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Economic integration across borders: The Polish interwar economy 1921–1937

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In this article we study the issue of cross-border economic integration in the context of Poland's reunification from 1919. We find that the Polish interwar economy can be regarded as integrated with some restrictions. Moreover, a significant negative impact of the former partition borders on the integration level in the early 1920s vanishes in the middle of the 1920s. This suggests that the post-reunification integration was surprisingly successful. The degree of integration is comparable to that of 19th century France.

1. Introduction

What happens to an economy if we change its borders? Since administrative and political borders tend to imply massive barriers to trade, information, and mobility, the removal of such borders should lead to better economic integration. However, the literature on integration due to changing borders remains inconclusive on these issues. Some authors find that massive barriers to integration remain even after borders are removed (for example, Engel and Rogers 1996). Others measure a high degree of integration across borders before their removal (for example, Moodley *et al.* 2000).

In this article we investigate in great detail how economic integration evolves across removed borders in a possibly unique historical setting: the reunification of Poland after the First World War. Already at the end of the 18th century, Poland had been partitioned among Tsarist Russia, the Habsburg monarchy, and the emerging Prussia. When Poland reappeared on the map of Europe in 1919, it consisted of three parts that were dramatically different with respect to their institutional framework (currencies, taxes, laws), and isolated from one another due to high transportation and communication costs. Accordingly, all Polish governments after 1919 attempted to unify and integrate the country. The Polish Statistical Office (GUS) monitored these efforts from 1921 until 1937 with respect to the price movements of several basic commodities, publishing monthly prices for all parts of the new state. Since the data are given in a single currency and originate from one single source we can exclude noise from exchange rate volatilities or data definitions that often plague cross-border studies.

Hence, we are in a position to evaluate the process of integration across the former partition borders of Poland. To this end, we draw on two strands in the literature and show that they can complement each other. Specifically, we consider the measurement of border effects in the tradition of Engel and Rogers (1996) and the use of threshold cointegration analysis following the examples of Balke and Fomby (1997) and Lo and Zivot (2001). Both approaches are based upon a definition of economic integration in terms of spatial arbitrage, to which the literature refers as the 'weak form of the Law of One Price (LOP)'. The weak LOP states that 'the prices of a homogeneous good at any two locations will differ by, at most, the cost of moving the good from the region with the lower price to the region with the higher price' (see Fackler and Goodwin 2001, p. 978). Hence, for a panel of price data in an economically integrated area this weak LOP should hold. However the two approaches differ with respect to the dimension of the panel on which they focus. Consider the approach of Engel and Rogers (1996). Since arbitrage is expected to keep price differentials between locations within the limits of transport and communication costs, the volatility of these price differentials can be a proxy for the degree of integration between locations. Engel and Rogers (1996) propose, then, to regress the cross-section of bilateral price volatilities on distance, location specific effects, and a border dummy variable. Obviously, if one estimates a positively significant coefficient on a border dummy, the border matters insofar as it reduces the degree of integration between locations divided by that border. The idea is appealing for large cross-sections, and within our regression specification we can deal with structural instability over time. However, it does not fully exploit the evidence on prices. A high volatility of price differentials could be caused by the absence of price adjustment between the locations but may also be caused by higher transportation and communication costs. Thus, on the intertemporal dimension the approach cannot distinguish between the situation where price differentials are volatile because prices do not adjust at all, and that where prices respond with a low speed or adjust only to some high level of transaction costs. Threshold cointegration models as discussed in Balke and Fomby (1997) and Lo and Zivot (2001) allow one to make these distinctions and to analyse the dynamics of bilateral price differentials in great detail.[1] However, this approach becomes very labour-intensive for

[1] Volatility in transaction costs can be a further reason for (higher) price volatility. However, within our threshold cointegration approach we have to assume constant transaction costs.

large cross-sections and shares the usual problems of time series analyses with respect to structural instability. Therefore, in order to fully exploit the information contained in our data, we will consider the two model frameworks as complementary to each other and apply both to the case of interwar Poland. In line with the large historical literature on economic integration, we explore this issue focusing on the grain market (see Persson 1999).

The rest of the article is organised as follows. In the next section we present the historical background of the study and describe our data set. Section 3 introduces the two econometric model frameworks and discusses their relationship. The empirical results are presented in Section 4. The last section summarises our findings and concludes. The Appendix contains a more detailed discussion of the econometric methods we have applied.

2. Historical background and data

Between 1772 and 1795 the Noblemen's Republic of Poland (*Rzeczpospolita Polska*) was split among the empires of Tsarist Russia, the Habsburg monarchy, and emerging Prussia. As a consequence of the partitions – 'the first very great breach in the modern political system of Europe' (Edmund Burke) – Poland disappeared from the map. Only the specific constellation of factors operating at the end of the First World War, when all three partition powers were severely weakened through war and revolution, paved the way for its restoration. Owing to the long period of partition, there arose in each region differing legislation governing virtually all aspects of social, political and economic life. The government in fact could rely on extensive programmes for legal, administrative, and economic unification that had been prepared since 1907 for an eventual reunification. However, the agenda was not guided by any political or economic 'master plan', but rather by the ongoing war that Polish troops fought with the Soviet army in the east (see Landau 1992, Roszkowski 1992).

This war required massive outlays and some mechanism to finance them. Since international credit was not available – the Paris peace conference did not start before January 1919 and Poland was yet to be formally recognised as a state – the government had to choose between the expropriation ('nationalization') of domestic private capital and ways to tax it (Landau and Tomaszewski 1999). The political compromise in 1919 relied on early concessions to the socialists on the one hand and observing private property rights on the other. As a consequence, the next step was to create the institutional framework necessary to tax capital and labour: a common currency and a working fiscal administration. The unification of the fiscal administration belonged to the very first institutional changes. While for most of the former Austrian and Russian parts this was already formally reached by April 1919, the former German parts remained separated until January 1922, and (Upper) Silesia even until June 1922 (Markowski 1927, Bielak 1931). A common income tax was decreed in July 1920, but it took several years to implement it in the former Russian territories. Business taxes in turn were introduced and unified throughout the whole territory by July 1925, following the Russian system of business certifications. Nevertheless, some differences in the tax system – for example, the real estate tax – persisted until 1936 (Weinfeld 1938).

Next, an important precondition of a functioning tax system would be the creation of a common currency area, namely the unification of the five (!) currencies that were in circulation in Polish territory: the German Mark, the Austrian Crown and the Russian Rouble, as well as the Polish Mark in the Kingdom of Poland and the 'Ost-Rubel' on the territory of 'Ober-Ost' – two currencies that the Germans introduced on former Russian territories after their occupation. Since the Warsaw government only controlled the Polish Mark, it adopted a stepwise strategy to get rid of the competing banknotes (Landau 1992). Some months after the introduction of the Polish Mark as a parallel currency in the different areas, the other currencies were withdrawn. For most of the Polish territory with the exception of Upper Silesia (November 1923) this was already realised as from April 1920 (Zbijewski 1931). While such a quick institutional change was an indisputable success, it could not create the necessary revenues to win a war. Nevertheless, it opened the way for the Polish government to effectively tax money holders via inflation. As estimated by Zdziechowski (1925) the money supply increased by 519 per cent between 1918 and 1919 and in the following year by another 929 per cent reaching in 1923 more than 12,000,000 per cent of the 1918 level. Obviously, the temporary gains from seigniorage and the devaluation of the budget deficit were quickly wiped out by the costs of hyperinflation, namely the loss of access to foreign capital. When Prime Minister Władysław Grabski tried to stabilise the currency in 1924, his specific aim was to link the Polish currency with some foreign currency that had successfully restored the gold standard in attempting to gain access to the international capital market. Indeed, Grabski managed to realise this task with the help of a temporary property tax fixed in Swiss gold francs and several international loans. By mid-January 1924 the nominal exchange rate had been stabilised and a new currency, the Zloty, was fixed to the Swiss gold franc. After a second wave of devaluations the exchange rate finally stabilised at a sustainable level around May 1926. From then on the government started to defend the parity at any cost.

The war in the east also induced a rapid improvement of the transportation system, since it required a network to transport men and material. After rather spontaneous takeovers of the railway networks in the different areas during the last months of the First World War, a Railway Ministry started its work in October 1918 and developed a ten-year plan for the completion and extension of the Polish railway network. At the same time the heritage of 129 types of cars and 165 types of engines had to be unified, new kinds of freight cars had to be developed (such as refrigerator wagons), the different densities of the network adjusted and the main economic centres of the former partition areas connected (Hummel 1939). The speed of the network and its capacity to transport goods was not only a function of the existence of railway connections themselves but also crucially depended on the material used. For example, the opening of a new connection between Warsaw and Krakow in 1934 did not only reduce the journey distance between the two cities by 12 per cent, but also reduced the travel time by 33 per cent (see Pisarski 1974, p. 58). Since nearly all freight transport over 50 km took place on railways with normal gauge (97.6 per cent in 1925 and 98.7 per cent in 1938 (see Brzosko 1982, p. 358) this development of the railway network can be expected to have had a strong integrating impact on the economy.

Hence, the most obvious non-tariff barriers to trade and mobility within the new Polish state, such as different currencies, different tax systems and a shortage of transport facilities, were considerably reduced if not completely removed by 1926. Moreover, as one of the first steps to unify the new economy, a common external tariff was introduced in November 1919. However, it took a little longer to get rid of internal tariffs and a system of widespread regulations of commodity and factor markets. This system was again motivated, in part, by the need to equip the Polish troops fighting the Soviet army in the east, but it also had aspects of political logrolling between different groups. In particular, the markets for agricultural products (such as bread, grain, potato, and sugar) and basic commodities (such as coal, soap, and matches) were affected by a variety of measures that discriminated between regions and social groups. For example, there remained a customs frontier between the former Prussian partition area and the rest. This kept grain prices in the Prussian area artificially low, thereby providing cheap supplies for the fighting troops (Kozlowski 1989, Landau and Tomaszewski 1999). After the armistice between Poland and Soviet Russia, the Polish government launched a programme to do away with the whole system of regulations. The internal customs frontier was removed in mid-1921 and by the end of 1921 most other regulations on the commodity markets had disappeared (Tomaszewski 1966).

To sum up, Poland after 1919 was characterised by a multitude of barriers to trade, information, and mobility, which may have given rise to border effects in line with the arguments of Engel and Rogers (1996). Between 1919 and 1926, Poland made massive efforts to remove these barriers that had divided the country for more than a century. But to what extent did this process of institutional change result in better economic integration across the former borders? To analyse this issue, we use monthly retail prices from several publications of the Polish Statistical Office (GUS) in Warsaw covering the period from January 1921 to December 1937.[2] All prices were reported to the GUS by the city administrations, for 1921–25 as monthly averages, for 1926–37 as prices of the last week in a month. We have evidence over the complete period for the cities of Warsaw, Kráków, Łódź, Lwów, Poznań, and for Wilno from 1924 onwards. This allows us to distinguish between the formerly Russian area (Warsaw, Łódź, Wilno), the formerly Austrian area (Kraków and Lwów) and the formerly German area (Poznań). Due to the currency reform in 1924 following the period of hyperinflation, the GUS published all price series until June 1923 in the Polish Mark and, beginning in January 1924, in the new currency, the Zloty. Therefore, we split our sample at June 1923 and obtain a first subsample including 30 observations (January 1921–June 1923) for a total of 10 city pairs and a second subsample including 160 observations (January 1924–April 1937) for 15 city pairs.

The GUS provided series of retail prices for several basic commodities including coal, soap, vegetables, and wheat flour. Some of these markets, however, were at least temporarily subject to high levels of concentration with prices set by inter-regional cartels rather than by competitive arbitrage traders. The analysis of such markets may be an interesting issue in its own right, since market power might reduce market integration through spatial price discrimination. However, this is beyond the scope of this article and we leave it for future analysis.

We choose to focus on the market for wheat flour (milled at the same grade) in different cities, for three principle reasons. First, the historical record does not show any major market concentration in this market, either at the level of wholesale trade or at that of retail trade. This is of crucial importance, since we use retail prices, which reflect inter-regional integration due to arbitrage only under the assumption of a competitive market.[3] As most of the wheat grinding was done in small mills that were evenly spread around the whole country, we can speak of a dense, decentralised network of mills from which the flour was shipped to the cities. Hence, we can assume that the prices for wheat flour were the outcome of a competitive market with arbitrage, adjusting for large price differentials between different locations.

Second, we have direct evidence that there were positive trade flows in wheat between the various parts of Poland under consideration, which is a necessary condition for arbitrage since price co-movements may

[2] The Głowny Urząd Statystyczny [Main Statistical Office] published this price series for 1921–1928 in its *Rocznik Statystyczny* [Statistical Yearbook]; the series for 1929–1937 were comprised in a publication *Statystyka Cen* [Price Statistics], published monthly for 1929 and quarterly for 1930–37.

[3] One may prefer to use wholesale prices to analyse market integration. But in the case of competitive wholesale and retail markets the retail prices form a stable link to wholesale prices.

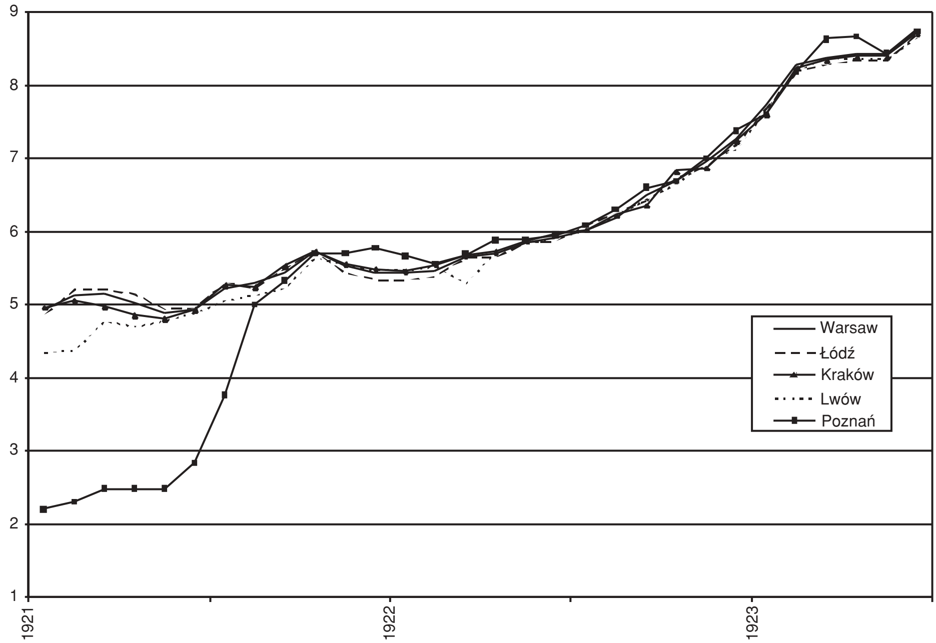

Figure 1. *Logarithm of wheat flour prices 1921–1923.*

well be caused by other factors, such as common climatic conditions.[4] Third the use of wheat prices fits into the historical literature on price integration, which is mostly based on the history of grain prices (see Persson 1999).

The log-price series are shown in Figures 1 and 2. Figure 1 shows the strong price increase due to high inflation at that time. Moreover, the lower flour prices in 1921 in Poznań are apparent, which reflect the regulated prices in the former Prussian area. Figure 2 refers to the period January 1924–December 1937, which is used for the threshold cointegration analysis. In line with general price developments in Poland, wheat flour prices increased until 1927 followed by a short period of stabilisation. Then, in line with the Great Depression, prices fell dramatically from 1929 onwards. This change from an inflationary to a deflationary environment may lead to possible breaks in the deterministic components of the time series. We have to address this issue within the threshold analysis. To determine a common breakpoint we refer to the food price index (FPI) since break dates may be different among the price series. The FPI is the most complete price index available for interwar Poland on a monthly basis. It consists of 16 agricultural goods wherein wheat accounts for slightly less than 5 per cent of the index. As a break date we choose the observation May 1929, since from this month

[4] The trade figures have been documented in Ministerstwo komunikacji (1925–1937) and are discussed in Wolf (2003).

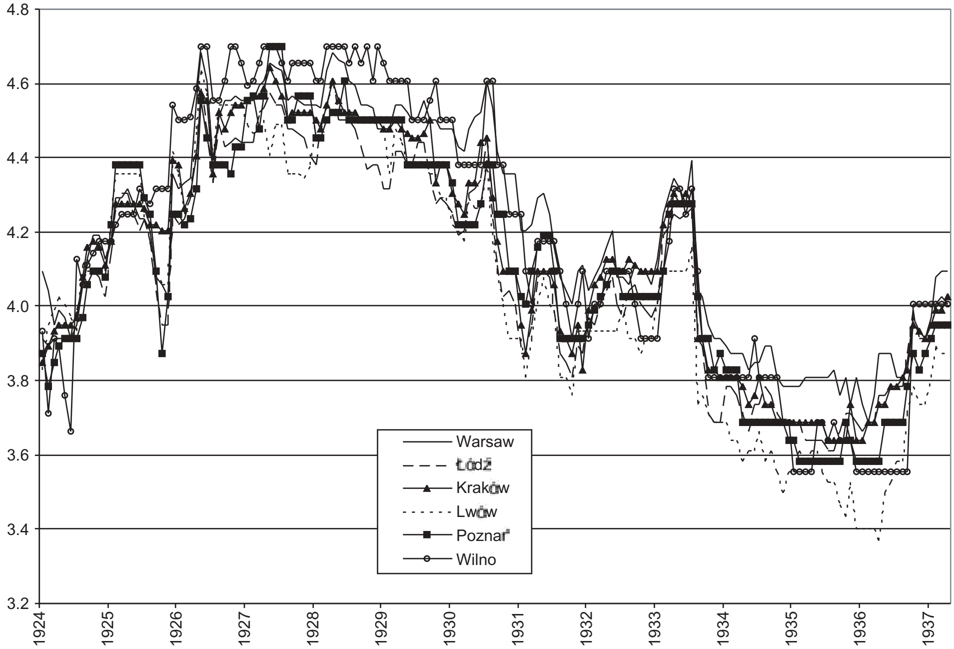

Figure 2. *Logarithm of wheat flour prices, 1924–1937.*

onward the FPI started to fall. A Chow breakpoint test confirms this break date. Accordingly, we have 64 and 96 observations respectively for the two subperiods.[5]

For the analysis of possible border effects according to the Engel-Rogers approach we can also take account of the data from January 1921–June 1923. As will be explained in the next section, our regression specification allows us to handle periods of missing data and structural differences between subperiods.

3. Economic model framework and econometric methods

In our study we measure economic integration by the LOP adjusted for transaction costs. In its strict form, the LOP states that the prices of the same good should not differ at two spatially separated market places if these markets are integrated. When the prices differ, arbitrage processes in a functioning integrated market would instantaneously equalise these prices. In the presence of transaction costs, arbitrage is only profitable and

[5] The cointegration framework employed later on does not allow for individual break points in the series of a multivariate system. We do not consider different break dates for the 15 price systems, since the results of the cointegration analysis are robust to slight changes in the breakpoint May 1929.

takes place if the price deviations between the two market places exceed the transaction costs.

Let us formalise the LOP more precisely. Consider two market places i and j, letting N_{ji} denote some export level of a good from place j to i. P_{it} and P_{jt} are the respective price of the good in locations i and j. Assume further that the transaction costs are proportional to the prices in the importing market place. In line with the recent economic geography literature let $(1 - e^{-\tau})P_{it}$ be the transaction costs, where $\tau > 0$ is a cost parameter. Then, $e^{-\tau}P_{it}$ is the per-unit revenue when the good is sold in location i. Intuitively, τ depends positively on the geographical distance between the locations i and j. Moreover, when border effects are present, τ also differs depending on whether or not the locations lie in the same area. Finally, trade from j to i is only profitable if $P_{it}N_{ji}e^{-\tau} > P_{jt}N_{ji}$. This results in the condition: $\log(P_{it}/P_{jt}) = p_{it} - p_{jt} > \tau$. Hence, arbitrage from j to i takes place when the log-price difference $p_{it} - p_{jt}$ is larger than the cost parameter τ. Equivalently, one trades from location i to j only if $p_{it} - p_{jt} < -\tau$. Thus, we obtain $[-\tau; \tau]$ as a band of no-arbitrage. Within this band, no trade occurs that could reduce price differences between the two markets because transaction costs exceed possible arbitrage profits. Obviously, the size of this band increases with τ. In the literature, the transaction cost view of the LOP is often referred to as a weak form of the LOP. It is equivalent to the so-called spatial arbitrage condition if it is only required that prices of the same good at two locations differ at most by the transaction costs (see, for example, Fackler and Goodwin 2001).

We have to be aware that there exists little agreement in the literature on the use of the term *integration* (for an overview, see Fackler and Goodwin 2001). Fackler and Goodwin (2001, p. 978) suggest using the term *market integration* ‘as a measure of the degree to which demand and supply shocks arising in one region are transmitted to another region’. In this sense, a market integration does not require a one-to-one price movement. A complete transmission of shocks, however, implies the LOP. On the other hand, we may observe the LOP holding even if two locations are not integrated in terms of shock transmission. In fact, by referring to the LOP we measure integration by price convergence. Hence, the LOP and the concept of market integration as used by Fackler and Goodwin (2001) have to be distinguished. Therefore, we refer to the term *economic integration* in our framework.

Both the framework of Engel and Rogers (1996) and the threshold cointegration analysis can be associated with the LOP. In the following, we first comment on Engel and Rogers’ approach. Then we describe the main steps of the threshold cointegration analysis and discuss the relationship between both model frameworks. The exposition in the next subsections is mainly descriptive. More technical details on the econometric methods and specifications applied can be found in the Appendix.

3.1. The pooled regression approach

Engel and Rogers (1996) suggest computing the standard deviation of $\Delta_{zt} = \Delta(p_{it} - p_{jt})$ over the whole sample for each city pair. These standard deviations are considered to be a measure of the volatility of the price differences. Then, they regress the standard deviations on the log of the distance between the cities, city-specific dummy variables and a border dummy variable indicating whether or not the border is crossed. Obviously, the border dummy variable summarises all possible effects and costs associated with crossing a border. Thus, border effects are present if the border dummy coefficient is positively significant implying higher variation in the price differences. The Engel-Rogers approach is linked to the LOP in the sense that larger transaction cost bands allow for more price dispersion. Accordingly, two locations with higher price volatility are characterised by higher transaction costs and a lower degree of integration.

The description of the development of the Polish interwar economy suggests that the importance of the former partition borders may have changed over time. Therefore, we pick up Engel and Rogers' idea of splitting the sample to allow for time varying border effects. In practice, this is done by introducing time-specific border dummy variables. A previous analysis with yearly border dummies indicates significant border effects in the years 1921 to 1924. The incorporation of a dummy variable for each year, however, requires the estimation of a large number of parameters. This may lead to high estimation uncertainty. Therefore, we propose using only two time-dependent border dummy variables: one for the period 1921–24 (B_{2124}) and one for 1925–36 (B_{2536}). To be precise, we compute the standard deviation of Δ_{zt} for each year, that is, over 12 months. Let $V(z_t^{i,j})$ denote this standard deviation for city pair (i, j) for year $t(t = t_1, \ldots, T)$. Then, we regress $V(z_t^{i,j})$ on B_{2124} and B_{2536} instead of a single border dummy and obtain the regression equation

$$V(z_t^{i,j}) = \beta_1 d^{i,j} + \beta_2 B_{2124} + \beta_3 B_{2536} + \sum_{m=1}^{6} \beta_{4m} C_m + u_t^{i,j} \;\; \forall i,\, j \text{ with } i \neq j, \tag{3.1}$$

where $d^{i,j}$ is the log of the distance between the cities i and j and the C_m's are city specific dummy variables, which take on a value of 1 for cities i and j if $m = i$ or $m = j$. The border dummy variables B_{2124} and B_{2536} are equal to one for the respective period if a border is crossed and zero otherwise. If $\beta_2 > 0$ or $\beta_3 > 0$, crossing the border increases price variation. We can trace possible changes in the importance of border effects through a comparison of β_2 and β_3. Regarding the error terms $u_t^{i,j}$ we allow for year-wise heteroscedasticity, hence $u_t^{i,j} \sim (0, \sigma_t^2)$. Accordingly, the unknown parameters $\beta_i, (i = 1, 2, 3), \beta_{4m}(m = 1, \ldots, 6)$ are estimated by a feasible GLS procedure.[6]

The specification (3.1) also enables us to take the period January 1921–June 1923 into account since the computation of standard deviations only for subperiods does not require a continuous data set as a time series approach would. We will treat the period January 1923 to June 1923 as a full year. Adding these six months to the year 1922 does not change our results in any important way. However, the data for January to April 1937 are excluded since this period is too short for a reasonable computation of volatilities. Note that the choice of yearly subperiods makes it more likely to have homogenous data for the computation of price variations. In addition, by allowing for heteroscedastic error terms we can deal with structural heterogeneity over time.

Another possibility for modelling time-varying border effects without estimating a large number of parameters is to refer to a smooth transition regression model as proposed by Lin and Teräsvirta (1994). Considering a transformation of the exponential function this regression framework allows one to describe a smooth change in the regression parameters. In our application, we are able to check whether a gradual transition from strong border effects in the early 1920s to no border effects from the mid 1920s onwards has occurred. As mentioned, this seems to correspond well to the results in Trenkler and Wolf (2004). The corresponding model will be estimated by nonlinear least square methods. The technical details are described in the Appendix.

Keep in mind that the impact of linear trends within the Engel-Rogers approach are eliminated because the standard deviations are computed with respect to the first differences of $p_{it} - p_{jt}$. Even if a linear trend does not cancel out when subtracting the prices, it does vanish after the first difference is taken. Similarly, the impact of breaks in the deterministic components diminishes due to differencing. Obviously, the presence of deterministic trends in relative prices would imply deterministically diverging prices and, hence, be at odds with economic integration. However, it is not possible to analyse deterministic price divergence using the Engel-Rogers approach since one has to eliminate trends and breaks for a reasonable computation of standard deviations. This creates another motivation to complement their approach with a time series framework.[7]

[6] We have first estimated (3.1) by OLS in order to obtain estimates for the yearly error variances, which are then used for the GLS estimation. A SUR approach is not possible since we would have to determine too many unknown covariances. This problem occurs because the number of years is higher than the number of city pairs (cf. for example, Griffiths *et al.* 2002).

[7] See also the corresponding discussion regarding the threshold cointegration analysis later on.

3.2. Threshold cointegration framework

The threshold cointegration analysis is directly motivated by the transaction cost view of the LOP. In fact, the latter is econometrically translated into a so-called threshold cointegration model (see Lo and Zivot 2001). The simplest possible form is the following symmetric three-regime BAND-threshold autoregressive (BAND-TAR) model.[8]

$$\Delta_{z_t} = \begin{cases} (\alpha - 1)(z_{t-1} - \tau) + \eta_t, & \text{if } z_{t-1} > \tau, \\ \eta_t, & \text{if } -\tau \le z_{t-1} \le \tau, \\ (\alpha - 1)(z_{t-1} + \tau) + \eta_t, & \text{if } z_{t-1} < -\tau, \end{cases} \tag{3.2}$$

where $z_t = p_{it} - p_{jt}$ is the log-price difference at time t and $\eta_t \sim$ i.i.d. $(0, \sigma^2)$.

The threshold band $[-\tau, \tau]$ corresponds to the band of no arbitrage in which no adjustment takes place. Hence, the log-price difference z_t evolves like a random walk, that is, Δz_t does not adjust to disequilibria ($\Delta z_t = \eta_t$). The limits of the threshold band are described by the thresholds which coincide with the transaction cost parameter. Therefore, they are also labelled as τ. In the outer regimes, for which we have $|z_t| > \tau$, economic forces move the prices back toward the price parity equilibrium. Since we use variables in logarithms, this suggests that the log-price series are cointegrated with a cointegrating vector (1, -1). In other words, the log-price difference forms a stationary relationship. Of course, adjustment ceases at the edges of the band of no arbitrage and does not continue until price parity because transaction costs are larger than the price difference. Thus, the relationships $z_{t-1} - \tau$ and $z_{t-1} + \tau$ enter the outer regimes of equation (3.2). To assure appropriate adjustment $-1 < \alpha < 1$ is required. This causes price differences to fall ($\Delta z_t < 0$) if $z_{t-1} > \tau$ and to increase ($\Delta z_t > 0$) if $z_{t-1} < -\tau$.[9] The autoregressive coefficient α measures the speed of price adjustment. The smaller the value of $|\alpha|$ the faster is the adjustment owing to price disequilibria. The parameter α can also be related to the so-called half-life $h = \ln(0.5)/\ln(|\alpha|)$. The half-life states the number of periods required to eliminate one-half of a deviation from price-parity. Obviously, a smaller value of $|\alpha|$ translates to a shorter half-life. Finally, note that the transaction cost view suggests symmetry regarding the adjustment coefficient α and the threshold τ since arbitrage should be induced in the same way, independent of where the prices are higher.

So far we have just considered the simple log-price difference $z_t = p_{i,t} - p_{j,t}$. The presentation of the data in the foregoing section has shown that the series may be characterised by a broken linear trend and different levels corresponding to the succession of inflationary and deflationary

[8] For a more general discussion of threshold models, see Lo and Zivot (2001) and Balke and Fomby (1997).

[9] Of course, we have to assume that η_t is zero for this thought experiment. In the general case of non-zero error terms, the described adjustment is only assured in the long run.

periods. The question is whether these broken or other deterministic terms have to be included in the price relationship to obtain stationarity. As long as both price series contain the same deterministic terms with the same slope and level, they cancel out when the series are subtracted. But if the deterministic parts differ, one may consider the extended relationship $z_t^* = p_{1,t} - p_{2,t} + \psi d_t$ instead of z_t, where the relevant deterministic terms are collected in d_t. The inclusion of deterministic components in z_t^*, however, has important economic implications. If z_t^* contains a linear trend or a broken trend, the prices in two market places deterministically diverge. Obviously, this contradicts economic (price) integration even if prices cointegrate, that is, if z_t^* is a stationary relationship. The situation is different for a constant or broken constant entering z_t^*. In this case, the prices still converge but not towards price parity. Instead, prices differ by some fixed (deterministic) amount in equilibrium. We may associate such kind of price convergence with a relative version of the LOP.

Deterministic price differences could be due to different local selling and buying costs which may have their origin in different wage and rent costs. This indicates that certain markets, for example, the labour market, are not perfectly integrated or are characterised by rather high transaction costs. We use retail prices in cities which may be quite strongly affected by regionally varying cost components like wages and rents. Note that deterministic price differences have to be distinguished from asymmetric transaction costs. Transaction costs refer to the occurrence of adjustment, but deterministic price differences affect the equilibrium towards which adjustment takes place. Since the occurrence of deterministic terms is important for inference on economic integration we address this issue in the empirical analysis.

In the literature, some objections have been raised as to the usefulness of the (threshold) cointegration concept to study integration (compare, for example, Fackler and Goodwin 2001, Baulch 1997, and McNew and Fackler 1997). We discuss the two main issues at this point. First, it can be shown that cointegration is an unnecessary condition for integration in the case of nonstationary transaction costs. Second, cointegration is not sufficient if the transaction costs are very large, such that all (or most) of the price differentials are within the transaction cost band. Then, no trade would take place. However, we observe trade flows between the different Polish regions. Thus, the second objection is of less importance in our empirical framework. The first issue is more complicated to deal with because we can never rule out nonstationary transaction costs. In fact, the assumption of constant transaction costs is also tested within the threshold cointegration framework. Hence, cointegration failure could result from nonstationary transaction costs. However, the situation is different if cointegration is detected. Following Fackler and Goodwin (2001), cointegration can then be interpreted as a sign of integration. This conclusion is most appropriate if transaction costs are low relative to the prices analysed. Given the weak evidence for threshold nonlinearity in our empirical study, we may infer, as done in Section 4, that the transaction costs are relatively low. Therefore, we are confident that the (threshold) cointegration framework is an appropriate tool for studying economic integration within our empirical application.

The threshold cointegration analysis for each city pair is performed in three steps according to Lo and Zivot (2001) and Balke and Fomby (1997). First we test for cointegration, then for threshold nonlinearity. Finally, the threshold models are estimated.

To test for cointegration we apply a generalisation of the multivariate Johansen testing procedure, which allows for broken linear trends and levels. This generalisation has been proposed by Johansen *et al.* (2000). It enables us to test which deterministic components affect the price cointegration relationship in line with the discussion above. Additionally, we can test whether the cointegrating vector can be restricted to $(1, -1)$ so that the log-price difference is, in fact, the relevant quantity for price adjustment. The basic idea of the Johansen procedure is to sequentially test for the number of cointegration relations among the variables of interest, starting with the null hypothesis of no cointegration relation. We denote this null hypothesis as $H_0 : r_0 = 0$. If it is rejected, one continues checking $H_0 : r_0 = 1$ and so on until a certain number of cointegration relations cannot be rejected. Since we consider bivariate systems, at most two test steps can occur. Further details on the estimation and testing procedures are deferred to the Appendix.

Provided we have found cointegration between the price series, we continue to test for threshold nonlinearity regarding the log-price difference $z_t = p_{1,t} - p_{2,t}$ with respect to each city pair.[10] To this end, several procedures are applied. They all test the null hypothesis of linearity against a threshold alternative. We first use the univariate and multivariate tests suggested by Tsay (1989, 1998). The idea of these procedures is to perform an arranged autoregression of order k based on data that are ordered according to the value of the threshold variable (in our case z_{t-1}). To be more precise, we order the threshold variable from the smallest to the largest value. The resulting order of the time index t is then applied to the variable(s) of interest (in our case z_t and $p_{i,t}$ and $p_{j,t}$) and its (their) lags.[11] Afterwards the autoregression is performed on these arranged data. The rearrangement

[10] The procedures mentioned in the following assume that actual rather than estimated time series are used. Furthermore they do not allow for linear trends or broken deterministic components. Therefore, it is problematic from a statistical point of view to refer to estimated cointegration relationships including certain deterministic terms. Nevertheless, we have also applied the tests to the log-price differences adjusted for the respective deterministic components. Our qualitative results do not change compared to the simple log-price difference. The cointegration test results support the use of $z_t = p_{1,t} - p_{2,t}$ as the cointegrating relation.

[11] In the univariate case, we consider the log-price difference z_t as the variable of interest; for the multivariate setup the single prices $p_{i,t}$ and $p_{j,t}$ are taken.

does not change the dynamic relationship between the dependent variable and its lags but transforms the threshold model into a changepoint problem. This occurs because the arranged data are implicitly ordered in accordance with their belonging to the different threshold regimes. Hence, a change in the regression parameters should be observed if we move from one regime into another as we move through the ordered sample. The Tsay procedures test for such structural changes.

The advantage of the Tsay tests is that they are independent of the threshold alternative. However, testing against a specific threshold alternative may result in higher small sample power if it is the appropriate alternative. Based on nested hypotheses, Hansen (1997, 1999) proposes testing the null of a univariate linear AR model against a general two-regime and three-regime TAR model respectively. In other words, one tests whether a two(three)-regime TAR model yields very different estimation results from a standard linear autoregressive model. If this is the case, the latter model is rejected in favour of a two(three)-regime TAR model.[12] Since the univariate procedures assume a known cointegrating relationship between the log-prices, we have also applied a multivariate SupLM test by Hansen and Seo (2001), which allows for an unknown cointegrating vector. However, as the results of their procedure do not provide additional insights, we do not comment in detail on this test.

One may directly test for threshold cointegration instead of following the two-step approach, which first applies the linear Johansen cointegration procedure and tests for threshold nonlinearity afterwards. However, we lack suitable threshold cointegration tests. Available tests only assume a two-regime threshold model (for example, Hansen and Seo 2001, Enders and Siklos 2001) and they often have rather low small sample power (compare Lo and Zivot 2001).[13] The former problem also applies to threshold unit root tests, which could be used in principle if one is willing to assume a known cointegrating relation, for the simple log-price difference z_t in our case. Finally, the treatment of deterministic terms, especially broken components, is much easier within a linear framework. On the other hand, we must note that the existence of threshold nonlinearity may distort linear cointegration tests or reduce their power.

If the tests indicate threshold nonlinearity one proceeds to estimate the threshold models. However, our price series show only weak signs of threshold nonlinearity as demonstrated later on in Section 4. Therefore,

[12] These TAR models allow for a general autoregressive structure of order k in each regime. Furthermore, no symmetry regarding the AR parameters of the outer regimes or the thresholds is imposed in the three-regime TAR model. Further details are given in the Appendix.

[13] Enders and Siklos (2001) have found that their cointegration test has lower power than a linear ADF test.

detailed results on the estimated threshold models are not presented in the empirical part of this article. Accordingly, we do not provide information about the applied estimation techniques at this time. The techniques are described in the Working Paper version of this article, which is available from the authors upon request.

3.3. Relationship between the econometric frameworks

Finally, we briefly discuss the relationship between the threshold cointegration framework and the Engle-Rogers approach in studying border effects and economic integration.

Within the threshold cointegration framework one can distinguish between two different degrees of border effects. First, a systematic border impact can prevent prices of across-border city pairs from adjusting. This strong form of border effect rules out cointegration between the prices of across-border pairs. Second, if the prices still adjust, weak border effects can lead to a slower price adjustment and higher transaction costs, that is, larger price differences are required to induce price adjustment. The former effect implies that the coefficient α in (3.2) is larger in absolute terms; the latter translates to larger threshold bands when corrected for the distance between the cities. As explained above, Engel and Rogers (1996) relate border effects to higher price variation. However, larger price deviations can be caused by the absence of price adjustment, but also by a slower adjustment speed and higher transaction costs. Thus, they cannot distinguish between high price volatility and lasting price deviations. In fact, this is the crucial difference from the threshold cointegration analysis. The estimate of the parameter α in (3.2) indicates whether adjustment towards a stable price relationship occurs, either slow or fast, or whether the prices stochastically diverge. This is not influenced by the fact that the estimate of the threshold parameter τ is also correlated with price volatility measures. Nevertheless, the Engel-Rogers approach can easily deal with large cross-sections and handle some types of structural instability as well as periods of missing data. In addition, non-significant border dummy variables are a strong result, since this excludes both strong and weak forms of border effect. In this sense, one may regard the framework of Engel and Rogers (1996) as a first step in studying the impact of borders, which should then be followed by a more detailed threshold analysis if the presence of border effects is indicated.

It should be clear from this Section that the question of price integration can only be addressed within the time series framework, which analyses the dynamic properties of the data. This cannot be done using the approach of Engel-Rogers. The threshold cointegration framework also delivers information on the speed of price adjustment and the magnitude of transaction cost bands. Nevertheless, one has to pay a price for that by assuming structural stability over time and running threshold analyses for

Table 1. *Results for regression equation (3.1) with time-specific border effects: 1921:01–1936:12.*

Variable	Coefficient	Std. error	p-value
Log distance ($d^{(i,j)}$)	0.0115	0.0059	0.0529
	City-specific dummy variables (C_m)		
Warsaw	−0.0024	0.0145	0.8689
Wilno	−0.0137	0.0201	0.4963
Poznań	0.0098	0.0171	0.5672
Łódź	0.0046	0.0145	0.7519
Kraków	0.0010	0.0164	0.9497
Lwów	0.0000	0.0183	0.9986
	Time-specific border dummy variables (B_{2124}, B_{2536})		
1921–1924 (β_2)	0.0638	0.0113	0.0000***
1925–1936 (β_3)	−0.0051	0.0039	0.1928

Note: We allow for heteroscedasticity in the error terms with respect to the different years. The adjusted R^2 and the standard error of regression are 0.306 and 0.021 respectively. Significance at the 1% level is denoted by ***. Computations have been done using EViews 4.1.

all city pairs. Hence, both econometric frameworks have advantages and disadvantages and tackle somewhat different problems. Therefore, we apply both approaches to exploit as much information from our data as possible.

4. Empirical results

4.1. Border effects

In this subsection, we present our findings on the importance of the former partition borders. Table 1 summarises the estimation results for equation (3.1) according to the Engel-Rogers approach with respect to the period 1921:01–1936:12. The estimation is based on 225 observations because we lack data for Wilno until 1923:06.

Obviously, border effects increasing the price variation are significant for the period 1921–1924.[14] By contrast, the border dummy coefficient for the years 1925–1939 is not significant at conventional levels.[15] Given the estimated border and distance coefficients, the border between 1921 and 1924 adds as much to the average price volatility as a distance of about

[14] We have also conducted a fixed-effect panel data estimation using the prices for potato and bread, for which market concentration did not play a major role. We obtained similar results: the estimated border dummy coefficient for the period 1921–24 is 0.0553 and significant at the 1 per cent level.

[15] The results do not change importantly if we sequentially delete the insignificant city dummy variables.

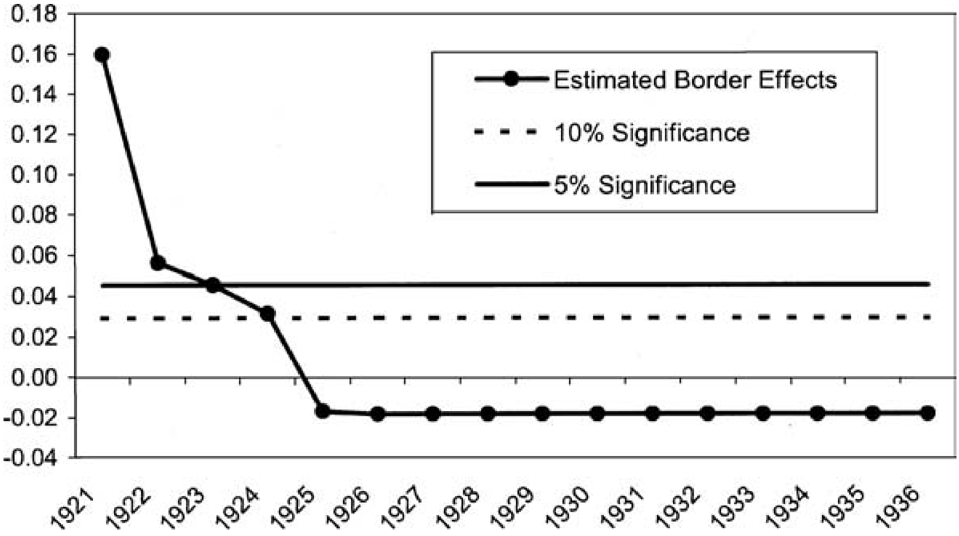

Figure 3. *Results from the smooth transition regression model.*

257 km.[16] This value is somewhat smaller than the average distance between all city pairs which is equal to 373 km. In contrast, Engel and Rogers (1996) found a much larger border effect of about 120,000 km between the US and Canada for the period June 1978 to December 1994. An explanation for these differences might be that our sample starts only in 1921. As indicated in Section 2 several steps towards formal integration were already under way at this point (such as the unification of the administrative framework) or even completed (such as the introduction of a common external tariff border in late 1919).

The results for the border effects derived from the smooth transition regression model are depicted in Figure 3. Clearly the impact of the border on price volatility decreases over time. The border matters most in the first half of the 1920s. Particularly in 1921, the border impact is very large and even exceeds the magnitude of the effect in Engel and Rogers (1996) if measured in distance. Afterwards a clear drop occurs and from 1925 onwards the impact of the former partition borders becomes insignificant.[17] The instability of the border dummy coefficient over time has also been confirmed by a formal test proposed in Lin and Teräsvirta (1994).

Our findings are in line with the regression results in Trenkler and Wolf (2004). The yearly effects in the early 1920s measured by the smooth transition approach correspond quite well to the coefficients of the yearly

[16] This has been calculated as $exp(0.0616/0.0107)$.

[17] Note in this respect that such a statement about the time point from which the effect is insignificant assumes that the transition path shown in Fig. 3 is known and not estimated. Compare the discussion in the Appendix.

border dummy variables in Trenkler and Wolf (2004). Moreover, the coefficient on the border dummy B_{2124} is slightly smaller than the average of the single yearly border effects given in that paper.

Some words of caution regarding the interpretation of our regression results are in order. First, the significant border impact in the 1920s may be part of general time effects hindering integration. In fact, our time-specific border dummy variables are similar to simple time dummy variables due to the limited number of 'within border' city pairs. Therefore, we have added such time dummy variables for the periods 1921–24 and 1925–36 into equation (3.1) in order to distinguish border and general time effects. It turns out that the border impact in 1921–24 is lower but still significant at the 5 per cent level. The general time effects are insignificant for both periods considered. Second, our results may be driven by the transition from monthly to weekly price reports in our underlying price statistics. However, the change in the magnitude and significance of the border effects already occurs in the first data collection period, which refers to monthly price averages. Hence, our findings cannot be a result of the different price collection methods.

The results on the importance of the former partition borders also have implications for the threshold cointegration analysis, according to the discussion in Section 3. Since the impact of the borders vanishes in principle after 1924 we should not expect to find evidence of border effects, either strong or weak, when examining the results of the threshold analysis covering the period 1924–37. Indeed, we could not detect any important systematic differences between the within-border and across-border city pairs with respect to the cointegration analysis and the estimated threshold models.

In conclusion, we have found significant border effects which increase price variation in the Polish wheat market during the first half of the 1920s. Hence, the integration policy in Poland was in fact successful in eliminating these negative effects. This seems to correspond to the historical description in Section 2. Whether arbitrage processes really lead to adjustment in the wheat flour prices will be studied in the next subsections within the framework of threshold cointegration analysis.

4.2. Threshold cointegration analysis

4.2.1. Preliminary analysis. The foregoing subsection has shown that only the period spanning 1921–24 displayed significant border effects. Therefore, we keep the assumption of parameter stability for our time series analysis regarding the sample January 1924–April 1937. However, we still have to consider a possible break in the deterministic terms in May 1929. In the following we present results for all city pairs instead of focusing on a certain benchmark city because some of our findings differ across pairs, although not systematically between within-border and across-border pairs.

Table 2. *Unit-root test statistics.*

	Level series	First differences
Warsaw	−2.26 (3)	−8.61 (2)***
Wilno	−2.97 (0)	−13.80 (0)***
Poznań	−3.64 (1)	−9.36 (1)***
Łódź	−3.39 (1)	−8.71 (1)***
Kraków	−2.55 (0)	−12.25 (0)***
Lwów	−3.20 (0)	−13.23 (0)***

Note: The statistics refer to Model A (level series) and C (first differences) in Perron (1989). The number of lagged differences included in the unit-root regressions is stated in parentheses, *** denotes significance at the 1% level. Critical values are −4.22 (5%) and −4.81 (1%) for the level series, and −3.74 (5%) and −4.34 (1%) for the first differences. They are taken from Tables IV.B and VI.B in Perron (1989) respectively and relate to the relative break point $\lambda = 65/160 = 0.4$. Computations have been done using EViews 4.1.

Before presenting the findings of the threshold cointegration analysis, we will study the integration properties of the price series and test for the presence of seasonality.

Table 2 summarises the results of the unit root analysis. Since the price series may exhibit a break in the trend and the level, we apply the unit root test by Perron (1989) with corrections given in Perron and Vogelsang (1993). This procedure is a generalisation of the ADF unit root test which allows for breaks in the deterministic components. For the level series we use the variant with a break in the linear trend and in the constant (Model C in Perron 1989). For the first differences, the version with a break only in the constant is applied (Model A in Perron 1989). Obviously, all log-series can be regarded as integrated of order one because the null hypothesis of a unit root is not rejected for the level series but rejected for the first differences.

The issue of seasonality is rather important since different seasonal patterns in the price series would raise doubts about integration. A test by Canova and Hansen (1995) clearly suggests modelling possible seasonality in a deterministic way and not stochastically for all cities. Therefore, we estimated, as a second step, univariate AR models for the first differences of the individual log-series including seasonal dummies in order to test for the significance of these dummy variables.[18] The dummy variables are jointly significant only for Warsaw (10 per cent level). Furthermore, the log-price differences of all 15 city pairs are not affected by deterministic seasonality at the 5 per cent level. This finding may result from the fact that flour can be produced from summer and winter wheat, both of which have the same degree of grinding. Therefore, they can be regarded as perfect substitutes.

[18] The first differences are considered to account for the non-stationarity of the log-price series.

Hence, only smaller storage capacities are required to eliminate seasonal price differences.

4.2.2. Results of cointegration analysis. The outcome of the generalised Johansen procedure is given in Table 3. The mis-specification test for vector autocorrelation described in Doornik and Hendry (1997) suggests no significant autocorrelation for the corresponding vector autocorrelation (VAR) models (compare col. 2 labelled AR (1–5)). We see that the log-prices are cointegrated at the 5 per cent significance level for all city pairs. The null hypothesis of no cointegration is rejected, whereas nonrejection regarding the null of one cointegration relation occurs. In a next step, we test whether the coefficients of the cointegrating relations with respect to the log-prices can be restricted to (1, −1). The results in the 5th column of Table 3 show that this assumption cannot be rejected at a 10 per cent level for all city pairs. Hence, the price differences of the respective city pairs are part of a stationary relationship, as required by the LOP.

As outlined in Section 3, a common stochastic dynamic of both price series is not sufficient for price integration as long as trend components enter the cointegrating relationship. In order to study which deterministic terms are present, we have applied a sequence of restriction tests that is described in the Appendix. The results are summarised in the last column of Table 3. Obviously, price convergence is rejected for the city pairs Warsaw-Lwów, Wilno-Lwów, Łódz-Lwów, and Kraków-Lwów because of a linear trend in the cointegrating relationship and for Warsaw-Łódź because of a broken linear trend. With respect to the remaining 11 city pairs, no trend has to be considered in the price relationship.

Thus, we conclude that we have evidence for the validity of a relative version of the LOP for 11 out of the 15 city pairs, since the corresponding price differences adjust appropriately to disequilibria without deterministic price divergence. However, the price differences do not eliminate the mean terms in the cointegrating relations. Hence, prices do not adjust toward price parity, but rather to some fixed difference. Interestingly, city-specific effects seem to matter for the occurrence of deterministic terms in the cointegrating relationship: all pairs including Warsaw require a modelling of broken components and four of the five Lwów pairs have a linear trend specification. Hence, price divergence is confined to city pairs including Lwów. Therefore, although integration did not evolve symmetrically across all parts of the country this asymmetry is not connected to the former partition borders.

4.2.3. Results of threshold nonlinearity analysis. Since the cointegration analysis clearly supports price integration for most of the city pairs we proceed to test for threshold effects regarding the log-price difference z_t by applying specific threshold nonlinearity tests. As described in Section 3 we apply the multivariate and univariate Tsay tests (Tsay-M, Tsay-U) and the

Table 3. *Results of generalised Johansen procedure.*

				Results of restriction tests for cointegrating relation		
City pair (k)	AR (1–5)	$H_0(r_0)$ p-value	Test statistic p-value	Price relationship is (1, −1) p-value	Deterministic terms	Half-life h in months
Warsaw-Wilno (1)	0.225	$r_0 = 0$	48.72 (0.001)***	0.066*	broken constant	2.71
		$r_0 = 1$	7.13 (0.612)			
Warsaw-Poznań (2)	0.134	$r_0 = 0$	40.72 (0.011)**	0.941	broken constant	2.10
		$r_0 = 1$	11.41 (0.233)			
Warsaw-Łódź (1)	0.111	$r_0 = 0$	50.04 (0.001)***	0.186	broken linear trend	3.41
		$r_0 = 1$	11.04 (0.260)			
Warsaw-Kraków (1)	0.501	$r_0 = 0$	49.27 (0.001)***	0.614	broken constant	2.53
		$r_0 = 1$	10.29 (0.311)			
Warsaw-Lwów (1)	0.562	$r_0 = 0$	49.61 (0.001)***	0.962	linear trend	5.24
		$r_0 = 1$	9.84 (0.347)			
Wilno-Poznań (1)	0.106	$r_0 = 0$	39.35 (0.013)**	0.622	constant	2.08
		$r_0 = 1$	7.55 (0.567)			
Wilno-Łódź (1)	0.406	$r_0 = 0$	51.02 (0.000)***	0.509	constant	3.16
		$r_0 = 1$	9.79 (0.351)			
Wilno-Kraków (1)	0.717	$r_0 = 0$	65.43 (0.000)***	0.167	constant	2.11
		$r_0 = 1$	10.59 (0.288)			
Wilno-Lwów (1)	0.319	$r_0 = 0$	63.72 (0.000)***	0.111	linear trend	1.80
		$r_0 = 1$	11.28 (0.241)			
Poznań-Łódź (2)	0.264	$r_0 = 0$	46.55 (0.002)***	0.643	constant	1.99
		$r_0 = 1$	15.27 (0.074)*			
Poznań-Kraków (2)	0.581	$r_0 = 0$	45.60 (0.003)***	0.288	constant	1.48
		$r_0 = 1$	13.43 (0.131)			
Poznań-Lwów (2)	0.537	$r_0 = 0$	39.93 (0.013)**	0.710	constant	2.47
		$r_0 = 1$	14.90 (0.083)*			

Łódź-Kraków (1)	0.365	$r_0 = 0$	51.72 (0.000)***	0.477	constant	1.58
		$r_0 = 1$	9.06 (0.416)			
Łódź-Lwów (1)	0.348	$r_0 = 0$	48.16 (0.001)***	0.313	linear trend	4.06
		$r_0 = 1$	9.55 (0.372)			
Kraków-Lwów(1)	0.841	$r_0 = 0$	45.10 (0.003)***	0.567	linear trend	3.20
		$r_0 = 1$	10.18 (0.320)			

Note: The number of lagged differences of the respective VAR is stated in parentheses behind the city pair. AR (1–5) represents the p-value for a misspecification test against vector autocorrelation for lags from one to five (see Doornik and Hendry 1997). This test is F (20,290–8k) distributed. The cointegration test statistics are trace tests where the respective null hypothesis is tested against a higher number of cointegrating relations. The p-values are computed from the response surface given in Johansen *et al.* (2000) for a relative break point $\lambda = 65/160 = 0.4$. The restriction test regarding the price relationship is $\chi^2(1)$ distributed. The half-lives are based on the estimates of a linear univariate AR(1) model for the log-price differences. These price differences have been adjusted for the deterministic terms which enter the respective cointegrating relation according to the restriction test outcomes. ***, **, * denote significance at the 1%, 5%, and 10% level respectively. The cointegration analysis has been done with PcFiml 9.10 (see Doornik and Hendry 1997).

Table 4. *Results of threshold nonlinearity tests.*

City pair	Tsay-M $\chi^2(4k)$ p-value (k)	Tsay-U $F(2, 141)$ p-value (k)	Hansen-U12 Bootstrap p-value (k, d)	Hansen-U13 Bootstrap p-value (k, d)
Warsaw-Wilno	0.006 (1)***	0.718 (1)	0.112 (1,1)	0.224 (1,1)
Warsaw-Poznań	0.008 (2)***	0.335 (2)	0.706 (2,1)	0.882 (2,1)
Warsaw-Łódź	0.254 (1)	0.392 (2)	0.342 (2,1)	0.491 (2,1)
Warsaw-Kraków	0.758 (1)	0.648 (2)	0.425 (2,2)	0.248 (2,2)
Warsaw-Lwów	0.157 (1)	0.030 (2)**	0.037 (2,2)*	0.040 (2,2)**
Wilno-Poznań	0.000 (1)***	0.000 (1)***	0.075 (1,1)*	0.157 (1,1)
Wilno-Łódź	0.015 (1)**	0.000 (1)***	0.329 (1,1)	0.579 (1,1)
Wilno-Kraków	0.007 (1)***	0.004 (1)***	0.026 (1,1)**	0.035 (1,1)**
Wilno-Lwów	0.307 (1)	0.002 (2)***	0.081 (2,1)*	0.431 (2,1)
Poznań-Łódź	0.631 (2)	0.451 (1)	0.067 (1,1)*	0.243 (1,1)
Poznań-Kraków	0.787 (2)	0.736 (1)	0.654 (1,1)	0.323 (1,1)
Poznań-Lwów	0.262 (2)	0.247 (1)	0.118 (1,1)	0.327 (1,1)
Łódź-Kraków	0.122 (1)	0.295 (1)	0.774 (1,1)	0.275 (1,1)
Łódź-Lwów	0.269 (1)	0.185 (1)	0.586 (1,1)	0.852 (1,1)
Kraków-Lwów	0.351 (1)	0.674 (2)	0.790 (2,1)	0.543 (2,1)

Note: Tsay-M and Tsay-U abbreviate the multivariate and univariate tests of Tsay (1989, 1998). Hansen-U12 and Hansen-U13 are short for the procedures of Hansen (1997, 1999) testing against a two-regime and three-regime TAR model respectively. The number of lags k used in the respective autoregressions and the delay order d of the threshold variable z_t are stated in parentheses behind the p-value. The delay order d for the Tsay tests is always 1, i.e. z_{t-1} is used as threshold variable. ***, **, * denote significance at the 1%, 5%, and 10% level respectively. The Tsay nonlinearity test statistics and their p-values are computed using own GAUSS programs. GAUSS programs from Bruce Hansen's web page (<http://www.ssc.wisc.edu/~bhansen/progs/progs_threshold.htm>) are applied to compute the test statistics for the procedures by Hansen (1997, 1999) and their respective bootstrap p-values.

univariate procedures by Hansen (1997; 1999), which test linearity against a two- and three-regime TAR model respectively (Hansen-U12, Hansen-U13).

The results in Table 4 indicate that threshold nonlinearity is not relevant for the price dynamics in general. Only with respect to some city pairs, including Wilno, there does seem to be robust evidence for threshold effects. However, one has to be careful in interpreting the outcome of these tests. None of the procedures allows for broken deterministic components or a linear trend (compare fn. 10). This may be problematic for the pairs containing Warsaw or Lwów. The cointegration analysis has shown that broken deterministic terms or a trend have to be included in the price relationships for the corresponding city pairs. Furthermore, the univariate nonlinearity tests assume stationarity under the null hypothesis. If the price differences are near unit root processes, the nonlinearity tests may be size-distorted (cf. Lo and Zivot 2001).

Nevertheless, our results are confirmed by similar findings from the Tsay and Hansen procedures when applied to the price relationships adjusted for the deterministic terms, and by a test derived from Hansen and Seo (2001), which takes account of a linear trend. Moreover, the results of the estimated BAND-TAR model (3.2) indicate that threshold bands are only important for the city pairs including Wilno.[19] This refers to both the band widths and the number of observations included in the bands. These findings are in line with the outcomes of the nonlinearity tests, which robustly suggest threshold effects only for the Wilno pairings. The fact that threshold nonlinearity is of importance with respect to these pairs should be no surprise since Wilno is the most remote city. Thus, threshold regimes are supposed to be wider owing to higher transportation costs. For the other pairs, the threshold bands are often quite small and contain a rather low number of observations. Moreover, estimates of unrestricted threshold models seem to be very unreliable. Since threshold dynamics are not very important for most of the pairs, we do not present more detailed results of the estimated threshold models. They can be found in the Working Paper version of this article, which is available from the authors upon request.

Our price convergence results are in line with similar studies by Persson (1999) and Ejrnæes and Persson (2000) on European grain and French wheat markets in the 19th century. This also refers to the estimated half-lives in Table 3, which are comparable to the ones computed by Persson (1999) for the French wheat market. He found a half-life of price parity deviations between Toulouse and Bordeaux of 2.2 months (1855–72), for Toulouse and Rouen of 5.1 (1860–80), and for Rouen-Marseille of 6.1 months (1885–1913). However, in contrast to our findings threshold nonlinearity is very important for the above-mentioned empirical studies. Therefore, we summarise as follows. The weak evidence for threshold nonlinearity and the rather small threshold bands point to relatively low transaction costs. We interpret this as a further sign of functioning price adjustment and economic integration in line with the main results of the cointegration analysis.

5. Summary and concluding remarks

In this article we study the topic of across-border integration for the Polish interwar economy by analysing the wheat flour market between 1921 and 1937 with respect to six cities. To be precise, we ask whether borders hinder economic integration whereby we consider an economy to be integrated if the LOP holds taking transaction costs into account.

Let us briefly summarise our findings. First, using the approach of Engel and Rogers (1996) we find evidence that the former partition borders matter until 1924, but their effect vanishes after that date. In fact, we have evidence that a relative version of the LOP holds for 11 out of our 15 city pairs during

[19] The model (3.2) is estimated for the price-differences z_t^*, which is defined to include the relevant deterministic terms entering the cointegration relations (see p. 211).

the period 1924–1937. To be precise, the prices of wheat flour adjust to disequilibria, but a (broken) constant enters the price relationship. Hence, some price differences remain between the market places so that integration is not perfect. It seems that city-specific effects matter in this respect, though there is no pattern that suggests a persistent effect of the former partition borders. Moreover, our findings on threshold bands of no-arbitrage also point to city-specific factors rather than border effects. For example, the multivariate Tsay Test evinces significant threshold-nonlinearity for city-pairs including Wilno, the most remote city in our sample, but this applied to both within-border pairings (Wilno-Warsaw, Wilno-Łódź) and across-border pairings (Wilno-Poznań, Wilno-Kraków). Hence, nonlinearities and the implied transaction costs seem to exist but they are generally small and not related to the former borders. Thus we regard the Polish interwar economy as integrated with some restrictions and interpret this as a success of the Polish integration policy after the reunification in 1919.

Our results for interwar Poland markedly differ from the findings of Moodley *et al.* (2000) and Engel and Rogers (1996), who analyse economic integration between the US and Canada in the wake of the Free Trade Arrangement (FTA) of 1990. Interestingly, neither study shows important effects of the FTA on the degree of integration, which contrasts with our results for border changes in Poland. Most probably, this is due to the fact that integration in the Polish case involved not only the removal of tariff barriers, but also improved regional mobility and communication in virtually all aspects.

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Appendix

In this Appendix we describe some of the more complicated technical details regarding the econometrics methods applied and provide some more detailed information on specification issues and empirical results.

Smooth transition regression model

Based on Lin and Teräsvirta (1994) we have estimated the following smooth transition regression model

$$\begin{aligned} &V(z_t^{i,j}) = \beta_1^* d^{i,j} + \beta_2^* B^{i,j} + \beta_3^* B^{i,j} F(t^*) + \sum_{m=1}^{6} \beta_{4m}^* C_m + u_t^{i,j} \quad \forall\, i, j \ with\ i \neq j, \\ &\text{with } F(t^*) = (1 + exp\{-\gamma(t^{*3} + \alpha_1 t^{*2} + \alpha_2 t^* + \alpha_3)\})^{-1} \end{aligned} \tag{5.3}$$

where $t^* = t/T$ and $B^{(i,j)}$ is a border dummy set to 1 if the cities i and j belong to different former partition areas and zero otherwise. The other variables are defined as in (3.1). The transition function $F(t^*)$ allows the border effect to change smoothly from β_2 to $\beta_2 + \beta_3$. The third-order polynomial representation with respect to t^* given here is the most flexible form for the transition process considered in Lin and Teräsvirta (1994).

The testing and estimation procedure works as follows. First, as suggested by Lin and Teräsvirta (1994), one tests whether the parameters of the model under consideration are constant over time. This test boils down to a standard WALD restriction test on the relevant parameters, using a first-order Taylor approximation of $F(t^*)$. The null hypothesis of constant border effects is clearly rejected in our case. Using the same restriction test approach, it is then determined which order of t^* should be used for $F(t^*)$. The corresponding test strongly suggests applying the third-order polynomial representation, as done above. Finally, the model (5.3) is estimated by nonlinear least squares methods. We used the program EViews 4.1. The starting values for the estimation procedure for $\beta_i^*(i = 1, 2, 3)$ and $\beta_{4m}^*(m = 1, \ldots, 6)$ are taken from the regression results for (3.1). The starting value for β_3^* is the sum of the estimators of $-\beta_2$ and β_3. The parameters γ and $\alpha_i(i = 1, \ldots, 3)$ were initially set to 75, –0.5, 0.5, and –0.1. These values approximately generate yearly changes in the border effects as found in Trenkler and Wolf (2004), who used yearly border

Appendix Table. *Results for smooth transition regression (5.3).*

Variable	Coefficient	Std. error	p-value
Log distance ($d^{(i,j)}$)	0.0061	0.0088	0.4876
City specific dummy variables (C_m)			
Warsaw	0.0086	0.0215	0.6907
Wilno	0.0040	0.0303	0.8948
Poznań	0.0455	0.0265	0.0874*
Łódź	0.0207	0.0217	0.3395
Kraków	0.0212	0.0256	0.4094
Lwów	0.0273	0.0273	0.3185
Border dummy variable $B^{(i,j)}$			
β_2	0.1681	0.0651	0.0104**
β_3	−0.1864	0.0648	0.0044***
Transition function			
γ	2266.5	2822.2	0.4228
α_1	−0.5558	0.0642	0.0000***
α_2	0.1012	0.0227	0.0000***
α_3	−0.0057	0.0022	0.0085***

Note: We allow for heteroscedasticity in the error terms with respect to the different years. The adjusted R^2 and the standard error of regression are 0.596 and 0.035 respectively. Significance at the 1%, 5%, and 10% level is denoted by ***, **, * respectively. Computations have been done using EViews 4.1.

dummy variables in a regression like (3.1). The estimation results are collected in the Appendix Table above.

As pointed out by Lin and Teräsvirta (1994), the large standard deviation of the estimate of γ does not indicate insignificance. It arises from the fact that there is a wide range of values of γ yielding almost the same transition function. In addition, γ is not identified if no transition takes place (parameter constancy). Thus, the standard interpretation of the t-ratio as a test of the hypothesis $\gamma = 0$ is invalidated (cf. Teräsvirta 2004).

The dots in Figure 3 have been computed by $\hat{\beta}_2 + \widehat{F(t^*)}\hat{\beta}_3$ where '$\widehat{(\cdot)}$' represents the use of the corresponding estimators. Accordingly, the line connecting the dots describes the estimated gradual change in the border effects over time. The significance bands in Figure 3 are determined as follows. We search for the value s which satisfies $\hat{\beta}_2 + s\hat{\beta}_3 = z_{\alpha/2}$ where $z_{\alpha/2}$ is the right-hand side α%-critical value of the normal distribution (or in other words the $(1 - \alpha/2)$-quantile of the normal distribution). Loosely speaking, we determine the earliest point on the transition path from $\hat{\beta}_2$ to $\hat{\beta}_2 + \hat{\beta}_3$ for which the border effect is not significantly different from zero at the level α. If we want to relate that point to a time point in our sample, we face the problem that the transition path is estimated. Thus, by comparing the transition path with the significance lines we implicitly assume $F(t^*)$ to be known. It would be more appropriate to consider the uncertainty involved in the estimation of $F(t^*)$ by considering a confidence interval around $\widehat{F(t^*)}$. However, this seems to be a rather difficult task since the estimation of $F(t^*)$ involves the parameter γ for which standard inference is not valid as mentioned above.

Threshold models

Our exposition of the threshold cointegration framework has been based on the BAND-TAR model (3.2). Of course, it is possible to consider more general TAR models regarding the log-price difference z_t. For example, one may introduce richer autoregressive dynamics or give up symmetry regarding the thresholds and AR parameters in the outer regimes. Later on, when explaining the nonlinearity tests by Hansen (1997, 1999) we will briefly describe the so-called SETAR model, which is one possible generalisation of (3.2).

In principle, it is recommended to test for the restrictions on the model parameters implied by the transaction-cost view within a more general TAR model. Unfortunately, the simulation results of Lo and Zivot (2001) demonstrate that possible Wald and LR restriction tests are heavily size-distorted in small samples, even for simple processes and rather large sample sizes. Therefore, Lo and Zivot (2001) conclude that these procedures are essentially useless. Furthermore, one may refer to multivariate threshold models with respect to the single time series $p_{i,t}$ and $p_{j,t}$. However, the simulation evidence given in Lo and Zivot (2001) does not indicate a general advantage of a multivariate model setup over a univariate one. Because of these problems we have focused on the BAND-TAR model (3.2) in our presentation.

Johansen cointegration testing framework

In our empirical study we have applied a generalisation of the multivariate Johansen testing procedure derived from Johansen *et al.* (2000). It not only allows for broken linear trends and levels, but also enables us to test which deterministic components affect the price cointegration relationship. The latter has important implications for the validity of the LOP as described in the main body of the text. Additionally, we can test whether the cointegrating vector can be restricted to (1, −1) so that the log-price difference is in fact the relevant quantity for price adjustment.

Assuming a bivariate price-system, one break in the deterministic terms at time $t = T_1$, and a lag order of $k = 1$, the Johansen procedure is based on a maximum likelihood estimation of the linear n-dimensional vector error correction (VEC) model:

$$\begin{aligned}\Delta p_t &= \alpha(\beta' p_{t-1} - \theta_1(t-1)D_{1,t} - \theta_2(t-1)D_{2,t}) + \nu_1 D_{1,t} + \nu_2 D_{2,t} + \gamma_2 d_{2,t} + \epsilon_t,\\ &\quad \epsilon_t \sim N(0,\Omega) \text{ and } t = p+1,\ p+2, \ldots, T,\end{aligned} \tag{5.4}$$

where $p_t = (p_{i,t}, p_{j,t})'$. $D_{1,t}$ equals one for all observations before T_1 and zero otherwise, $D_{2,t} = 1 - D_{1,t}$ and $d_{2,t}$ equals one for $t = T_1$ and zero otherwise. Hence, these variables describe the two regimes before and after the break in the deterministic components. The parameters $\theta_1, \theta_2, \nu_1$ and ν_2, are the corresponding $(n \times 1)$ parameter vectors for the linear trends and constants of the two regimes. In the empirical analysis we have set T_1 to May 1929. The Johansen procedure tests for the rank r of the matrix $\Pi = \alpha\beta'$, where $\alpha(n \times r)$ is the matrix of adjustment coefficients and the matrix $\beta(n \times r)$ contains the coefficients of the cointegrating vectors related to the prices $p_{i,t}$ and $p_{j,t}$. Hence, the rank r determines the number of cointegration relations. We consider the trace test version, that is, the pair of hypotheses is $H_0(r_0) : \text{rk}(\Pi) = r_0$ *vs*. $H_1(r_0) : \text{rk}(\Pi) > r_0$. We expect a cointegrating rank of one since the LOP implies a cointegrating relationship between the log-prices. Critical or p-values of the test can be computed by using a response surface given in Johansen *et al.* (2000). Of course, one has to augment (5.4) by lags of Δp_t and $d_{2,t}$ if necessary.

Provided we have found a cointegrating rank of one, we have tested whether β can be restricted to $(1, -1)$. This restriction has not been rejected for all city pairs at the 10 per cent significance level. Then, we have performed restriction tests on the parameters $\theta_1, \theta_2, \nu_1$, and ν_2 in order to examine which deterministic terms are present and enter the cointegrating relationship. We applied the following sequence of tests.

1. Test of null hypothesis $H_0^1 : \theta_1 = \theta_2 = 0$ (no trends in cointegrating relation).
2. If H_0^1 is rejected:
   Test of null hypothesis $H_0^2 : \theta_1 = \theta_2$ (common trend in cointegrating relation).
   (a) If H_0^2 is rejected: Rejection of price convergence due to broken deterministic trend.
   (b) If H_0^2 is not rejected: Rejection of price convergence due to a common deterministic trend.
3. If H_0^1 is not rejected:
   Test of null hypothesis $H_0^3 : \nu_1 = \Pi\mu_1$ and $\nu_2 = \Pi\mu_2$
   (no trends in the VEC model)
   H_0^3 has not been rejected regarding all relevant city pairs
   (a) Test of null hypothesis $H_0^{3a} : \mu_1 = \mu_2$
   (common mean term in cointegrating relation)
   (aa) If H_0^{3a} is rejected: Non-rejection of price convergence, but broken mean term in cointegrating relation
   (ab) If H_0^{3a} is not rejected: Non-rejection of price convergence, but common mean term in cointegrating relation

Note that all tests are conducted conditional upon the outcome of the previous tests in the sequence. All restriction tests are asymptotically $\chi^2(m)$ distributed where m refers to the number of restrictions tested. More details on these tests and the Johansen procedure can be found in Johansen *et al.* (2000) and Johansen (1995). The final results of the testing sequence on the deterministic terms are summarised in Table 3.

Threshold nonlinearity tests

The main idea in Tsay (1989, 1998) of using arranged autoregressions in order to test for threshold nonlinearity has been already explained in Section 3. If the data are generated by a threshold model, the arrangement of the data according to the value of the threshold variable groups the data into the different threshold regimes. Hence, performing a linear autoregression of order k on the arranged data should create changes in the parameters at the threshold points. The univariate and multiple Tsay procedures test for structural changes in the parameter of the autoregressions with the help of a recursive least squares estimation scheme. If the evidence for parameter changes is strong enough, the null hypothesis, of linearity is rejected. The test statistics are F (univariate test) and χ^2 (multivariate test) distributed respectively. For further details see Tsay (1989, 1998).

The autoregressive orders for the regressions of the multivariate test have been chosen in accordance with the respective orders of the VECMs on which the Johansen tests rest. In the case of the univariate test we took the orders which are appropriate if a ADF unit root test regression is performed on z_t. We only obtained orders of $k = 1$ and $k = 2$. In line with model (3.2) we have chosen z_{t-1} as the threshold variable. Regarding the city pairs for which an order $k = 2$ has been suggested, we also considered the threshold variable z_{t-2}. The change in the delay order from one to two has not altered the test results in any important way.

The procedures by Hansen (1997, 1999) are based on comparing the estimation results for a linear AR model, which represents the null hypothesis, with results for a two-regime or three-regime TAR model respectively. The three-regime TAR (TAR (3)) model considered by Hansen (1997, 1999) can be written as

$$z_t = \begin{cases} c_1 + \alpha_{11} z_{t-1} + \ldots + \alpha_{1k} z_{t-k} + \eta_t, & \text{if } z_{t-d} > \tau_1 \\ c_2 + \alpha_{21} z_{t-1} + \ldots + \alpha_{2k} z_{t-k} + \eta_t, & \text{if } -\tau_3 \leq z_{t-d} \leq \tau_1 \\ c_3 + \alpha_{31} z_{t-1} + \ldots + \alpha_{3k} z_{t-k} + \eta_t, & \text{if } z_{t-d} < \tau_3, \end{cases} \tag{5.5}$$

where τ_1 and τ_3 are the (non-symmetric) thresholds and d is the delay order regarding the threshold variable with $0 < d \leq \bar{d}$. The upper bound for d is typically set to k and we apply that bound as well. Since a lag of z_t is chosen as the threshold variable the model (5.5) is called a self-exciting threshold autoregressive (SETAR) model.

The TAR(3) model is then estimated by sequential least squares methods. The sum of squared residuals of this model (S_{TAR}) is compared to the sum of squared residuals obtained from the linear autoregression of the same order k (S_{LIN}). To be precise, the test has the so-called sup-F-type (sup-Wald) form $F_{13} = T((S_{LIN} - S_{TAR})/S_{TAR})$ where T is the number of observations. Hence, if S_{LIN} is much larger than S_{TAR}, the linear AR models provide a much worse fit than the TAR(3) model and we reject the null hypothesis of linearity. Since the threshold parameters are not identified under the null hypothesis the asymptotic distribution of F_{13} is nonstandard. Therefore, p-values have to be determined by bootstrap methods. The test setup can be adjusted accordingly for a two-regime TAR model. Further details are given in Hansen (1997, 1999).

The orders for the autoregressions are chosen as for the univariate Tsay test. The delay order d is automatically determined within the test procedures as the optimal order out of the range $0 < d \leq k$.