. The market was rapidly expanding and thus was experiencing considerable change that may have influenced spatial linkages. Second, this period also encompasses the significant strengthening of the U.S. housing market as well as the subsequent contractions which are still being experienced. Such events suggest the likelihood of significant price volatility and shocks to regional markets and thus raise interesting questions about how these shocks may have been transmitted among regional markets.

The basic unit of analysis used throughout this study is the log of the price ratio $\ln(p_{it}/p_{jt})$, where i and j are indices indicating regional location (i.e., $i, j = 1, \dots, 4$) and a subscripted

t is a time index such that $t = 1, \dots, T$, where $T = 589$. It is common to consider price relationships in such a “log-of-ratios” specification as it allows the results to be interpreted in proportional difference terms. Plots of the six price pairs, expressed in logarithmic form, are presented in Figure 2. The plots indicate that prices in Eastern Canada are generally lower than those in various regions of the United States, although there has been considerable variation over time. As well, the OSB price in Eastern Canada has been rising steadily relative to that of the Southwest throughout much of the sample period (Figure 2). Likewise, OSB prices in the Northeast U.S. are, on average, slightly lower than those in the rest of the country. Similar patterns are observed for the price relationship between the Southeast and the Southwest, although prices in the Southwest have been increasing relative to those in the Southeast in recent years. The figure also illustrates a considerable degree of volatility in the price ratios, suggesting the potential for significant market interactions and reactions to shocks.

3 Econometric Methods

3.1 STAR-Type Models

The fundamental building block of any nonlinear time series model, is of course, a linear model. Specifically, let $y_t = \ln(p_{it}/p_{jt})$ for some i and j . We may then specify a linear p^{th} -order autoregressive model for the price pair as follows

$$\Delta y_t = \boldsymbol{\phi}' \mathbf{x}_t + \varepsilon_t, \tag{1}$$

where $\boldsymbol{\phi} = (\phi_0, \phi_1, \dots, \phi_{p-1}, \theta)$ and where $\mathbf{x}_t = (1, \Delta y_{t-1}, \dots, \Delta y_{t-p+1}, y_{t-1})$.³ As well, ε_t is a mean-zero *iid* error term with finite variance. An appropriate lag length p may be chosen by using, for example, a model selection criterion such as Akaike’s Information Criterion (AIC),

³In the ensuing discussion we assume that y_t is stationary in the levels. As we shall see in subsequent sections this assumption is entirely consistent with our preliminary analysis of the OSB regional price data.

the Schwarz Information Criterion (SIC), or by using a general-to-specific modeling strategy.

In the STAR modeling framework, the linear autoregression in (1) is modified as follows

$$\Delta y_t = \phi_1' \mathbf{x}_t (1 - G(s_t; \gamma, \mathbf{c})) + \phi_2' \mathbf{x}_t G(s_t; \gamma, \mathbf{c}) + \varepsilon_t. \quad (2)$$

where $\mathbf{c} = (c_1, c_2)$. In (2) $G(s_t; \gamma, \mathbf{c})$ is the so-called “transition function” that varies in a potentially smooth manner between zero and one according to the transition variable s_t , and whose properties are determined, at least in part, by the values of the speed-of-adjustment parameter γ and the location parameter or parameters, $\mathbf{c}$. The transition variable s_t may be a function of of nearly any observed variable, but in practice it is taken typically to be a function of lagged the lagged dependent variable, y_t . For example, a commonly used transition variable is $s_t = \Delta_d y_{t-1} = y_{t-d} - y_{t-1-d}$, $d = 1, \dots, D_{max}$, where d is known as the delay parameter (see, e.g., Teräsvirta, 1994). This specification has a natural economic interpretation in terms of departures from equilibrium parity conditions. The transition variable determines the nature of adjustment (or the transition). In our case, the transition variable represents a lagged, difference in logarithmic prices at two different locations. The larger the (absolute) value of the transition variable, the bigger will be the difference in prices and thus the larger is the deviation from a presumed parity condition. We anticipate that larger deviations will provoke faster or larger market adjustments. Thus, this specification accommodates market adjustments that follow (in accordance with the delay parameter d) departures from spatial parity among prices. The overall implication is that market shocks that lead to departures from the LOP at time t should lead to adjustments that tend to restore the LOP at time $t + d$.

In any case, as $G(\cdot)$ varies between zero and one, the model’s autoregressive parameters reflect a weighted average of the parameters in ϕ_1 and ϕ_2 , with the weights being given by the value of the transition function. While there are several possible specifications for the transition function $G(\cdot)$ in (2), one that seems especially suitable for present purposes is the second-order or quadratic logistic function, proposed initially by Jansen and Teräsvirta

(1996), where $\mathbf{c} = (c_1, c_2)$, and given by

$$G(s_t; \gamma, \mathbf{c}) = [1 + \exp\{-\gamma(s_t - c_1)(s_t - c_2)\}]^{-1}, \gamma > 0, c_1 \leq c_2. \quad (3)$$

An interesting feature of (3) is that, when combined with (2), the parameters $\phi_1(1-G(s_t; \gamma, \mathbf{c})) + \phi_2 G(s_t; \gamma, \mathbf{c})$ change symmetrically as a function of s_t around $(c_1 + c_2)/2$, the quadratic logistic function's mid-point. In this manner the model that combines (2) and (3), the so-called “QSTAR,” allows for the possibility of a transactions cost band wherein the movement in or out of the band is potentially smooth. Note also that as $\gamma \rightarrow 0$ the model in (2) becomes linear and, as well, as $\gamma \rightarrow \infty$, and assuming that $c_1 \neq c_2$, then $G(\cdot)$ assumes a value of one for $s_t < c_1$ and $s_t > c_2$ and zero otherwise. In this manner the QSTAR nests as a special case a three-regime self exciting threshold autoregression (SETAR). This later property is useful in part because several previous studies of spatial price relationships have employed three-regime SETAR models to account for nonlinearities introduced by transactions costs. See, for example, Goodwin and Piggott (2001). Even so, to date the QSTAR has not been explicitly used to examine the LOP or purchasing power parity (PPP).⁴

There are intuitive reasons to suspect that the patterns of price adjustment in regional markets will be smooth rather than discrete, even though the economic behavior underlying the adjustments is of a discrete nature (i.e., arbitrage is either profitable or it is not). Weekly prices of the sort used in our analysis are comprised of many transactions among many agents, all averaged or otherwise aggregated to obtain a single price. To the extent that agents are not all identical and that markets are not entirely static within the week, differences will exist across individual transactions. Averaging or otherwise aggregating prices to obtain a weekly price quote smooths these differences and should result in smooth patterns of price adjustment. Thus, it is preferable to allow for such smooth adjustments to market shocks rather than imposing discrete breaks that define market regimes.

⁴Eklund (2003) considers QSTAR models in the context of PPP but does not actually estimate such models.

A STAR-type model more commonly used in testing the LOP and PPP is the exponential or ESTAR model. See, for example, Kilian and Taylor (2003), Taylor, Peel, and Sarno (2001), Paya and Peel (2004), and Fan and Wei (2006), among others. In this case the transition function embedded in (2) is specified as

$$G(s_t; \gamma, c) = 1 - \exp\{-\gamma(s_t - c)^2\}, \gamma > 0. \quad (4)$$

This transition function (4) includes only one location parameter, c . The combination of (2) and (4) yields the ESTAR model. In the case of the ESTAR the parameters $\phi_1(1 - G(s_t; \gamma, c)) + \phi_2 G(s_t; \gamma, c)$ change symmetrically about the parameter c in (4) with s_t , again allowing for the possibility of a transactions cost band. As with the QSTAR, the ESTAR also becomes linear as $\gamma \rightarrow 0$ but as well when $\gamma \rightarrow \infty$. As such, the ESTAR model is also capable of modeling a spatial market setting wherein there are transactions costs associated with altering trade flows. From a modeling perspective the primary advantage to the ESTAR vis-à-vis the QSTAR is there is one less parameter to estimate. Alternatively, the limitation is that the ESTAR does not, in fact, nest a TAR model, as is the case for the QSTAR. This latter feature of the ESTAR seems undesirable in as much as previous studies on spatial price relationships have successfully employed TAR specifications; it would be desirable to employ a modeling framework that nests the TAR. For these reasons we examine the role of transactions costs in the LOP for OSB in North American markets by utilizing the more flexible QSTAR specification. In Section 4 we discuss the empirical results and Section 5 concludes.

3.2 A Combined Unit Root and Linearity Testing Framework

A key component of any study of spatial price relationships is: (1) testing for unit roots (i.e., stochastic trends) in the respective $\ln(p_{it}/p_{jt})$ price pairs, and (2) testing for linearity against nonlinear alternatives. Of course, the classic (Augmented) Dickey-Fuller (1979, 1981) unit root test, or ADF test, examines the null of a linear model with a unit root against an alternative

model that imposes stationarity in the levels of the price pairs.⁵ The Dickey–Fuller approach does not, however, allow for nonlinear models under the alternative. For this reason prior research in this area has generally approached testing in a two–step process. First, stationarity is examined by using standard unit root tests such as the ADF test. Second, the data are then examined for nonlinear features by specifying an appropriate linear model and testing this model against various nonlinear alternatives. See, for example, Skalin and Teräsvirta (2002) study of unemployment data for several countries.

It may be desirable to test for the unit root hypothesis and linearity simultaneously, that is, to test the null of a linear model with a unit root against a model that incorporates nonlinear features along with mean reversion. In recent years several studies have examined testing issues of this sort, most typically in the context of TAR alternatives. See, for example, Enders and Granger (1998) and Caner and Hansen (2001). As well, Eklund (2003) presents a bootstrapping framework for testing a linear unit root model against a stationary STAR–type alternative, specifically, a QSTAR alternative. Rothe and Sibbertsen (2006) extend the Phillips–Perron type test of unit roots to consideration of a stationary ESTAR alternative. Of interest is that both Eklund (2003) and Rothe and Sibbertsen (2006) were motivated by questions about model specification in the context of purchasing power parity.

In this application, we adopt a testing approach that is motivated largely by Eklund’s (2003) proposed bootstrapping framework. To formalize the testing framework, consider that we wish to test the validity of the linear unit root model

$$\Delta y_t = \mu + \tilde{\phi}' \mathbf{x}_t + \varepsilon_t \tag{5}$$

⁵Given that log relative prices are the basic unit of analysis, it follows that in this case the ADF test is equivalent to a test that $\beta = (0, 1)$ is a cointegrating vector in the log–linear price relationship $\ln(p_{it}) - \beta_0 - \beta_1 \ln(p_{jt})$.

against the stationary STAR alternative

$$\Delta y_t = (\mu_1 + \theta_1 y_{t-1} + \tilde{\boldsymbol{\phi}}_1' \mathbf{x}_t)(1 - G(s_t; \gamma, \mathbf{c})) + (\mu_2 + \theta_2 y_{t-1} + \tilde{\boldsymbol{\phi}}_2' \mathbf{x}_t)G(s_t; \gamma, \mathbf{c}) + \varepsilon_t, \quad (6)$$

where now $\mathbf{x}_t = (\Delta y_{t-1}, \dots, \Delta y_{t-p+1})$ and where $G(\cdot)$ is given by the quadratic logistic function in (3). As noted previously by Luukkonen, Saikkonen, and Teräsvirta (1988) and Teräsvirta (1994), even in the absence of the unit root question it is not possible to directly test (5) against (6) by using any standard testing framework. The reason is that there are two ways to obtain the linear model in (5) from the STAR model in (6). To illustrate, assume for the moment that $\theta_1 = \theta_2 = 0$; that is, the model contains a global unit root. Then (5) may be derived from (6) either when $H_0 : \mu_1 = \mu_2, \tilde{\boldsymbol{\phi}}_1 = \tilde{\boldsymbol{\phi}}_2$ or when $H'_0 : \gamma = 0$. In the second case, that is, under H'_0 , the location parameters c_1 and c_2 as well as the parameter vectors $\tilde{\boldsymbol{\phi}}_1$ and $\tilde{\boldsymbol{\phi}}_2$ are unidentified.⁶ In other words, directly testing (6) against (5) results in unidentified nuisance parameters under the null, the classical Davies (1977, 1987) problem. What Davies (1977, 1987) shows is that in this case standard tests such as the log likelihood ratio test do not have known limiting distributions; test results must generally be obtained numerically via simulation. See, for example, Hansen (1996).

Alternatively, Luukkonen, Saikkonen, and Teräsvirta (1988), or LST, discuss an approach for testing H'_0 that does not involve simulation. Specifically, they suggest replacing the transition function $G(\cdot)$ in (6) with a suitable Taylor series expansion in s_t , where the expansion is evaluated at $\gamma = 0$. In the case of the ESTAR or QSTAR models, LST suggest using a first-order approximation to $G(\cdot)$, which yields, after substitution into (6) and collecting terms, the second-order auxiliary regression

$$\Delta y_t = \delta_0 + \lambda_0 y_{t-1} + \boldsymbol{\vartheta}' \mathbf{x}_t + \sum_{i=1}^2 \delta_i s_t^i + \sum_{i=1}^2 \lambda_i y_{t-1} s_t^i + \sum_{i=1}^2 \boldsymbol{\vartheta}_i' \mathbf{x}_t s_t^i + \xi_t. \quad (7)$$

⁶Alternatively, if the null hypothesis $H_0 : \mu_1 = \mu_2, \tilde{\boldsymbol{\phi}}_1 = \tilde{\boldsymbol{\phi}}_2$ is used the parameters γ and c_1 and c_2 are unidentified.

See van Dijk, Teräsvirta, and Franses (2002) for additional details. Escribano and Jordá (1999) present evidence that a second-order approximation to $G(\cdot)$ may have some advantages in testing in terms of identifying the transition function's inflection points when ESTAR or QSTAR models are being examined. In this case the following auxiliary regression obtains

$$\Delta y_t = \delta_0 + \lambda_0 y_{t-1} + \boldsymbol{\vartheta}' \mathbf{x}_t + \sum_{i=1}^4 \delta_i s_t^i + \sum_{i=1}^4 \lambda_i y_{t-1} s_t^i + \sum_{i=1}^4 \boldsymbol{\vartheta}_i' \mathbf{x}_t s_t^i + \xi_t. \quad (8)$$

Note that in either case that the error term ξ_t in the auxiliary regression equations is a linear function of the original error term ε_t in (6) plus approximation error. Even so, under the null hypothesis of linearity the approximation error is zero and $\xi_t = \varepsilon_t$.

In what follows we use (8) to test the null hypothesis that the log relative price pairs for regional OSB markets are best characterized by a linear model containing a unit root. Specifically, by imposing the restrictions $\delta_0 = \dots = \delta_4 = \lambda_0 = \dots = \lambda_4 = \vartheta_{1,1} = \dots = \vartheta_{4,p} = 0$ on (8), a linear model containing a single unit root obtains. We refer to the null hypothesis being tested in this case as H_0^{lur} . Of course it is then possible to conduct an F test of H_0^{lur} . We refer to the resulting F statistic as F_{lur} , which will be associated with $(9+3p)$ and $T-(10+5p)$ degrees of freedom.

Of course the F_{lur} test statistic will not, in fact, be associated with a limiting F-distribution, at least to the extent that the unit root hypothesis is also embedded in the test. Here we choose to simply approximate the limiting distribution of the F_{lur} test statistic empirically by using nonparametric bootstrapping methods in a manner similar to that discussed by Li and Maddala (1996), Enders and Granger (1998), and MacKinnon (2002). In fact, Leybourne, Newbold, and Vougas (1998) and Sollis (2005) have employed similar methods in testing for unit roots in the residuals of time-varying autoregressions, a family of models similar in many respects to STAR models. As well, and as previously noted, Eklund (2003) has proposed a similar albeit more restrictive framework for testing for unit roots against QSTAR alternatives in the context of purchasing power parity.

The testing procedure is implemented as follows. We use sample data to estimate the model under both the null H_0^{lur} in (5) and the (modified) alternative, (8). The F_{lur} test statistic is then constructed in the usual manner. We then use the residuals from the model estimated under the null to perform B (dynamic) bootstrap simulations of size T , where B is some suitably large number. In so doing we obtain B new (pseudo) sets of T observations for the dependent variable, y_t . See Li and Maddala (1986) for details regarding the dynamic bootstrap. Both the null and the alternative models are then re-estimated by using the simulated data and for each of the B simulated data sets the test statistic, which we refer to as $\tilde{F}_{lur}$, is constructed. In this manner it is possible to use the set of B estimates $\tilde{F}_{lur}$ to obtain the empirical distribution for the test statistic F_{lur} . Moreover, this empirical distribution may then be used to derive an empirical p -value corresponding to the original test statistic F_{lur} . This is the procedure performed here to test the linear unit root specification against a nonlinear model that, moreover, may include (possibly local) mean reversion.

A final issue of concern is the choice of the transition variable, s_t and, accordingly, the delay parameter, d . For the OSB price data we consider two sets of transition variables. The first is defined simply as the set $s_t = y_{t-d} - y_{t-d-1}$ for $d = 1, \dots, 8$, and the second as $s_t = y_{t-1} - y_{t-d-1}$, again for $d = 1, \dots, 8$. Similar candidates for the transition variable have been considered, by, for example, Persson and Teräsvirta (2003). As a practical matter, the optimal d may be determined by choosing the value that corresponds with the strongest rejection of the null hypothesis (Teräsvirta, 1994). Recall that in this case the null hypothesis H_0^{lur} may be rejected because (1) the data are in fact stationary in the levels; (2) the preferred model specification is nonlinear, presumably due to transactions cost bands; or (3) for both reasons.

4 Empirical Results

The econometric results obtained for the weekly OSB regional North American price data for the 1995–2006 period are presented in three parts. First, we discuss the results of linear unit

root (i.e., ADF) tests applied to the price pairs. We then discuss the results for the tests of the linear unit root model versus a nonlinear and (possibly locally) stationary alternative. Finally, we present the results of estimated QSTAR models for the six price pairs.

4.1 ADF Test Results

Three ADF tests were performed for each logarithmic ratio of price pairs, and include the case where (1) the null model is a random walk (the τ test); (2) the null model is a random walk with drift (the τ_μ test); and (3) the null model model is a random walk with drift and a linear trend (the τ_τ test). In all cases we retained the first sixteen observations for which to determine the optimal lag length, determined, moreover, by choosing the model with the lowest AIC, and for conducting further diagnostic analyses. The result is that after first-differencing we have a usable sample of 568 weekly observations for each market pair.

The ADF test results, including empirical p -values obtained in each case by bootstrapping the respective null model $B = 999$ times, are reported in Table 1. Overall there is little support for the unit root hypothesis in these data, at least when tested against linear alternatives. Among other things, the implication is that $\ln(p_{it}) - \ln(p_{jt})$ may be thought of as a cointegrating relationship for all i and j .

Results of testing the unit root hypothesis against stationary but nonlinear alternatives are presented in Table 2 for each market pair. Candidate transition variables include lagged first differences, $s_t = \Delta y_{t-j}$, $j = 1, \dots, 12$, and increasing differences, $s_t = y_{t-1} - y_{t-j}$, $j = 2, \dots, 13$, where Δ denotes a first difference operator such that $\Delta z_{t-j} = z_{t-j} - z_{t-j-1}$. As before, empirical p -values are constructed by using 999 non-parametric bootstrap draws of the null model's residuals. Results in Table 2 indicated that the null of a unit root and linearity may be rejected in every case at conventional significance levels. Indeed, for most price pairs the minimal p -value (0.001) is obtained for more than one candidate transition variable, the implication being that the decision about which transition variable is ultimately most appropriate must be made during the model estimation and validation stage of the analysis. Given the results in Table

1, that is, the results of standard Dickey–Fuller tests, it seems reasonable to conclude that the results in Table 2 provide strong evidence in favor of nonlinearity in the data generating processes for each price pair, and most notably nonlinearity that is potentially consistent with a QSTAR specification.

As specified, the QSTAR model given by (2) and (3) is nonlinear in parameters. Thus, nonlinear estimation methods must be employed. Additional details regarding our approach to estimation may be found in van Dijk, Teräsvirta, and Franses (2002). In what follows we retain the first seventeen observations to initialize the lags and, as well, to test for any remaining nonlinearity. The result is there are 568 sample observations used in estimation. As before, optimal lag lengths are determined by applying the AIC to the linear model. Finally, as indicated in Table 2 there are several instances in which more than one transition variable seems appropriate. The final choice of the transition variable is determined by carefully examining the results of several model fit statistics as well as various diagnostic tests. These are reported in Table 3 for the final model specifications. Plots of the estimated transition functions are reported in Figure 3.

As recorded in Table 3, for every price pair the estimated standard error for the QSTAR model is smaller than that for the respective linear model. Although not reported in order to conserve space, in every case the respective QSTAR model also has a lower AIC than its linear counterpart. In this regard the estimated QSTAR models represent an improvement in fit relative to their respective linear analogues. Results in Table 3 also suggest that there is no evidence of skewness in the estimated residuals of the QSTAR models, but in each case there is substantial evidence of excess kurtosis. Of course such results are not surprising given that weekly price data are being employed. For these reasons the Lomnicki–Jarque–Bera (LJB) test overwhelmingly rejects the null hypothesis of normality of the estimated residuals in each case (Table 3). Correspondingly, diagnostic test results also reveal there is considerable evidence of ARCH-type heteroskedasticity in each case.

In addition to the foregoing tests, the diagnostic tests developed by Eitrheim and Teräsvirta

(1996) were also employed to test for any remaining autocorrelation as well as remaining non-linearity and/or parameter nonconstancy. These tests are implemented as F -test versions of Lagrange Multiplier (LM) tests. The results, also reported in Table 3, suggest that there is little evidence of remaining autocorrelation and, as well, there is virtually no evidence of remaining nonlinearity or parameter nonconstancy in the estimated QSTAR models. Finally, the plots in Figure 3 indicate that in every case except that involving the relationship between OSB prices in the Southeastern and Southwestern U.S., that the estimated QSTAR models approach a SETAR.

4.2 Model Dynamics

While the model fit and diagnostic results indicate the estimated QSTAR models for the regional OSB price relationships appear to do an adequate job of explaining the data, it remains, of course, to examine the implications for the dynamics of price linkages in each case. To obtain additional insights into the behavior of the estimated QSTAR models, we performed stochastic, forward simulations of the models by using a bootstrap routine similar to that suggested by Clements and Smith (1997) for use in SETAR models. In essence bootstrapping is used to perform numerical integration in order to obtain estimates of expected values of forward iterations once stochastic shocks are introduced. The stochastic forward iterations of each model, along with approximate 95-percent confidence bands, are reported in Figure 4. In every case the mean forward iterations converge to a stable path. As well, the 95-percent confidence bands show that, in every case, the simulations generally encompass the range of the observed data. Based on these results it seems reasonable to further assess the dynamics properties of the estimated models by developing generalized impulse response functions.

Generalized impulse response functions (GIRFs) are also obtained by simulating the model ahead both with and without a shock and for different histories. The basic methodology for obtaining GIRFs is developed in some detail by Koop, Pesaran, and Potter (1996), and proceeds

as follows. Let δ denote a specific shock, that is, let $\delta = \varepsilon_t$ for the initial period $t = 0$. As well, let a given history of data associated with time t be given by $\Omega_{t-1} = \omega_{t-1}$. The GIRF for time period (i.e., history) t at forward iteration n is then given by

$$GIRF_{\Delta y}(n, \delta, \omega_{t-1}) = E[\Delta y_{t+n} | \varepsilon_t = \delta, \Omega_{t-1} = \omega_{t-1}] - E[\Delta y_{t+n} | \varepsilon_t = 0, \Omega_{t-1} = \omega_{t-1}]. \quad (9)$$

where the histories are determined as follows. We randomly draw (with replacement) 142 histories from the 569 available histories, that is, we use approximately one quarter of the available histories for each price series. We then use normalized shocks taking on a range of values given by $\delta/\sigma_\varepsilon = 3.0, 2.8, \dots, 0.2$, where σ_ε denotes the estimated standard deviation of the residuals from the respective QSTAR model. For each combination of history and initial shock we compute $GIRF_{\Delta y}(n, \delta, \omega_{t-1})$ for $n = 0, 1, \dots, 78$, or 1.5 years. The expectations in (9) are computed, both with and without the shock, by using 800 bootstrap draws of the model's estimated residuals. Impulse responses for the log levels of the price pairs are obtained by simply summing those obtained for the first differences, that is, by

$$GIRF(n, \delta, \omega_{t-1}) = \sum_{i=0}^n GIRF_{\Delta y}(i, \delta, \omega_{t-1}). \quad (10)$$

It is also possible to obtain GIRFs for a subset of the histories wherein $G(\cdot)$ is either greater than or less than 0.5.

The estimated GIRFs for each estimated model are illustrated in Figure 5. There are several noteworthy results. Regarding the OSB price pairs associated with Canada and the North Central U.S., Canada and the Southeastern U.S., Canada and the Southwestern U.S., and the North Central U.S. and Southeastern U.S., there seems to be evidence that, on average, the GIRFS return to zero, but only after a rather lengthy period of time has passed. This result must, in turn, be attributed to the potential role of transactions costs as captured by the nonlinear features of the estimate models. Secondly, in at least three instances (i.e., those

involving Canada and the North Central U.S., Canada and the Southeastern U.S., and the North Central U.S. and Southeastern U.S.) there is a notable difference in the time paths of the GIRFs depending on which state the market is in at the time the shock is introduced (i.e., either $G(.) = 0$ or $G(.) = 1$). Overall, the results of the generalized impulse response function analysis indicate that the effects of market shocks do eventually dissipate, but likely at a much slower rate than would be implied by linear models that do not account for the role of transactions costs.

5 Summary and Conclusions

We have reviewed and evaluated alternative linear and nonlinear smooth transition autoregressive models that can advance our understanding of spatial market integration in oriented strand board markets. Our results show that all tested price pairs for oriented strand board in North America create stationary linear combinations, or cointegrated price pairs, and that these pairs also exhibit nonlinearities in the data generating process. Moreover, the nonlinearities in the price relationships support the application of quadratic smooth transition autoregressive models to evaluate whether the nonlinearities derive from transactions costs, and therefore the threshold parameters in the DGP. By using several measures of fit and performance we find that the QSTAR models do a superior job of explaining relative price movements than do their linear counterparts. Aside from providing confirmation that the Law of One Price—augmented to account for transactions costs bands—holds for oriented strand board markets in North America, the QSTAR results imply that market shocks are more persistent than linear models would suggest.

The primary implication of these findings is that market models that ignore the existence of transactions cost bands in these markets would tend to overestimate the rate of adjustment to market shocks brought about by catastrophic events or regional housing market shifts. For example, small shocks to local or regional housing markets would likely not transmit quickly

to distant markets, but large shocks could register quickly, through a spatial arbitrage process. Another implication of the finding is that transmission of market shocks of these kind would depend on initial price differentials, yielding rapid responses when prices are far apart but slow ones when they are not.

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Table 1: Results of Dickey-Fuller Tests Applied to Six Regional OSB Price Pairs.

Price Pair	No Intercept/Trend			Intercept Included, No Trend			Intercept and Trend Included		
	$\hat{\rho}$	τ	p -value	$\hat{\rho}$	τ_μ	p -value	$\hat{\rho}$	τ_τ	p -value
$\ln(p_1/p_2)$	0.976	-1.968	0.061	0.924	-3.497	0.009	0.895	-4.110	0.009
$\ln(p_1/p_3)$	0.973	-2.217	0.035	0.908	-4.268	0.001	0.885	-5.019	0.001
$\ln(p_1/p_4)$	0.965	-2.818	0.005	0.902	-5.433	0.001	0.866	-6.434	0.001
$\ln(p_2/p_3)$	0.886	-4.857	0.001	0.843	-5.833	0.001	0.843	-5.830	0.001
$\ln(p_2/p_4)$	0.836	-6.667	0.001	0.825	-6.921	0.001	0.809	-7.263	0.001
$\ln(p_3/p_4)$	0.899	-3.414	0.002	0.848	-4.336	0.003	0.793	-5.164	0.002

Note: $\hat{\rho}$ is the estimated root. The test statistics τ , τ_μ and τ_τ are t -ratios for $(\hat{\rho}-1)$, and correspond, respectively, to: (1) the case where the estimated model does not include an intercept or a trend; (2) the case where the estimated model does include an intercept but no trend; and (3) the case where the estimated model includes both an intercept and a linear trend. Columns headed p -value record approximate p -value's based on $B = 999$ bootstrap simulations.

Table 2: Results of Testing a Linear Unit Root Model Against STAR-Type Alternatives for Regional OSB Relationships.

Candidate Transition Variable, s_t	<i>Price Pair</i>					
	$\ln(p_1/p_2)$	$\ln(p_1/p_3)$	$\ln(p_1/p_4)$	$\ln(p_2/p_3)$	$\ln(p_2/p_4)$	$\ln(p_3/p_4)$
Δy_{t-1}	0.003	0.001	0.001	0.018	0.001	0.001
Δy_{t-2}	0.223	0.001	0.002	0.014	0.004	0.002
Δy_{t-3}	0.548	0.001	0.001	0.003	0.003	0.003
Δy_{t-4}	0.631	0.001	0.001	0.001	0.005	0.015
Δy_{t-5}	0.420	0.125	0.027	0.073	0.330	0.013
Δy_{t-6}	0.346	0.020	0.002	0.026	0.034	0.022
Δy_{t-7}	0.454	0.003	0.002	0.019	0.370	0.144
Δy_{t-8}	0.007	0.019	0.062	0.109	0.104	0.021
Δy_{t-9}	0.031	0.049	0.085	0.008	0.013	0.002
Δy_{t-10}	0.327	0.220	0.345	0.061	0.374	0.001
Δy_{t-11}	0.482	0.017	0.097	0.096	0.112	0.001
Δy_{t-12}	0.523	0.080	0.393	0.003	0.111	0.008
$(y_{t-1} - y_{t-2})$	0.003	0.001	0.001	0.018	0.001	0.001
$(y_{t-1} - y_{t-3})$	0.007	0.008	0.003	0.007	0.015	0.001
$(y_{t-1} - y_{t-4})$	0.010	0.001	0.001	0.001	0.002	0.082
$(y_{t-1} - y_{t-5})$	0.006	0.001	0.001	0.001	0.001	0.025
$(y_{t-1} - y_{t-6})$	0.028	0.001	0.001	0.001	0.001	0.001
$(y_{t-1} - y_{t-7})$	0.021	0.001	0.001	0.001	0.001	0.008
$(y_{t-1} - y_{t-8})$	0.058	0.001	0.001	0.001	0.004	0.047
$(y_{t-1} - y_{t-9})$	0.084	0.001	0.001	0.001	0.003	0.002
$(y_{t-1} - y_{t-10})$	0.027	0.001	0.001	0.001	0.001	0.004
$(y_{t-1} - y_{t-11})$	0.017	0.001	0.001	0.001	0.001	0.058
$(y_{t-1} - y_{t-12})$	0.018	0.001	0.001	0.001	0.001	0.211
$(y_{t-1} - y_{t-13})$	0.020	0.001	0.001	0.003	0.001	0.269
Opt. Lag Length	9	4	7	6	12	8

Note: All fractions are p -values associated with testing a linear unit root model against a STAR-type alternative using the indicated transition variable, s_t , and by using a fourth-order approximation to the transition function. All p -values are obtained by performing $B=999$ recursive bootstraps of the model under the null hypothesis of linearity and a unit root process. Bolded values indicate the minimum p -value(s) obtained for each price pair. The final row indicates the optimal lag length for each price pair, determined by minimizing the AIC.

Table 3: Measures of Model Fit and Diagnostic Test Results for Estimated QSTAR Models.

Price Pair	Trans. Var.	$\hat{\sigma}_\varepsilon$	AIC	R^2	$\hat{\sigma}_{NL}/\hat{\sigma}_L$	SK	EK	LJB	LM_{AR}	LM_{ARCH}	Diagnostic Tests				
											Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	Δy_{t-4}	t
$\ln(p_1/p_2)$	Δy_{t-1}	0.335	-2.102	0.132	0.976	0.159	3.877	370.80(3.02E-81)	0.880	0.135	0.058	0.583	0.290	0.475	0.121
$\ln(p_1/p_3)$	Δy_{t-3}	0.333	-2.148	0.115	0.980	0.004	2.822	188.82(9.95E-42)	0.316	0.001	0.203	0.300	0.219	0.113	0.009
$\ln(p_1/p_4)$	$y_{t-1} - y_{t-4}$	0.370	-1.916	0.122	0.956	0.002	3.148	234.73(1.07E-51)	0.077	1.47E-05	0.848	0.441	0.600	0.551	0.100
$\ln(p_2/p_3)$	$y_{t-1} - y_{t-9}$	1.031	0.124	0.148	0.977	0.003	0.770	14.32(7.79E-4)	0.796	0.045	0.298	0.852	0.573	0.111	0.308
$\ln(p_2/p_4)$	$y_{t-1} - y_{t-6}$	2.022	1.514	0.165	0.969	0.109	1.992	104.18(2.38E-23)	0.804	2.91E-04	0.262	0.417	0.621	0.883	0.229
$\ln(p_3/p_4)$	Δy_{t-2}	1.251	0.526	0.188	0.976	0.010	1.363	44.91(1.77E-10)	0.327	0.024	0.662	0.428	0.494	0.055	0.297

Note: The column headed Trans. Var. denotes the transition variable used in the final model specification; $\hat{\sigma}_\varepsilon$ is the residual standard error; AIC is Akaike information criterion; R^2 is the unadjusted R^2 , and $\hat{\sigma}_{NL}/\hat{\sigma}_L$ is the ratio of the residual standard error from the respective STAR-type model relative to the linear AR model. As well, SK is skewness, EK is excess kurtosis, and LJB is the Lomnicki-Jarque-Bera test of normality of residuals, with asymptotic p -values in parentheses. LM_{AR} denotes the F variant of Eitrheim and Teräsvirta's (1996) LM test of no remaining autocorrelation in the residuals based on four lags. Likewise, LM_{ARCH} denotes an LM test for ARCH-type heteroskedasticity based on four lags. Remaining columns report p -values for F variants of Eitrheim and Teräsvirta's (1996) LM tests for remaining nonlinearity, $(\Delta y_{t-1}, \dots, \Delta y_{t-4})$, and parameter constancy, t .

Table 4: Estimated Half Lives for the Estimated AR and QSTAR Models for Regional OSB Price Relationships.

Price Pair	Parameter	AR Model	QSTAR Model
$\ln(p_1/p_2)$	$1 - \hat{\rho}$	-0.072	-0.169
	Half Life:	9.288	3.752
$\ln(p_1/p_3)$	$1 - \hat{\rho}$	-0.092	-0.262
	Half Life:	7.180	2.281
$\ln(p_1/p_4)$	$1 - \hat{\rho}$	-0.074	-0.076
	Half Life:	8.986	8.713
$\ln(p_2/p_3)$	$1 - \hat{\rho}$	-0.150	-0.850
	Half Life:	4.276	0.365
$\ln(p_2/p_4)$	$1 - \hat{\rho}$	-0.166	-0.196
	Half Life:	3.810	3.175
$\ln(p_3/p_4)$	$1 - \hat{\rho}$	-0.152	-0.539
	Half Life:	4.208	0.896

Note: $1 - \hat{\rho}$ is the estimated coefficient on the lagged level term in the respective model. The half life for QSTAR models is constructed by using the estimated $1 - \hat{\rho}$ coefficient corresponding to the regime implied when $G(s_t; \gamma, \mathbf{c}) = 0$. All half lives are expressed in weeks.

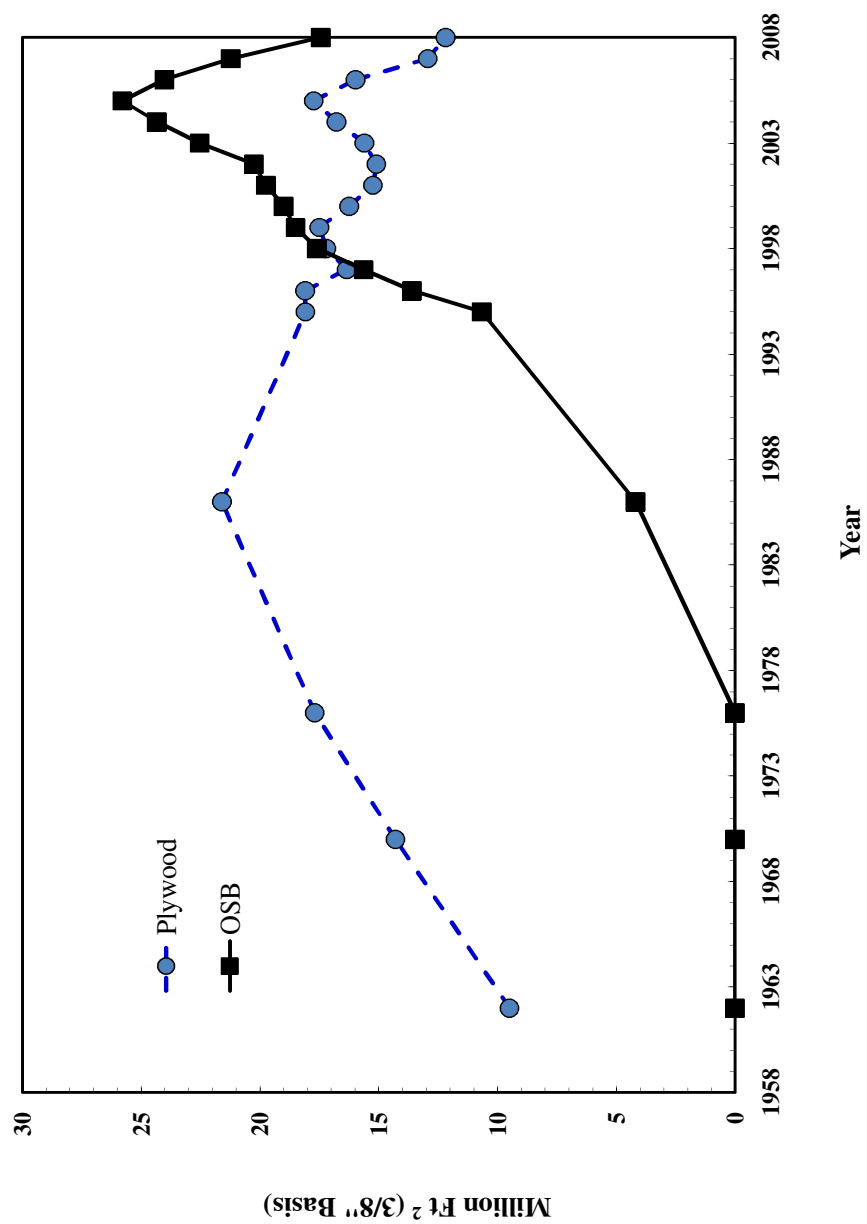


Figure 1: U.S. Plywood and OSB Consumption, 1958–2007 and Projected for 2008.

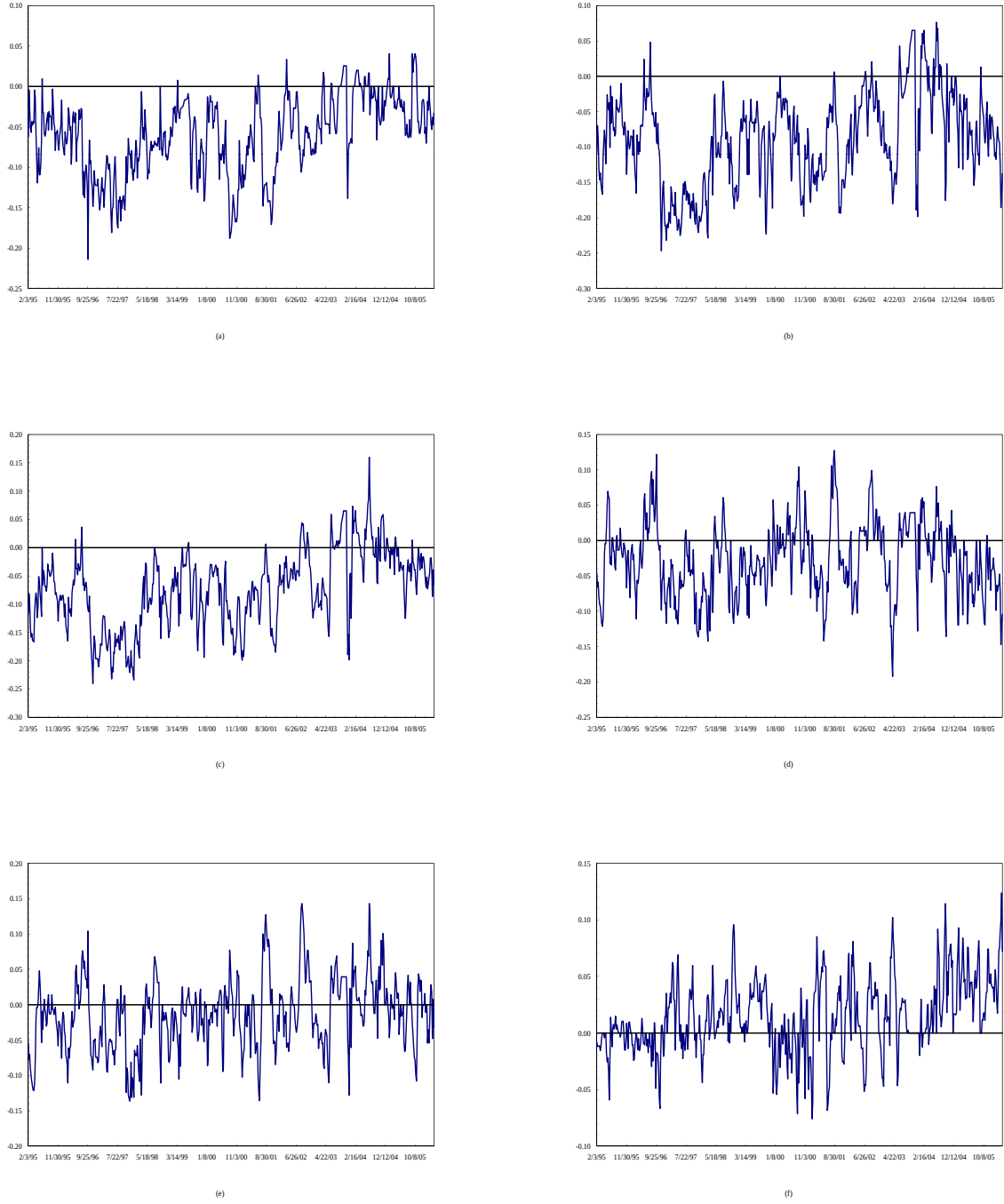


Figure 2: Time-Series Plots of Natural Logarithms of Regional Relative Prices for Oriented Strand Board. (a) OSB price in Eastern Canada relative to North Central U.S.; (b) OSB price in Eastern Canada relative to Southeastern U.S.; OSB price in Eastern Canada relative to Southwestern U.S.; (d) OSB price in North Central U.S. relative to Southeastern U.S.; (e) OSB price in North Central U.S. relative to Southwest U.S.; and (f) OSB price in Southeastern U.S. relative to Southwest U.S.

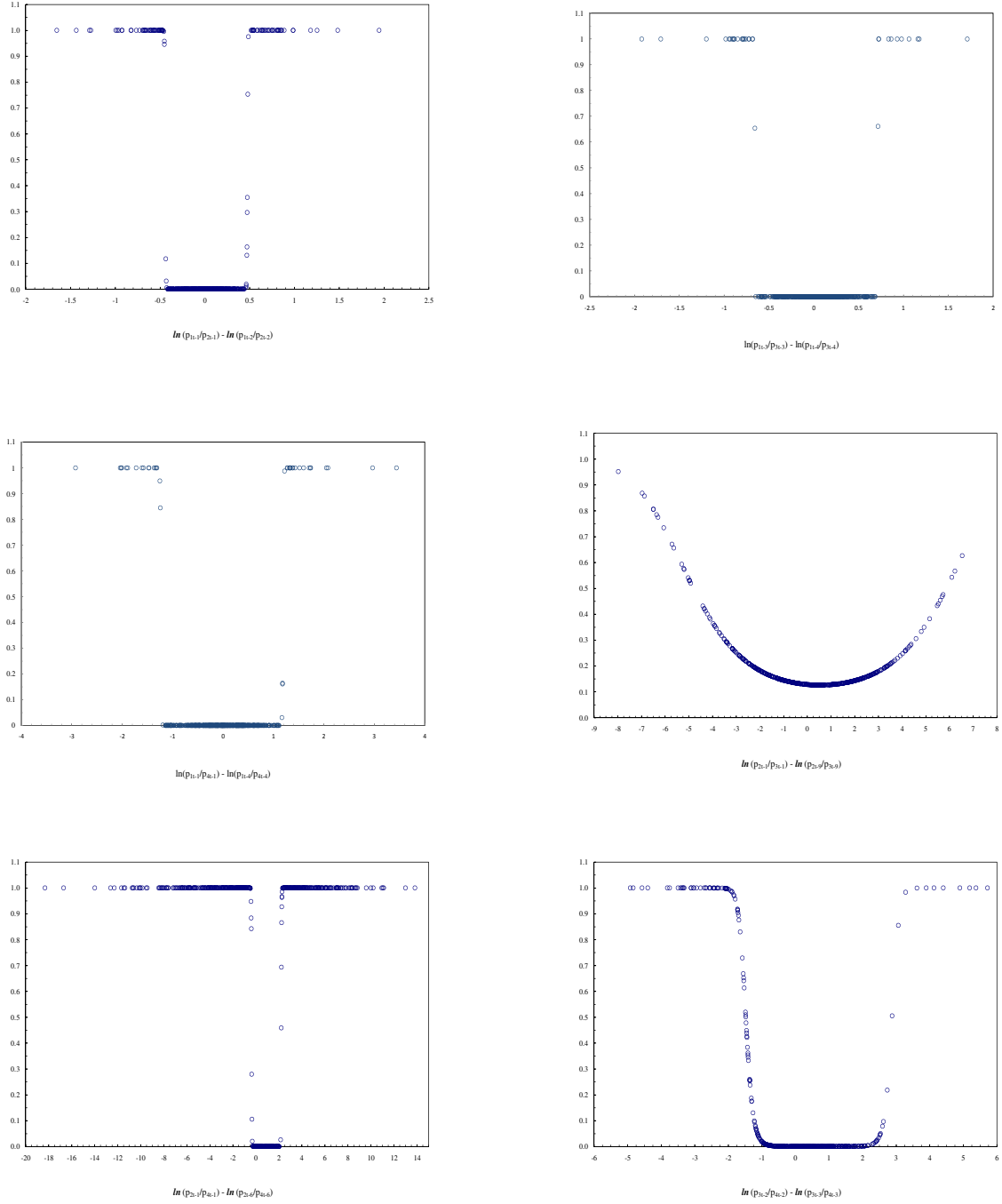


Figure 3: Estimated Transition Functions for Six Regional Oriented Strand Board Price Relationships. (a) OSB price in Eastern Canada relative to North Central U.S.; (b) OSB price in Eastern Canada relative to Southeastern U.S.; OSB price in Eastern Canada relative to Southwestern U.S.; (d) OSB price in North Central U.S. relative to Southeastern U.S.; (e) OSB price in North Central U.S. relative to Southwest U.S.; and (f) OSB price in Southeastern U.S. relative to Southwest U.S.

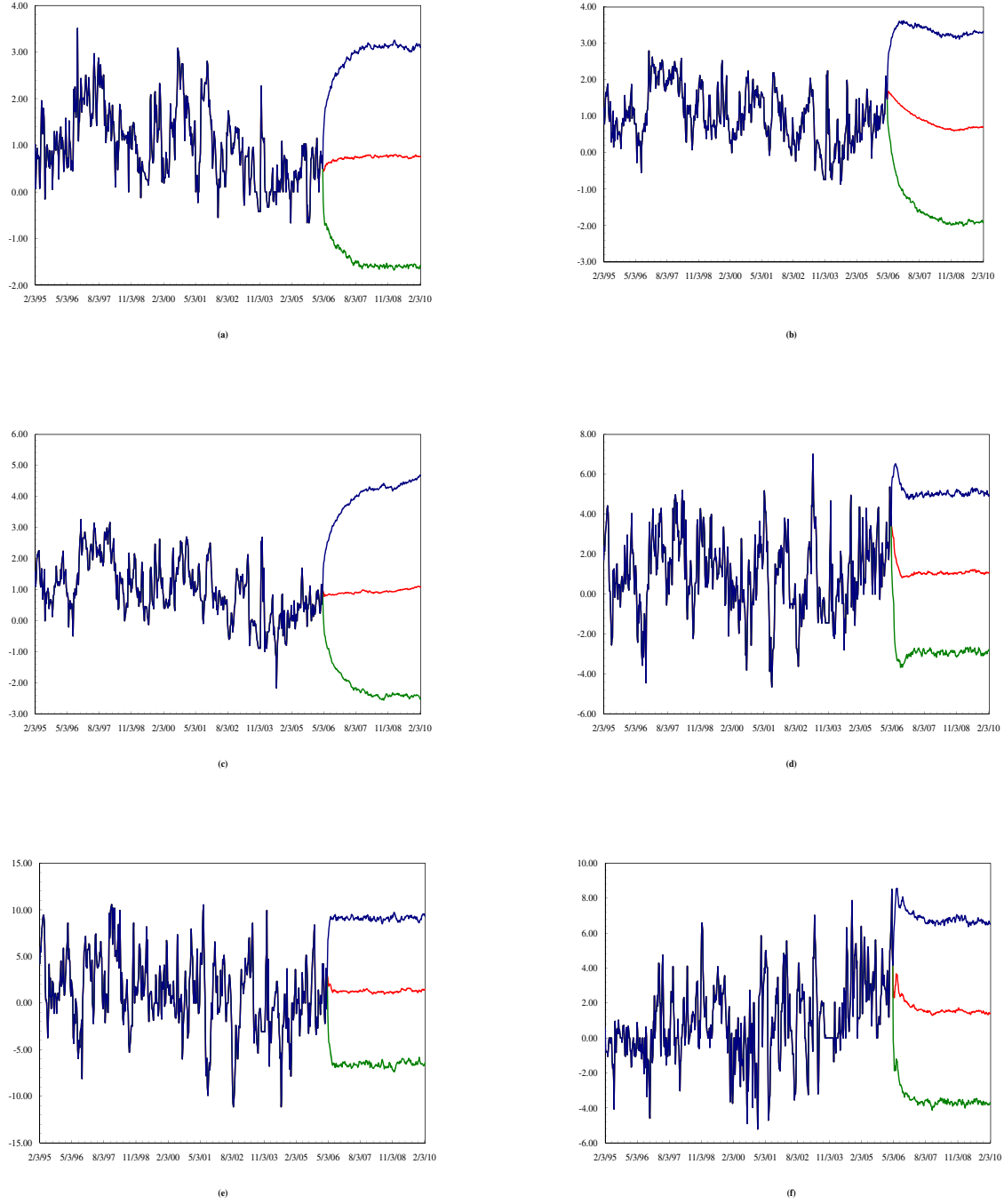


Figure 4: Observed Data and Stochastic forward Simulations of the Estimated QSTAR Models from the End of the Sample with 95-percent Confidence Intervals. (a) OSB price in Eastern Canada relative to North Central U.S.; (b) OSB price in Eastern Canada relative to South-eastern U.S.; OSB price in Eastern Canada relative to Southwestern U.S.; (d) OSB price in North Central U.S. relative to Southeastern U.S.; (e) OSB price in North Central U.S. relative to Southwest U.S.; and (f) OSB price in Southeastern U.S. relative to Southwest U.S.

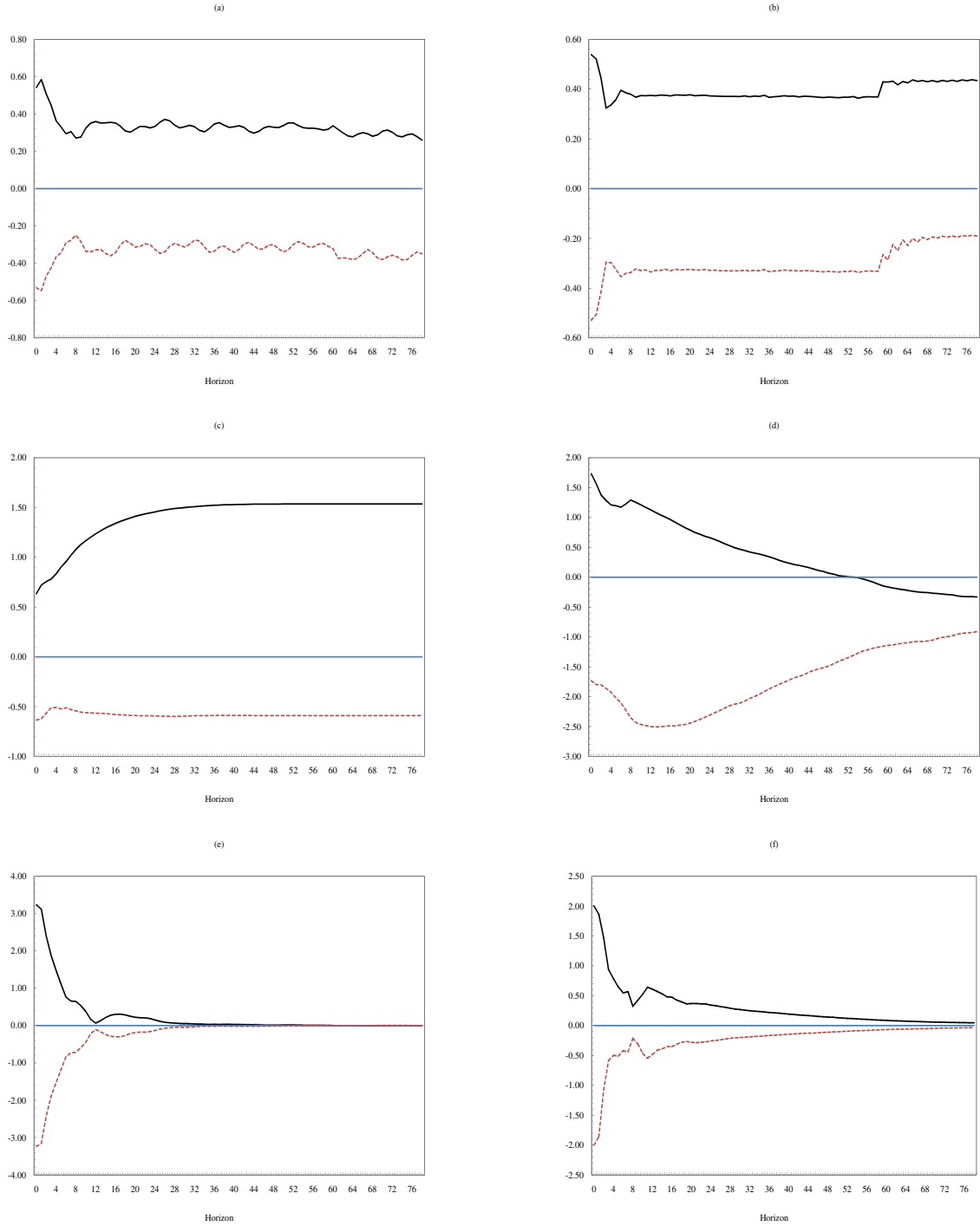


Figure 5: Estimated Generalized Impulse Response Functions for Six Regional Oriented Strand Board Price Relationships. (a) OSB price in Eastern Canada relative to North Central U.S.; (b) OSB price in Eastern Canada relative to Southeastern U.S.; (c) OSB price in Eastern Canada relative to Southwestern U.S.; (d) OSB price in North Central U.S. relative to Southeastern U.S.; (e) OSB price in North Central U.S. relative to Southwest U.S.; and (f) OSB price in Southeastern U.S. relative to Southwest U.S. A solid line denotes response to a positive shock and a dashed line response to a negative shock.