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State-Space Cointegration Modeling for the Analysis of Exogenous Shocks to Prices in Israeli-Palestinian Food Trade

Rico Ihle[a], **Linde Götz**[b] **and Ofir D. Rubin**[c]

Abstract: The Israeli-Palestinian conflict constitutes a prominent example of a long-lasting political conflict which has major consequences for the livelihoods of the people on both sides. The agricultural sectors of the Palestinian and Israeli economies are tightly connected. However, various security measures occasionally implemented in consequence of the political conflict between Israelis and Palestinians strongly inhibit the movement of people and commodities. In order to obtain evidence on the impacts of such abrupt trade impediments on price dynamics, we estimate a vector error correction model in state space form employing the Kalman filter. The time-varying cointegration parameters suggest that the security measures indeed impacted price interdependencies on a short-term scale.

Keywords: cointegration, error correction, Israel, Palestinian territories, price transmission, state space model, time-varying parameters.

JEL: C32, D74, Q11, Q13, F15

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1 Introduction

The Israelis and the Palestinian people residing in the occupied territories have close trading relationships since the Six-Day war in 1967. The small geographical scope of the region and a number of apparent differences in production factors give rise to successful traffic of agricultural produce between Israel and the Palestinian territories in the course of time. The Israeli agriculture sector is capital intensive, publicly supported and with advanced marketing systems in international markets. The Palestinians agriculture sector on the other hand consists on relatively high supply of unskilled labor and depends for some extend on Israel to conduct trade relationships with markets overseas. Agriculture has an important significance in the region. Due to the tense political situation in the Middle East, the ability to be food-independent is of highest political priority for Israel. For the Palestinians, the agriculture sector comprises a significant part of the national income. Cultivated land in the West Bank and Gaza jointly is approximately 96,000 hectares which accounts for 15.7% and 20.6% of total land in the respective region (PCBS 2011a). Accordingly, the share of households earning a living from agricultural activities in the Palestinian territories has always been a significant segment in the population ranging from 12% to 16% between 1995 and 2008 (PCBS 2011b). The choice of crop production is a direct implication of the comparative advantages of each side. Fruits which require relatively more capital are grown in Israel and exported to the Palestinians territories while vegetables which are labor intensive flow the other way.

The aim of this paper is to provide empirical evidence on the impacts of exogenous factors on agricultural trade in the context of a political conflict. In the context of the Israeli-Palestinian conflict, economic relationships in the region are often subject to challenging conditions. At times of intensified violence, trade is cut off completely while in periods which are relatively more quiet, security measures put forward by the Israeli Defense Forces (IDF) generate costly barriers for trade. In this paper, we examine a recent period of the conflict during which movement restrictions and other trade barriers were rarer. Movement restrictions that we look at correspond by large only to times of Israeli national holidays in which atmosphere are sensitive by nature in this region. Typically, during these days a complete closure is put on the Palestinian territories and bilateral trade is terminated. These security measures are usually implemented on very short notice due to security alerts and hence tend to interrupt economic relationships and trade in an abrupt manner.

In this paper we focus on the economic consequences of the complete closures of the West Bank security barrier implemented by the IDF between Israel and the West Bank on wholesale price relationships in both agricultural markets.[1] We restrict our attention to the fruit and vegetable wholesale markets in Tel Aviv in Israel and Hebron in the West Bank and focus on the price interdependencies of apples and cucumbers, two products which are traded continuously through the year between these regions. The complete closures of the barrier interrupt movements of goods and people between both regions on very short notice which suggests an alternative modeling approach to the traditional cointegration framework which

[1] Since decades, this conflict is followed by the international media and dealt with extensively by the international community. Hence, data availability and reliability are in general comparatively satisfying in order to be able to carry out quantitative analyses on this subject. We do not regard Gaza because data is difficult to obtain and conditions differ fundamentally from the ones existing in the West Bank.

assumes constant parameters generating the data. Hence, this temporal characteristic of the implementation of the security measures will guide the model choice for the analysis.

We suggest an alternative modeling framework which allows the model parameters to vary over time in order to model the short-term impacts of security measures on price relationships. The state space approach for time-varying parameters seems adequate in this context. Besides, this model approach is expected to provide further insights into price dynamics since the condition of stable parameters is relaxed.

State-space models have been developed by control engineers and physicists to model time-varying dynamics of time-series data which are governed by an unobservable variable. They were made available to applications in financial economics by the works of Harvey (1981) and Hamilton (1988) among others. This model type is for example used to analyze financial integration (e.g. Haldane and Hall, 1991) or capital market integration (e.g. Bekaert and Harvey,1995).

However, state-space models are applied to study market integration in other sectors well. Often, an error correction model approach is chosen to study integration or convergence (Harvey, 1981) of markets. For example, Neumann et al. (2006) investigate convergence of the European natural gas market by modelling the long-run equilibrium relationship with a time-varying transmission parameter between 3 spot markets for natural gas. Also, an error correction model with unobserved components (time-varying parameters) is estimated within a state space model approach by Carvalho and Harvey (2005) to investigate growth, cycles and convergence of real income per capita in 8 regions of the USA. Also, Li et al. (2006) develop an error correction model with time-varying parameters to forecast the level and growth rate of tourist demand, assuming the existence of a constant long-run equilibrium relationship, whereas the speed of the correction of deviations from the long-run equilibrium may change over time. Results suggest that the time-varying error correction model exhibits superior forecast performance over all alternative fixed-parameter single equation models, e.g. of the Engle-Granger error correction, ARIMA or the vector autoregressive model types.

However, a state space version of the error correction model to estimate time-varying parameters has not yet been used to investigate spatial market integration in the agricultural sector[2]. This paper aims to close this gap by investigating the long-run price relationship between markets in Israel and the Palestinian territories within a state space model approach. We focus on the influence of the complete boarder closings on the integration of these two markets. Section 2 presents the market context and the data used in the analysis. Section 3 elaborates on the relationship between the traditional error correction model with constant parameter and the suggested time-varying version. Section 4 explains the model framework and section 5 presents the empirical results. In section 6 we provide a discussion.

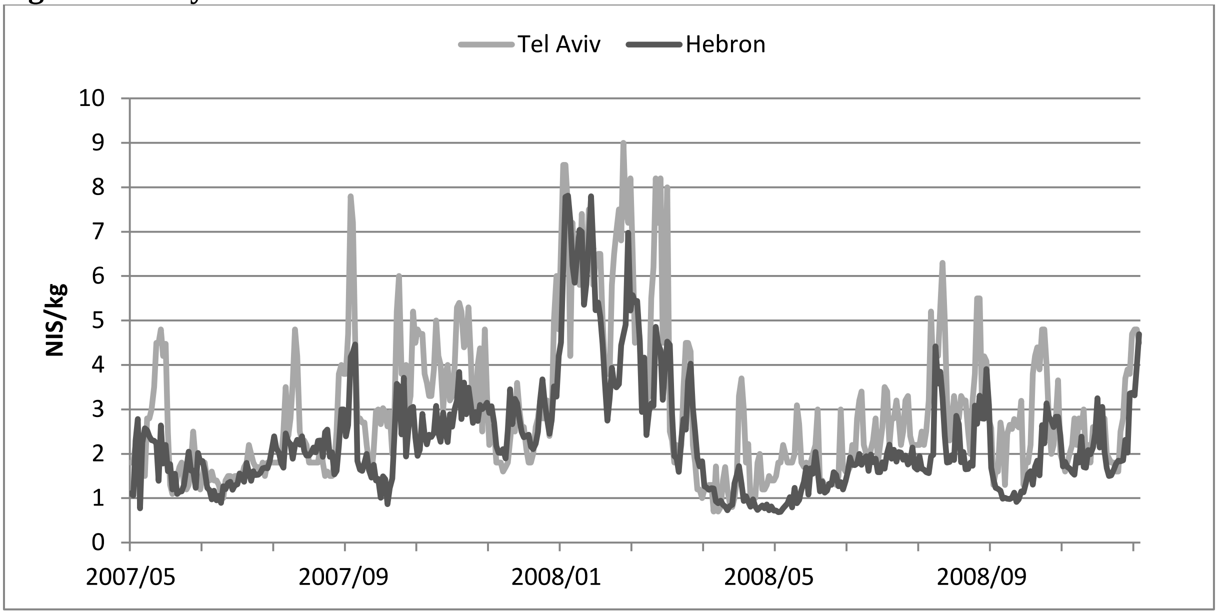
2 The market context and the data

This study analyses price time series observed in the largest fruits and vegetables wholesale markets in Israel (Tel Aviv) and the West Bank (Hebron). We focus on apples and cucumbers, two commodities which are extensively traded during the entire year on wholesale markets in

[2] The study by Walburger and Foster (1998) on regional price dynamics of the fed cattle markets in the USA is based on the state space technique suggested by Aoki (1989). They apply a multivariate ARMA model to investigate dynamic factors.

Israel and the West Bank. While apples almost exclusively flow from Israel to the West Bank, cucumbers are mainly produced in the West Bank and represent one of the most important vegetables shipped from the Palestinian territories to Israel (while only minor quantities traded in the opposite direction). Each series of daily prices has 437 observations and ranges from May 2007 until the end of December 2008 (Figures 1 and 2). Since the markets almost operate from Sundays to Thursdays, we use datasets without observations for Fridays and Saturdays. Missing values were imputed using an adaptation of the R package Amelia II (Honaker et al., 2009). Table 1 compares the original and the complete datasets containing the imputed values suggesting that the imputations resemble the structure of the original time series very well.

Figure 1: Daily Cucumber Prices in Tel Aviv and Hebron



Source: IMA (2011) and Hebron wholesale market (2011).

Table 1 shows that the imputations resemble the structure of the original time series very well. Several characteristics of the trade relationships are illustrated by the graphs and the descriptive statistics. The most remarkable difference between the two commodities are the large coefficients of variation for cucumbers and the large difference in mean prices between Hebron an Tel Aviv for apples. The mean prices of cucumbers and even more of apples are lower on the Hebron market than on the Tel Aviv market. This might be due to, e.g., quality differences of the varieties sold (higher quality produce is sold on the market in Israel for a higher price), price discrimination of Israeli marketing companies or higher marketing costs resulting from e.g. storage costs or higher labour costs.

Table 1: Descriptive Statistics of the Prices

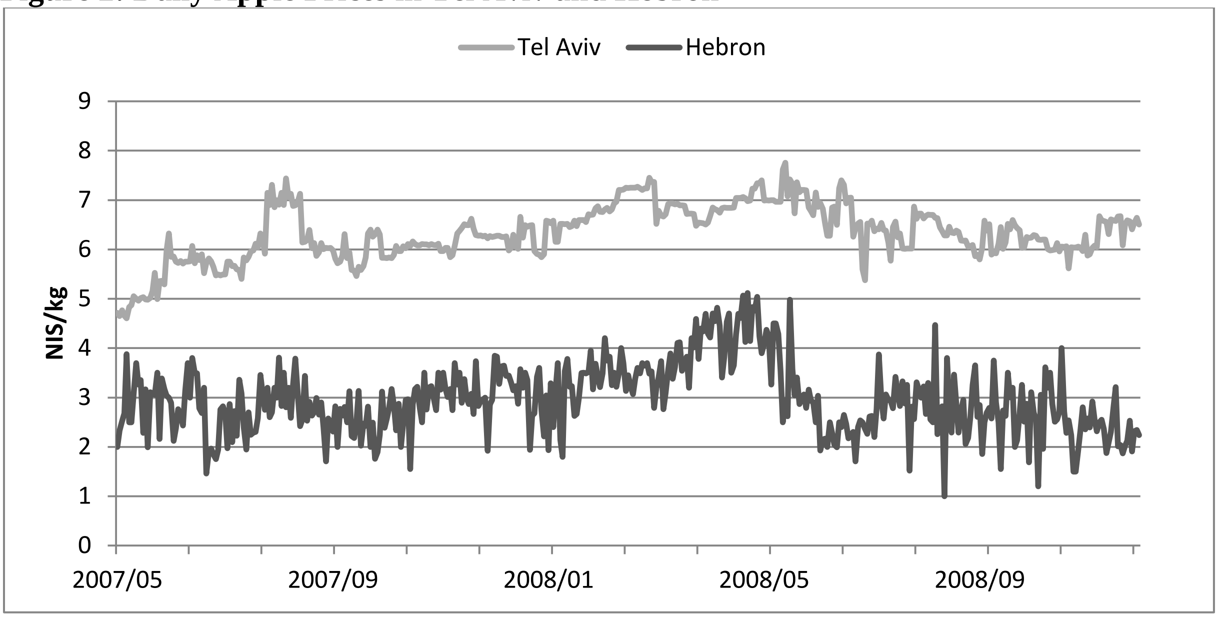
	Cucumbers		Apples	
Wholesale market	Hebron	Tel Aviv	Hebron	Tel Aviv

data	observed	Incl. imputed values	observed	Incl. imputed values	observed	Incl. imputed values	observed	Incl. imputed values
Nb. obsv.	427	437	406	437	374	437	402	437
Mean	2.28	2.31	2.97	2.99	2.95	2.97	6.33	6.34
Median	1.95	1.96	2.50	2.50	2.87	2.91	6.29	6.32
Stand. dev.	1.28	1.29	1.68	1.65	0.74	0.72	0.57	0.56
Min	0.68	0.68	0.70	0.70	1.00	1.00	4.61	4.61
Max	7.81	7.81	9.00	9.00	5.12	5.12	7.76	7.76
Var. coef.	0.56	0.56	0.56	0.55	0.25	0.24	0.09	0.09

Source: Authors' calculations.

The coefficient of variation indicates that the price volatility of cucumbers is considerably higher than for apples in both markets. While the coefficient of variation for cucumbers does not differ between both markets, the coefficient of variation for apples in the West Bank is substantially higher than in Tel Aviv.

Figure 2: Daily Apple Prices in Tel Aviv and Hebron



Source: IMA (2011) and Hebron wholesale market (2011).

The trade relationships between the agricultural markets in Israel and the West Bank can abruptly be interrupted by the consequences of security alerts. The measures to curb security threats can take a number of forms and can be implemented on short notice on a regional basis (see, e.g., OCHAoPt, 2011). A major reaction to the unstable state of security in both regions was the establishment of a comprehensive system of institutions to monitor and control the movement of people and goods within the Palestinian territories (currently only inside the West Bank) and between these regions and Israel, for example permanent and temporary road blocks and check points (World Bank, 2007; World Bank, 2008). Since reliable data on movement restrictions within the West Bank is difficult to obtain, we focus on the effects of the comprehensive closures of the West Bank barrier (OCHAoPt, 2011) implemented by the IDF. Information on the closure incidences is published by the Israeli humans rights NGO B'Tselem based on data from the IDF (Table 2). The comprehensive closures yield any permits obtained for residents of the West Bank to be invalid so that they cannot leave the West Bank (B'Tselem, 2011). That is, the security barrier encircling the West Bank is completely closed so that movements of people and goods across the barrier are effectively blocked. The closures are announced in advance some days before the implementation; they are usually implemented for only short periods of a few days as Table 2 illustrates.

Table 2: Days of Comprehensive Closures

Month	2007	2008
January		2
February		
March		4
April		7
May		9
June		2
July		
August		
September	5	2
October	4	10
November		
December		
Sum	9	36

Source: B'Tselem (2011).

Given this market setting of price relationships and trade directions, expectations on the effects of the complete closures can be deduced. The strength of the price effects to these exogenously induced supply shocks is likely to depend on the responsiveness of the wholesale prices to local supply changes as induced by the short-termed closures of the trade relationships. The absolute price response depends on the duration of the exogenously enforced interruption of the trade relationships and the degree of perishableness of the food product (Ward, 1982; Cutts and Kirsten, 2006). Since cooling storage capacities are cheaper and easier available for Israeli wholesalers than for Palestinian wholesalers, the former might be better able to cope with such short term supply shocks either by withholding stocks intended for delivery or by storing the increasingly sparse commodity for a longer period. Moreover, Israeli wholesalers might potentially source the commodity from other regions while a complete closure in the West Bank implies also the impossibility of trade across its Eastern border towards Jordan. Consequently, price responses in the West Bank might tend to be more pronounced than in Israel which is also suggested by the coefficients of variation in Table 1 above.

An interruption of the apple trade from Israel to the Palestinian territories increases apple supply on the Israeli market and reduces supply on the West Bank market since the Israeli excess supply cannot enter the Palestinian territories. Apple prices there are therefore likely to increase in the Palestinian territories leading to a lower (negative) price differential between Hebron and Tel Aviv. The stronger the reaction of Tel Aviv wholesale prices is, the larger the resulting effect on the price differential will be. On the other hand, a disturbance of cucumber trade might push Tel Aviv prices markedly above Hebron prices so that the margin becomes increasingly negative.

3 Time-varying parameters in the vector error correction model

Price interdependencies are often represented by cointegration models. One prominent model type is the vector error correction model (VECM) which typically takes the form:

$$\Delta p_t = \alpha\beta' p_{t-1} + \sum_{i=1}^{k} \Gamma_i \Delta p_{t-i} + \varepsilon_t = \alpha\ eqe_{t-1} + \sum_{i=1}^{k} \Gamma_i \Delta p_{t-i} + \varepsilon_t = \Pi p_{t-1} + \sum_{i=1}^{k} \Gamma_i \Delta p_{t-i} + \varepsilon_t. \qquad (1)$$

In spatial price transmission analysis, the model is apt to consider a vector $p_t = \{p_t^1, \ldots, p_t^v\}'$ of price series each of which represents the price of a homogenous good in a certain location. The operator Δ denotes the first difference operator, i.e., $\Delta p_t = p_t - p_{t-1}$. The model partitions the current price change Δp_t into two partial sources. The first term represents the deviations from equilibrium (the equilibrium errors $\beta' p_{t-1} = eqe_{t-1}$). The matrix β estimates the long-run coefficients (the weights) of the prices in equilibrium and α quantifies the feedback coefficients, that is, the adjustment speeds with which a certain equilibrium error is corrected in each period. The second term quantifies the partial impact of past price changes (the short-run parameters Γ_i). ε_t and *k* denote the Gaussian white noise and the lag length, respectively.

The core assumption of this model is parameter constancy, i.e., the coefficients of the model are assumed not to change (to be constants). In recent years, a number of more flexible versions of the VECM have been developed in order to account for various characteristics of price dynamics of agricultural markets. These models have in common that they no longer assume globally constant parameters but allow for changing model coefficients (for an overview see, e.g. Ihle, 2010). Most of these approaches look at parameters as being regime-dependent, that is, some or all parameters take locally constant values, that is, a constant value in each regime while the parameters are allowed to vary across regimes. The most prominent example is certainly the threshold VECM (Goodwin and Piggot, 2001) which estimates the regime-classifying parameter from the data in order to account for transaction costs. The Markov-switching VECM of Brümmer et al. (2009) estimates the regime classification based on unobserved states in the price relationships (see also Götz et al. 2010). While also being regime-dependent, does the smooth transition VECM of Mainardi (2001) and Serra et al. (2011) rather view the parameters as smoothly varying. Ihle and Amikuzuno (2010) develop a multivariate exogenous regime switching model which is a regime-dependent model suitable for a multivariate context.

Very recently, the methodological development was extended to time series models with parameters which are allowed to take any smooth form. The parameters are no longer thought of as locally constant but as smoothly varying instead. The advantage of these approaches is their flexibility and their explorative power since they estimate the functional forms of the model parameters based on semi- or nonparametric regression techniques. Although extremely flexible and useful in applied analysis, these varying coefficients models rely on cross

The VECM has three types of parameters to be estimated: the cointegration matrix β, the loading matrix α containing the feedback coefficients between the past deviation from equilibrium and the current price movement and the autoregressive short-run parameters Γ_i.

Although often the first two parameters are of interest for interpretation, all three can be allowed to be time-varying in the context of state space models. Time-varying adjustment speeds α_t and time-varying short-run parameters Γ_{it} may be used to obtain evidence of high interest regarding the short-run price dynamics in a similar fashion as the above-mentioned regime-dependent models without changing the usual interpretation. The main additional information offered by the state space view is the temporal development of the parameters. This evolvement over time is modelled in the most flexible way without imposing assumptions of global (as in the case of the traditional VECM) or local, that is, regime-dependent parameter constancy (as for example in the cases of the threshold or the Markov-switching VECMs). Parameter estimates for each observation may be obtained instead which leads to similar results as recursive parameter estimation approaches (see, e.g., Juselius, 2008, chapter 4).

Moreover, the time-varying parameter approach opens a novel perspective if it is applied to the cointegration parameters in β which we outline in some detail in the following since it is not straightforward to notice from the slightly changed notation in the time-varying approach. The parameters in β take the form of a two-dimensional vector if only prices in markets *A* and *B* are regarded, that is, $p_t = \{p_t^A, p_t^B\}'$. The equilibrium errors of the traditional VECM (indicated as **model I**) of two variables may hence be written as:

$$eqe_t^I = \beta' p_t = p_t^{A,ob} - p_t^{A,eqI} = p_t^{A,ob} - (\beta^0 + \beta^1 p_t^{B,ob}) \qquad (2)$$

where the superscripts 0 and 1 only represent parameters indices, so that $\beta = (1 \quad -\beta^0 \quad -\beta^1)'$. The variables $p_t^{A,ob}$ and $p_t^{B,ob}$ denote the observed prices in both markets. Moreover, $p_t^{A,eqI}$ signifies the (hypothetical) equilibrium price in market *A* which is calculated based on the parameters in β quantifying the *long-run price equilibrium relationship* between both markets. This notion of "long-run" means that the parameters in β represent the stable average relationship between the prices.[3]

If these parameters are allowed to be time-varying (**model II**)[4], then a time index has to be added to the cointegration parameters in (2):

$$eqe_t^{II} = p_t^{A,ob} - p_t^{A,eqII} = p_t^{A,ob} - (\beta_t^0 + \beta_t^1 p_t^{B,ob}). \qquad (3)$$

Although this change of notation seems to be very minor, it implies a fundamental change of model logic. First, note that the equilibrium price in *A* is denoted as $p_t^{A,eqII}$ since it might differ from $p_t^{A,eqI}$ of the model (2) with globally constant parameters. Consequently, also eqe_t^{II} and eqe_t^I might differ from each other, as we will show below, in a systematic way. The larger the differences of β_t^0 and β_t^1 from β^0 and β^1, respectively, the larger the difference between the equilibrium errors of models I and II for the time period *t* will be. The *resulting absolute partial impacts* of the equilibrium deviations (lagged by one period) on the current price

[3] This corresponds to the way of obtaining estimates of these parameters which also leads to an average relationship of the prices, that is, the conditional average of $p_t^{A,eq}$ given $p_t^{B,ob}$.

[4] For the sake of clarity, we assume that both models differ only in the time-varying β parameters. Hence, both of them have identical relative partial disequilibrium impacts and autoregressive impacts quantified by the globally constant α and Γ_i coefficients, respectively.

changes Δp_t will also differ accordingly although both models have *identical relative partial disequilibrium impacts* α. In other words, the implied error correction behaviors will be markedly different in both models.

The reason for this major difference is that model II partitions eqe_t^I into two components in the following way. First, assume that the time-varying parameters in (3) are related to the globally constant ones in (2) as:

$$\beta_t^0 = \beta^0 + \delta_t^0 \text{ and } \beta_t^1 = \beta^1 + \delta_t^1 \tag{4}$$

where δ_t^0 and δ_t^1 denote the deviations of the two time-varying β parameters from their globally constant counterparts in (2). Substituting (4) into (3) leads to:

$$eqe_t^{II} = p_t^{A,ob} - p_t^{A,eqII} = p_t^{A,ob} - (\beta^0 + \delta_t^0 + (\beta^1 + \delta_t^1) p_t^{B,ob}) \tag{5}$$

so that

$$\begin{aligned} eqe_t^{II} &= p_t^{A,ob} - p_t^{A,eqII} \\ &= p_t^{A,ob} - (\beta^0 + (\beta^1 p_t^{B,ob}) - (\delta_t^0 + \delta_t^1 p_t^{B,ob}). \end{aligned} \tag{6}$$

Finally, the systematic discrepancy of eqe_t^{II} from eqe_t^I becomes obvious by using (2):

$$\begin{aligned} eqe_t^{II} &= p_t^{A,ob} - p_t^{A,eqI} - (\delta_t^0 + \delta_t^1 p_t^{B,ob}) \\ &= eqe_t^I - (\delta_t^0 + \delta_t^1 p_t^{B,ob}). \end{aligned} \tag{7}$$

Hence, model II of the time-varying cointegration parameters β partitions the deviation from the (average) long-run price relationship eqe_t^I into two components $eqe_t^I = (\delta_t^0 + \delta_t^1 p_t^{B,ob}) + eqe_t^{II}$. This in turn means that it decomposes the error correction behavior of model I into two sub dynamics which are the (short-run) dynamics of the disturbance of the stable long-run equilibrium relationship $\delta_t^0 + \delta_t^1 p_t^{B,ob}$ in (2) and the (short-run) correction of the equilibrium errors $eqe_t^{II} = p_t^{A,ob} - p_t^{A,eqII}$, that is, the deviations of the observed prices in *A* from the disturbed long-run equilibrium price in *A* which represents itself a temporary equilibrium state.

Consequently, the equilibrium errors from the globally constant and the time-varying cointegration parameter models eqe_t^I and eqe_t^{II} will differ systematically from each other. The difference for a certain period *t* will be the larger, the larger the deviations of the time-varying β parameters from the average long-run price equilibrium parameters β^0 and β^1 are ($|\delta_t^0|$ and $|\delta_t^1|$). Hence, the more pronounced the time-varying (non-constant) behavior of the cointegration parameters is, the more the parameter estimates can be expected to differ. While the equilibrium errors eqe_t^I of model I with globally constant parameters might occasionally strongly differ from zero, the errors eqe_t^{II} will in tendency remain much closer to zero since model II with time-varying cointegration parameters attributes a part of the magnitude of eqe_t^I to the disturbance of the stable long-run equilibrium relationship. Hence, using state space modeling for regarding the cointegration parameters as time dependent might open an additional dimension into the price dynamics of agricultural markets since the dynamics of the disturbance of long-run price equilibria can explicitly be accounted for.

4 A state space cointegration relationship

We use the state space approach to investigate of the long-run price relationship (3) between the wholesale markets in Tel Aviv and Hebron. State space modeling entails two levels of modeling. First, it models the time series as a function of potentially unobservable components in the so-called *measurement equation*. Second, the components themselves are modeled to follow certain stochastic processes in the *state space* or *transition equations* (for an overview, see Mergner, 2008, ch. 3.5).

The state space model is very rich in flexibility (see for an overview, e.g., Zivot, 2002, or Koopman et al., 2008). The components can be thought in a quite comprehensive way. They may be, similar to the Box-Jenkins approach, trend, seasonal component etc.. Alternatively, they may be, similar to non-parametric approaches, a local level, or, as in the case of regression models, a number of explanatory variables including their coefficients. Some or all of the components can be modeled to be time-varying.

For the given case, if both parameters of the long-run price equilibrium (3) are allowed to be time-dependent, a linear regression model of the following state space form emerges:

Measurement equation:

$$p_t^{A,ob} = \beta_t^0 + \beta_t^1 p_t^{B,ob} + \varepsilon_t \quad \varepsilon_t \approx \text{iid } N(0,\sigma_\varepsilon^2) \tag{8}$$

State space equations:

$$\begin{aligned} \beta_{t+1}^0 &= f_0(\beta_t^0) + \eta_t \quad \eta_t \approx \text{iid } N(0,\sigma_\eta^2) \\ \beta_{t+1}^1 &= f_1(\beta_t^1) + \varsigma_t \quad \varsigma_t \approx \text{iid } N(0,\sigma_\varsigma^2) \end{aligned} \tag{9}$$

The parameters β_t^0 and β_t^1 are allowed to be time-varying in (8). This implies, that the stochastic process they follow has to be specified. A wide range of functional forms for $f_0(\bullet)$ and $f_1(\bullet)$ in (9) can be specified. A simple and quite general assumption for these functional forms is the identity, so that the parameters are basically modelled as random walks. Equations (8) and (9) can conveniently be summarized into the so-called state-space form:

$$\begin{pmatrix} \xi_{t+1} \\ p_t^{A,ob} \end{pmatrix} = \begin{pmatrix} 1 & 0 \\ 0 & 1 \\ 1 & p_t^{B,ob} \end{pmatrix} \xi_t + \begin{pmatrix} \varepsilon_t \\ \eta_t \\ \varsigma_t \end{pmatrix} = \Phi_t \xi_t + \zeta_t \quad \zeta_t \approx \text{iid } N(0,\Omega) \text{ and } \Omega = \begin{pmatrix} \sigma_\varepsilon^2 & 0 & 0 \\ 0 & \sigma_\eta^2 & 0 \\ 0 & 0 & \sigma_\varsigma^2 \end{pmatrix} \tag{10}$$

where $\xi_t = \begin{pmatrix} \beta_t^0 & \beta_t^1 \end{pmatrix}'$ contains the parameters modelled in the state space equation.

Maximum likelihood estimates for the unobserved state variable and the parameters are obtained by applying the *Kalman filter* (Kalman, 1960) which is a recursive estimation algorithm. The procedure is initialized by a set of chosen initial parameters. Each step of the forward recursion consists of two operations. First, predictions of the mean and covariance matrix of the as normally distributed assumed state vector ξ_{t+1} are calculated based on the observations from the first period up to period *t*. Subsequently, the estimates of ξ_{t+1} are updated incorporating the the observation of period *t+1*.Subsequently, the smoothed estimates of all ξ_t are obtained by using the information of the entire sample. This backward recursion is also referred to as *Kalman smoother*.

5 Results

The results of the Augmented Dickey-Fuller (ADF) test for unit roots (Table 3) strongly suggest for all data series except the apple price of the Tel Aviv market that the price series are I(1). However, we follow Juselius (2008, ch. 2)[5] and also regard this series as having a unit root.

Table 3: Results of the Unit Roots Tests

Product	Location	ADF (price levels)	ADF (first differenced prices)
		test statistic[a]	test statistic[a]
Apples	Hebron	-2.331	-12.601*
	Tel Aviv	-3.874*	-19.121*
Cucumbers	Hebron	-2.550	-9.889*
	Tel Aviv	-2.857	-10.852*

Source: Authors' calculations.
Notes:
* Significance at the 5 per cent level, lag length selection according to the Hannan-Quinn Criterion.
[a] The critical value at the 5 per cent significance level is -2.86 for levels and -1.94 for differences.

The results of the Johansen trace test statistic and the Saikkonen-Lütkepohl test statistic in Table 4 show that the price pairs for both commodities are cointegrated at the 5 per cent significance level. We hence regard both markets to be integrated.

Table 4: Results of the Cointegration Tests

Product	Market pair	Johansen trace[a]	Saikkonen-Lütkepohl[b]
		test statistic	test statistic
Apple	Hebron – Tel Aviv	20.03*/ 5.32	21.67*/ 0.47
Cucumber	Hebron – Tel Aviv	32.29*/ 5.71	19.60*/ 2.31

Source: Authors' calculations.
Notes:
* Significance at the 5 per cent level, lag length selection according to the AIC.
[a] The critical values at the 5 per cent level for a cointegration rank of zero/ one are 20.16 and 9.14, respectively.
[b] The critical values at the 5 per cent level for a cointegration rank of zero/one are 12.26 and 4.13, respectively.

As mentioned above, we restrict our interest to the time-varying estimation of the long-run price equilibrium (model (3)), which is estimated for the apple and cucumber price series, respectively. We first estimate model (2) with constant parameters (Table 5) without restricting and with restricting the cointegration relationship ($\beta^{TA} = -1$) using Johansen's reduced rank regression. The parameter estimates are presented in Table 5.

[5] She argues that the unit root property is a useful approximation for empirical analysis. Furthermore, a unit root is not a characteristic of atime series per se but rather depends on the time frame considered.

Table 5: Estimates of the Cointegration Relationships

	Product	β^0	β^H	β^{TA}	p-value of Wald test
Unrestricted	Apple	7.579	1	-1.639	-
	Cucumber	-12.052	1	-0.905	-
Restricted	Apple	3.475	1	-1	0.089
	Cucumber	0.677	1	-1	0.206

Source: Authors' calculations.

Notes: β^0, β^H and β^{TA} are the constant and the coefficients of the Hebron (market *A*) and the Tel Aviv (market *B*) prices, respectively.

Since the Wald test does not provide evidence against the restrictions imposed on the long-run equilibrium we proceed with the restricted model in which the coefficients of the two prices in each equation are restricted to unity in absolute value. Hence, the estimates of β^0 denote simply the average price differentials (the average price margins) between the wholesale markets in Israel and the West Bank for each commodity. Consequently, the resulting estimates of eqe_t^I in (2) represent the short-run deviations from the average price differentials. Based on the restrictions, the state space models which are finally estimated are a simplified version of model (10). The models correspond to a local level state space model:

$$\begin{aligned} p_t^{A,ob} - p_t^{B,ob} &= \beta_t^0 + \varepsilon_t \quad \varepsilon_t \approx \text{iid } N(0,\sigma_\varepsilon^2) \\ \beta_{t+1}^0 &= \beta_t^0 + \eta_t \quad \eta_t \approx \text{iid } N(0,\sigma_\eta^2). \end{aligned} \tag{11}$$

Figures 3 and 4 plot the time-varying estimates and their averages of the margin parameter β_t^0 and the 95% confidence bands for apples and cucumbers, respectively.

In this model setup, the parameter estimates are identical to the smoothed equilibrium errors eqe_t^{II} in (3). The figures illustrate that the equilibrium errors eqe_t^I of model (2) vary around their smoothed counterparts eqe_t^{II} from (3). The observed margin is basically modeled as being composed of a systematic part, i.e. the estimates of the time-varying margin β_t^0, and the white noise error term η_t which is represented by the vertical difference between the solid blue and the black lines. For the given model, the (global) mean of the time-varying state variable β_t^0 represents the long-run equilibrium relationship between the two time series. The estimates of β_t^0 can be interpreted as the systematic part of the deviation from the long-run equilibrium, that is, its temporary disturbance.

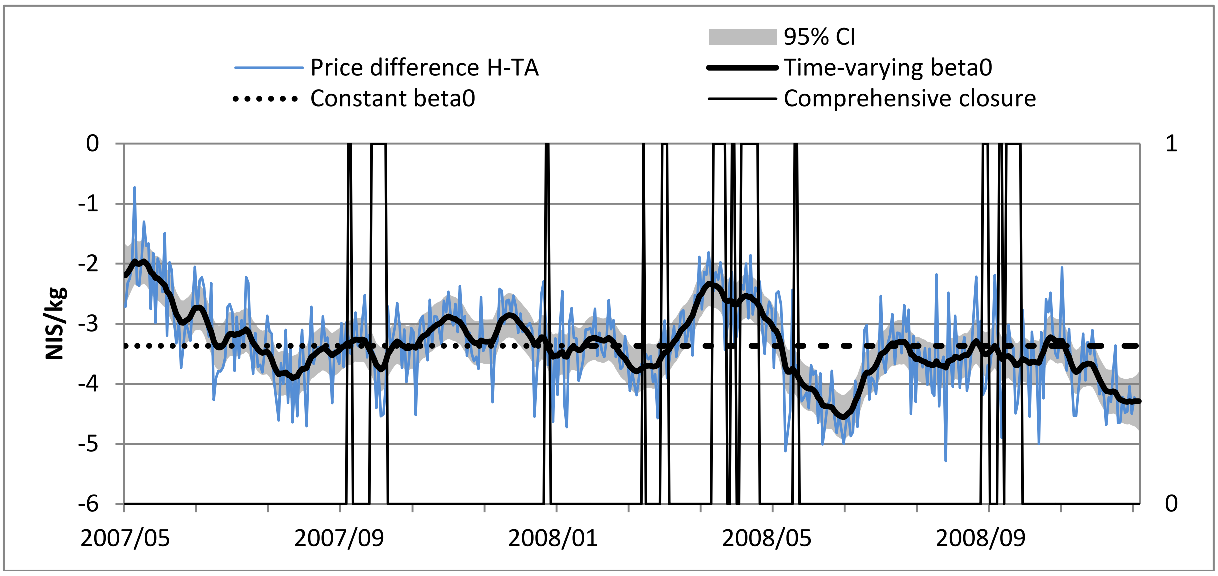
Figure 3 illustrates the differences between the constant parameter and the time-varying parameter approaches. The parameter β_t^0 varies markedly around its mean value of -3.37 NIS/kg[6], although it appears to be reasonable close to it for most periods. In May 2007 and April/ May 2008 the margin is significantly smaller than its average reaching -2 NIS/kg. This is however followed by a significantly increased margin in June/July and December 2008. The estimates of β_t^0 appear to be quite smooth though varying between -4.6 and -2.0 NIS/kg.

[6] Note that the left hand side variable in (11) is calculated as the price difference between Hebron and Tel Aviv with Hebron having considerable lower prices (compare Figure 2).

Temporarily, the margin is trending downwards, upwards or sideward which is interrupted by short periods of clear negative “bumps” which occur every three to four weeks, so for example the two bumps in the negative trend up to the end of July 2007 at the beginning of the observation period. After August 2008, the volatility of the margin is markedly increased.

There is some evidence for a discernable immediate impact of the incidences of the complete closures on the β_t^0 parameters7. The closure incidences in March and May 2007 are characterized by pronounced short term margin decreases which result from a sudden price decrease in the Tel Aviv market and a considerable price increase, respectively (Figure 2).

Figure 3: Smoothed estimates of the time-varying β_t^0 for apples



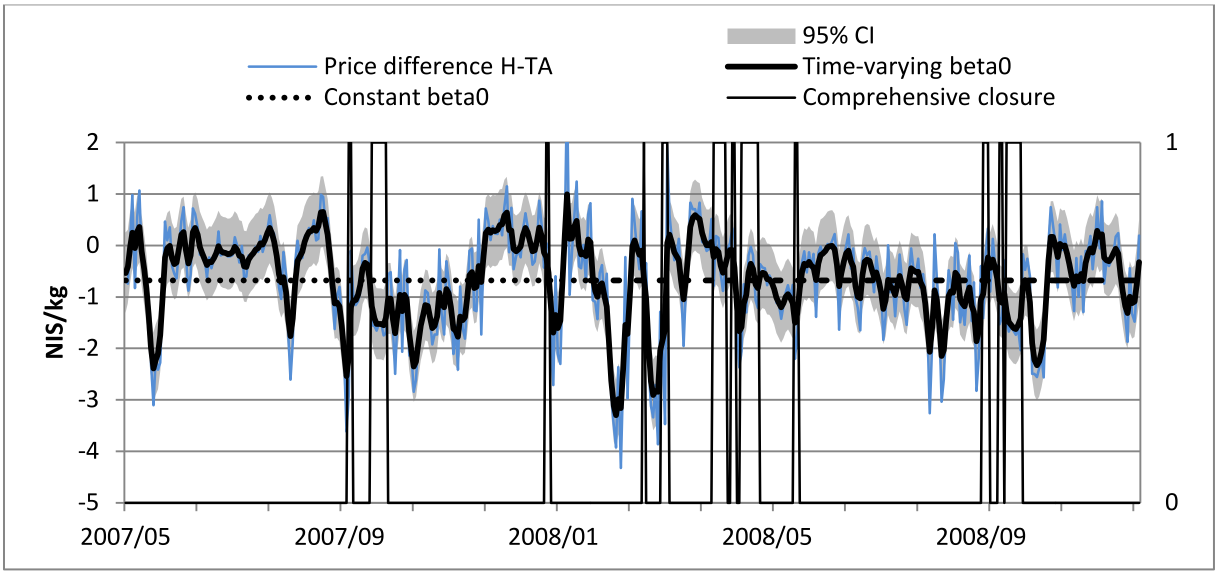
Source: Authors’ calculations.

Note: The axis on the left hand side measures the price. The incidences of the comprehensive closures are indicated by the axis on the right hand side (1 means closure). They were implemented on the following dates: 26.9.-4.10.2007; 9.1.-10.1.2008; 9.3.2008; 19.3.-23.3.2008; 20.4.-27.4.2008; 30.4.-1.5.2008; 6.5.-15.5.2008; 8.6.-9.6.2008; 29.-1.10.2008; 8.10.-9.10.2008; 13.10.-21.10.2008.

Figure 4 contains the estimates of β_t^0 for cucumbers. The time-varying parameter seems to be much less smooth, but to stay somewhat closer around its mean of -0.68 NIS/kg than in the case of apples. Both, negative and positive margins occur varying between 1 and -3.3 NIS/kg. In contrast to apples, no downward or upward trending can be observed. Instead, the relative stability of the parameter is interrupted by strong short termed negative shifts in the magnitude of 2 NIS/kg or more. The closure incidences from January to May 2008 have virtually no missings, so that the short term price responses during this period are reliable. In all cases during this period, the (negative) margin considerably widens during or after the closures, that is, the margin between Hebron and Tel Aviv strongly increases in absolute value (Figures 4 and 5). This is mostly due to the fact that the Tel Aviv price quickly changes (increases) strongly while the Hebron price does not show pronounced changes (see Figure 1).

[7] Since the closures in September 2007 and September/ October 2008 coincide with several missing values in the price series which were imputed, we refrain from interpreting these periods.

Figure 4: Smoothed estimates of the time-varying β_t^0 for cucumbers



Source: Authors' calculations.

Note: The scale on the left hand side measures the prices. The incidences of the comprehensive closures are indicated on the right hand side by unity.

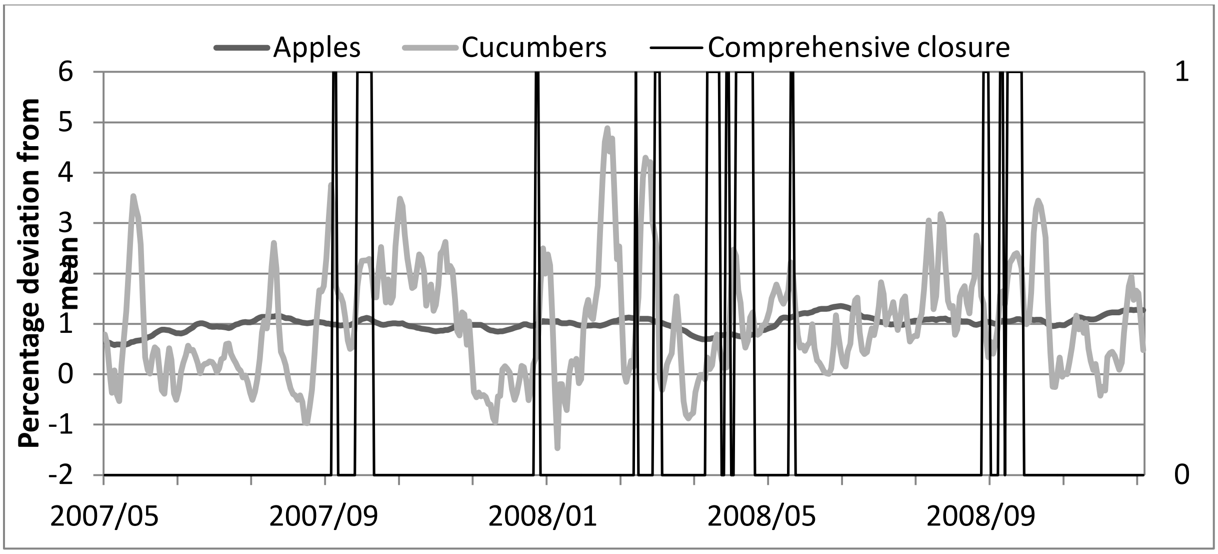
The time-varying margin estimates of apples and cucumbers are compared in Table 6 and Figure 5. The coefficient of variation is substantially higher for cucumbers than for apples. The cucumber margin shows extraordinary strong fluctuations around its mean, i.e., its long-run equilibrium value, while the variations of the apple price margins are small. The estimates of the skewness and kurtosis indicate that the estimates of both models are reasonably normally distributed.

Table 6: Characteristics of the β_t^0 Estimates

	Apples	Cucumbers
Mean	-3.37	-0.68
Median	-3.4	-0.56
Maximum	-2.0	1.0
Minimum	-4.6	-3.3
Standard deviation	0.5	0.8
Coefficient of variation	-0.2	-1.1
Skewness	0.4	-0.8
Kurtosis	3.5	3.4
Final state	-4.3	-0.3
Observations	437	437

Source: Authors' calculations.

Figure 5: Development of the time-varying β_t^0 parameters relative to their means



Source: Authors' calculations.

Note: The normalized prices are measured on the left hand scale. Closure incidences are indicated on the right hand scale by unity.

6 Conclusions

Economic relationships face often challenging conditions in the conflict of violent political conflict. We assess the economic side effects of security measures implemented in response to security threats occasionally arising due to the conflict on agricultural trade and price relationships. We focus on the effects of abruptly implemented movement restrictions in the context of the Israeli-Palestinian conflict. In particular, we assess the effects of temporary complete closures of the West Bank barrier on the price relationships between Israel and the Palestinian territories because they prevent agricultural trade on a daily basis. We investigate the long-run price relationship between the fruit and vegetable wholesale markets in Hebron and Tel Aviv. We use daily data of the two of the most traded products (apples and cucumbers) between May 2007 and December 2008.

We assess the price interdependencies by employing a cointegration model. However, the pronounced abrupt and short-term character of the trade impediments suggests a modeling approach which is able to capture such fierce, but temporary disruptions of economic relationships. Therefore, globally constant parameters as assumed in the traditional cointegration framework appear to be an inadequate choice. Instead, we suggest a varying parameter cointegration model estimated using the Kalman filter in order to be able to explore the short-run impact of the security measures on price interdependencies. We model the cointegration relationship as time-varying because the interest of this paper focuses on temporary disturbances of the long-run price equilibria. Based on the Johansen estimation procedure, we do not find evidence against restricting the slope coefficients to unity in absolute terms. The final models hence estimate the time-varying margins between both wholesale markets for each commodity.

The results for (mainly in Israel produced) apples suggest that the margin markedly decreases due to the rise of the apple price on the Hebron market which tends to be below the Tel Aviv price. This appears to be plausible due to the price response due to the decrease of the apple supply on the Hebron market as consequence of the prevented trade by the complete closures. On the Tel Aviv market, the price response will be downward due to the increased supply. The results concerning the cucumber prices are in the opposite direction since this commodity is mainly produced in the West Bank. The estimates of the margin strongly widen during or immediately after closure incidences. Again, this result appears to be plausible since the excess supply which cannot be shipped through the West Bank barrier to Israel leads to slight price decreases on the Hebron market while Tel Aviv prices respond with strong rises.

Palestinian consumers are harmed by increased levels and volatility of apple prices but profit from the increased supply of cucumbers during the implementation of the security measures. The total welfare effects depend on the magnitudes of the price responses for each product in Hebron. However, Palestinian traders may try to take advantage of increased apples prices by stretching their profits. Israeli consumers benefit from the closures in the case of apples which seem to be outweighed by the stronger negative effects of the closures concerning cucumbers.

The suggested method of a time-varying cointegration model is found to offer comprehensive additional insights into price dynamics since it estimates one expected value of each parameter of interest for each time period considered. Market events which exert a strong impact on price relationships can be identified by this modeling technique. In the given context, the timings, magnitudes and durations of such disruptions of the long-run price equilibria appear to be very informative. The framework also offers a large flexibility concerning the parameters allowed to be time-varying.

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