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by

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Abstract

In this paper we study the issue of economic integration in Poland between 1924 and 1937 by means of a threshold cointegration analysis of the law of one price. We find that the interwar economy can be regarded as integrated but with obvious restrictions which refer to the existence of relevant transaction costs for arbitrage and differences in the level of prices between cities. Moreover, the former partition borders within Poland did not affect economic integration suggesting that the integration policy after the reunification of Poland in 1919 was successful. However, we are not able to assess the impact of the aggregate price level which changed strongly during our sample period.

Keywords: Economic integration, Law of one price, Transaction costs, Poland, Threshold cointegration, Threshold nonlinearity tests

JEL classification: C22, C32, F15, E31, N74, N94

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1 Introduction

After the First World War Poland was reunified in 1919 which ended up the long period of partition beginning at the end of the 18th century. The new state consisted of three different parts which belonged to the former tsarist Russia, the Habsburg monarchy and the German Empire. In 1919 these areas were not only dramatically different with respect to their institutional framework (administration, law, currency), and their social and economic structures, but they were separated by high costs of transportation and communication. Accordingly, all Polish governments after 1919 attempted to unify and integrate the country.

We want to evaluate the integration process with a focus on the question whether the former partition areas have economically integrated into one single market. By integration we mean that the law of one price (LOP) holds. In its strict form the LOP says 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 equalize the prices. Of course, one has to consider the transaction costs associated with arbitrage. Only if the price deviations between the two market places exceed the transaction costs, arbitrage is profitable and takes place.

In our empirical analysis covering the period from 1924 to 1937 we focus on the wheat market by using monthly retail prices for wheat flour. This market has been chosen since wheat flour is a rather homogenous good so that the LOP is in general applicable. Furthermore, we have no evidence for monopolistic structures or existing cartels in this market implying that the prices can be regarded as an outcome of a competitive market. To be specific, we have data for six of the biggest Polish cities: Warsaw, Lodz, and Wilno, belonging to the former Russian area, Lwów and Kraków belonging to the former Austrian part and Poznań from the former German part. This gives us the possibility to analyze 15 city pairs.

The question whether Poland was successful in creating a unified market is an helpful information when studying Polands history and for evaluating the political institutions in charge in the 1920s and 1930s. Additionally, after the dramatic political changes at the end of the 1980s many political officials referred to this interwar-period when justifying and setting up new institutional structures. It would be of great interest to know whether this reference is solely a psychological one or whether it can be based on verifiable facts at least with respect to the specific economic issue we look at. Moreover, the general question whether areas or countries with different institutional backgrounds can be successfully integrated is an important one up to now and is, for example, crucial for the process of extending the European Union.

Obviously, it is not only of interest whether market integration can be observed but also what forces have an effect on the process of integration. The first issue in this respect relates to the already mentioned question of market integration of the former partition areas. In other words, we ask how important the old borders were, i.e. whether there existed effects of the former partition borders within Poland. Usually, the literature examines border effects between different countries. However, it is also very interesting to analyze how long it takes to make the impact of a political or administrative border disappear between areas that enter a political union.

Engel & Rogers (1996) mention some reasons for border effects that may also be relevant for Poland during the interwar period. Among other things, different degrees of local labor market integration, direct costs of crossing borders like tariffs and other regulations, and the existence of different currencies and productivity shocks are such reasons. The description of the historical background later on will show that these factors were present in Poland, especially in the first years after the unification. Since our sample starts in 1924, five years after unification, we are able to asses whether the institutional changes implemented in these five years have created a framework so that there exist no systematic differences between city pairs with cities from the same area ("within-border pairs") and pairs of cities from different partition areas ("across-border pairs").

A second factor probably relevant for the process of integration is the influence of a changing aggregate price level in the presence of nominal fixed transaction cost. Until the end of the 1920s we can observe an important increase in the aggregate price level in Poland. This inflationary period was followed by strong deflation. Only at the middle of the 1930s the aggregate price level stabilized. At the same time, most of long-distance freight transport in Poland has been done by trains, but the nominal tariffs for using the railway network remained remarkably stable in the 1920s and 1930s.[1] Thus, the real transportation costs changed mainly with the price level. During the inflationary period the real transportation costs increases and in times of deflation the costs increase. Insofar these nominal fixed

[1]Tariffs for merchandize were changed three times during our sample period: they were somewhat increased in December 1926, and in October 1929, and decreased in March 1936 (Gieysztor 1939).

transport costs are a crucial part of the whole transaction costs a changing aggregate price level should drive the integration process: inflation fosters and deflation slows economic integration owing to falling and increasing real transaction costs respectively.

The importance of the old borders and the changing aggregate price level are studied along with the general question of a successful market integration by performing a threshold cointegration analysis. Threshold cointegration models can be regarded as the econometric model equivalent to the LOP taking transaction costs into account. The transaction costs view imposes certain parameter restrictions on these nonlinear models. In line with Balke & Fomby (1997) and Lo & Zivot (2001) this analysis is conducted in three steps testing for cointegration, threshold nonlinearity, and estimating the threshold models. These three steps aim to study whether the markets are integrated, the transaction costs approach is appropriate and whether the model parameters satisfy the economic theory respectively.

Our time series approach has a number of advantages over a cross-section analysis like in Engel & Rogers (1996). They compute volatility measures of price pairs over time for several U.S. and Canadian cities and regress them on distance measures, a dummy variable describing border crossing and other variables. Thereby, they can evaluate whether distance and border explain price volatility what would violate the strict LOP. In contrast, the time series approach allows us to capture the dynamics of the data explicitly. We are able to analyze whether the prices still adjust, i.e. whether they cointegrate, although they vary over time. In this sense lasting price deviations have to be distinguished from simple price volatility. Additionally, by estimating adjustment coefficients and threshold bands we can describe the speed of adjustment and the importance of transaction costs. These estimates can still be related to measures describing distance and border effects. Hence, we are able to derive results which cannot be obtained from a cross-section analysis like in Engel & Rogers (1996) who do not exploit the integration and cointegration properties of the time series explicitly when computing price volatility measures.

The rest of the paper is organized as follows. In the next section we describe the historical background. Section 3 introduces the economic model framework and we comment on some characteristics of the Polish wheat market. The econometric framework and methods we apply are described in Section 4 and the empirical results are presented in Section 5. The last section summarizes and concludes.

2 Historical Background

At the end of 1918 Poland might be described as a power vacuum in central Europe, with several political and military authorities struggling for influence on a territory without clear shaped borders. The devastations of the First World War affected 90% of this area, destroyed the harvest and the livestock, buildings and machines, bridges and railways. Even more damage was done by the exploitation from the German and Russian occupants during the war and sabotage during their retreat (Duda & Orlowski 1999, p. 231). However, this chaos got structured with an amazing speed. In official statistics, the state was from 1921 on organized in 17 administrative units (vojvodships), which can often found to be aggregated into several groups that followed the old lines of partition between the former occupants: the western, southern and central vojvodships, covering approximately the former partition areas of Germany, Austria and Russia, and the eastern vojvodships, covering former Russian areas in the east that were claimed by Polish nationalists.[2] When the political situation in November 1918 gradually stabilized with the return of Jozef Pilsudski to Warsaw, all parties saw the necessity to create a unified institutional framework with adequate infrastructure in order to establish that new state. Actually, the government could rely on extensive programs for a legal, administrative and economic unification, that had been prepared since 1907 for a future Polish state (Roszkowski 1992).

The unification of the fiscal administration - a conditio sine qua non for the survival of the new state - belonged to the very first institutional changes. While for the southern and central vojvodships this was formally reached already in April 1919, the former German parts remained separated until January 1922, Upper Silesia even until June 1922 (Markowski 1927, Bielak 1931). Several differences of the tax system however - e.g. the real estate tax - remained persistent until 1936 (for details see Weinfeld 1935). An even more demanding task was the creation of a common currency area, unifying the five (!) currencies, that were in circulation on the 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

[2] Western vojvodships comprised: Poznań, Pomerania, Silesia; central: city of Warsaw, Warsaw, Lódz, Kielce, Lublin, Białystok; southern: Kraków, Lwów, Stanisławów, Tarnopol; eastern: Wilno, Nowogrod, Polesia, Wolhynia; see Głowny Urząd Statystyczny (1939).

Table 1. Important railway-connections between main cities and average length of the trip.

Date of opening	Connection	Distance	Length of the trip (measured for 1937)
1848	Warsaw-Kraków via Czestochowa	ca.364 km	8.00 hrs
25/ 11/ 1934	Warsaw-Kraków via Radom	ca.320 km	5.20 hrs
1872	Warsaw-Poznań via Torun	ca.376 km	7.00 hrs
1/ 11/ 1921	Warsaw-Poznań via Wrzesnia	ca.304 km	4.45 hrs
1857	Poznań-Kraków via Wroclaw	ca.380 km	n.a.
1/11/ 1926	Poznań-Kraków via Wielun	ca.330 km	n.a.
1861	Kraków-Lwów	ca.341 km	5.00 hrs
1917	Warsaw-Lwów via Lublin	ca.500 km	8.30 hrs

Sources: Pisarski (1974, p. 58); Obraz Polski dzisiejszej (1938, p. 223).

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 delegalized. For the central, southern and western vojvodships this was realized already in April 1920, with the exception of Upper Silesia (Nov. 1923) (Zbijewski 1931).

During the first years of reunified Poland however these activities towards market integration were hindered by a system of regulations concerning most commodity goods, as well as factor markets. For the most part this was motivated by the need to furnish supply for the Polish troops, fighting with the Soviet army in the east, but it had also aspects of political logrolling between different groups. Especially the markets for agricultural products (e.g. bread, grain, potato, sugar) and basic commodities (e.g. coal, soap, matches) were affected by a variety of measures that discriminated between regions and social groups. While for instance a common external tariff was introduced in November 1919 and domestic tariff barriers between the different parts of Poland had been removed for the most part already during the First World War, there remained a customs frontier between the former Prussian partition area and the rest (see Kozlowski 1989, p.157 and Landau & Tomaszewski 1999, p.69). After the armistice between Poland and Soviet Russia the polish government launched a program to liquidate the whole system of regulations. The internal customs frontier were removed in mid-1921, and until the end of 1921 most other regulations on the commodity markets had disappeared (Tomaszewski 1966, Kozlowski 1989, p.158.)

The transportation system of the new state in turn apparently profited from the war in the east. After rather spontaneous takeovers of the railway networks in the different areas during the last months of the First World War, already in October 1918 a railway ministry started its work and developed a 10-years 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 (e.g. refrigerator wagons), the different densities of the network adjusted and the main economic centers 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 depended also crucially on the material used. Table 1 gives an overview for the development of important new built railway lines and the changes in speed.

Since nearly all freight transport took place on railways (97,6% in 1925 and 98,7% in 1938)[3] this development in the railway network together with the extensive unification of the institutional framework until the end of 1922 can be expected to have had a strong impact on commodity market integration.

However, as indicated in the introduction it is possible that the impact of these factors on the course of economic integration was completely dominated by another effect, namely by changes in the aggregate price level. Figures 1 and 2 give the food price (FPI) as published from 1921 on by Głowny Urząd Statystyczny (GUS), the Polish Statistical Office. Calculated on the base of price data for 16 food products from 175 Polish cities it is the most complete price index available for interwar Poland. The FPI indicates massive inflation until 1923, followed by a short period of stabilization and a second inflation until the end of 1926. After just two years of stable prices, the index fell by more than 50 points, and stabilized not before mid 1935. What drove the aggregate price level? The discussion will show that the mentioned steps towards an institutional unification of the Polish economy - including the creation of a common currency and an efficient fiscal administration - can be seen as functional of these factors.

[3]See Brzosko (1982, p. 358). This information obviously refers only to that part of transportation, that was comprised in some kind of official statistics, i.e. transport over longer distances. There exists an estimation, that during the period 1934-1936 just 39% of the total Polish wheat surplus (production not consumed by producers themselves) was transported on railways. The rest was mainly shipped on horse-drawn vehicles within a 50km radius which we can drop for the purpose of our investigation (Buczyński 1939, pp. 91ff).

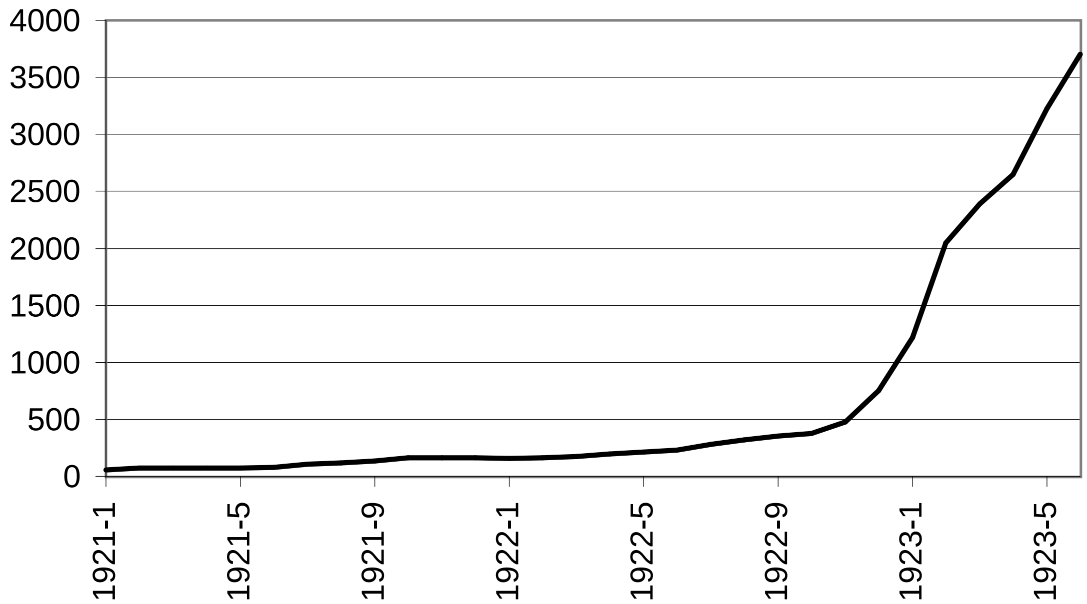

Figure 1. Food price index 1921-1923 (1921=100)

Sources: Głowny Urząd Statystyczny [Polish Statistical Office], Rocznik Statystyczny [Statistical Yearbook] RP 1920/1922 (II), p. 215 and RP 1923, p. 88.

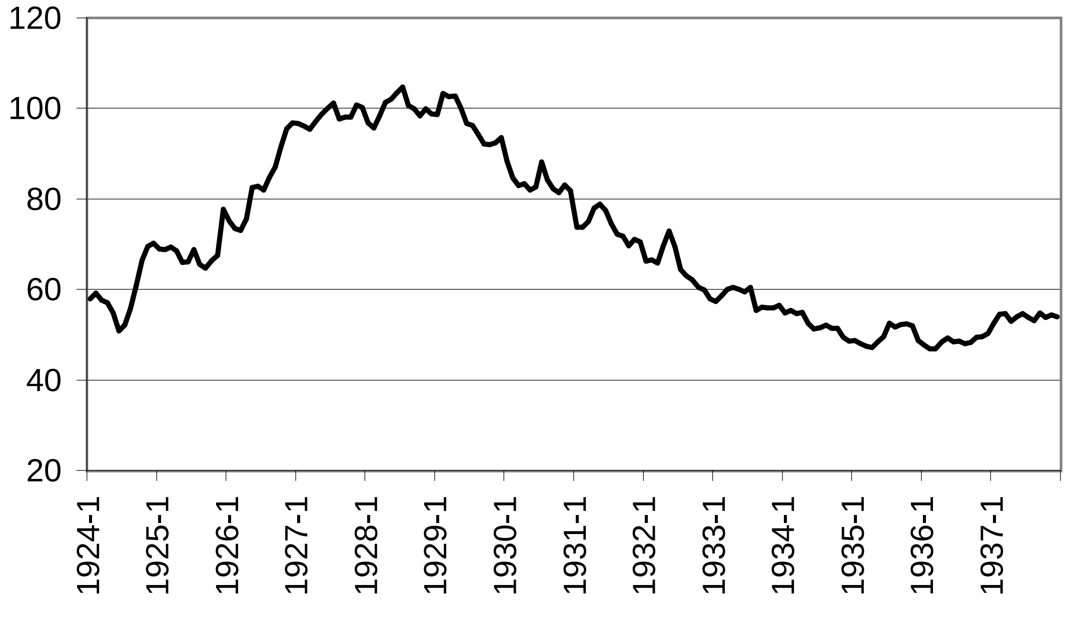

Figure 2. Food price index 1924-1937 (1928=100)

Sources: 1924-1927: own calculations based on Głowny Urząd Statystyczny [Polish Statistical Office], Rocznik Statystyczny [Statistical Yearbook], different years; 1928-1937: Instytut Badania Koniunktur Gospodarczych i Cen [Institute for Business Cycle and Price Research], Koniunktura gospodarcza Polski (zeszczyt specjalny), pp. 22-23.

Table 2. Budget deficit, money supply, and the nominal exchange rate.

Year	Budget deficit[a] (Billion Polish mark)	Money supply[b] (Million Polish mark)	Nominal exchange rate[c] (Polish mark/USD)
1918	—	1,024	9
1919	7,503	5,316	110
1920	54,000	49,316	590
1921	116,000	229,538	2,922
1922	445,000	793,437	17,800
1923	33,000,000	125,371,955	6,375,000

Sources from Landau & Tomaszewski (1999): (a) Kempner (1924); (b and c) Zdziechowski (1925).

Again, a starting point is the war in the east, which required massive revenues and some mechanism to finance them. Since international credit was not yet available, the government had to choose between a "nationalization" of domestic private capital and some mechanism to tax it (Landau & Tomaszewski 1999). The political compromise in 1919 relied on early concessions to the socialists on the one hand (the eight-hour working day was introduced already in November 1918, see Landau 1992) and observing private property rights on the other. As a consequence, the next steps were to create the institutional framework necessary to tax capital and labour: a common currency and a working fiscal administration. As described, these aims were achieved rather quickly. While this was an indisputable success it could not create the necessary revenues to win a war and finance its public infrastructure. But it opened the way for the Polish government to effectively tax money holders by inflation. Table 2 gives some estimations of the budget deficit, of money supply in Polish marks 1918-1923 and the nominal exchange rate with USD.

As the gains from seigniorage and the devaluation of the budget deficit were wiped out by the costs of hyperinflation, namely a complete breakdown of capital markets, the government of prime minister Wladyslaw Grabski had to stabilize the currency. The definite aim was to link the Polish currency with some foreign currency that had successfully restored the gold standard in order to get access to the international capital market. Indeed, Grabski managed to realize this task by help of a temporary property tax, fixed in Swiss gold francs, and several international loans. Already in mid January 1924 the nominal exchange rate was stabilized and a new currency, the Zloty, was fixed at the parity of the Swiss gold francs, i.e. 1 Zloty= 9/31 gram of pure gold. A new institution, the Bank Polski S.A. was introduced with the exclusive right to issue banknotes, while the government kept the right to issue coins (Zbijewski 1931).

However, this currency stabilization gave the government only a short breather. Since the Polish current account, mainly the balance of trade, was negative from its very first statistical registration in 1922, the consequent outflow of capital brought the newly established parity of the Zloty under pressure (see Rocznik Handlu Zagranicznego RP 1922-1925). The problem of a passive balance of trade was intensified by a tariff conflict with Germany that started as the bilateral trade regulations from the Versailles conference ended in January 1925. In this situation, the government was unable to maintain a policy of hard money, and resorted to a mechanism of hidden inflation. In 1925 the government issued paper notes, which were first designed as a substitute for the new Zloty coins and banknotes. When the Bank Polski did not accept these notes in exchange for gold and foreign currency, this resulted in their de facto legalization as a second currency and finally in inflation (see Landau & Tomaszewski 1999, p. 137f.). As a consequence, the Bank Polski had to admit the first of a whole series of devaluations of the Zloty in July 1925 until the exchange rate settled down in May 1926. Formally, only in October 1927 the new parity was fixed at 1 Zloty = 1000/5924,44 gram of pure gold (see Dziennik Ustaw RP 1927).

From now on the government started to defend the parity at any cost. This was possible since the economic policy stopped to follow a path of democratic trade-off after the coup d'état in May 1926, when Jozef Pilsudski installed his authoritarian regime. The new regime launched several programs in order to attract foreign capital to the country. A first success was the stabilization credit of October 1927 over 62 million U.S. dollar and 2 million Pound (Landau & Tomaszewski 1999), followed by some minor credits and an inflow of short term credits till 1929. The improved situation of public finance was in turn used to restore the budget of public enterprises, such as the state railways.

With the onset of the great depression and the dramatic decline of industrial output and aggregate prices, the new policy of hard money was maintained against a growing pressure. Especially the peasants had to suffer, since the decline of agricultural prices was not held in check by cartel agreements, as in the case of industrial prices (see Głowny Urząd Statystyczny 1935). Accordingly, the so-called price scissors squeezed agricultural producers to death. With the exception of some action in the banking sector (Feinstein, Temin & Toniolo 1997), and a rather small-scale program of export promotion and supporting purchase of agricultural products in late 1932, the government pursued a tough liberal policy. The main strategy was quite old-fashioned: improve the balance of trade through a coordinated program of lowering domestic industrial prices, which simultaneously should give relieve to agricultural producers but leave the Zloty parity unchanged (see Knakiewicz 1967). Remarkably, Poland left the gold standard only in 1936, together with Switzerland.

As said above, the following sections explore the extent to which the changes in aggregate prices affected the course of economic integration of Poland, compared to the impact of an improved infrastructure, and a unified institutional framework. The prior from the simple reasoning on arbitrage in the introduction would be that the first years of heavy inflationary tendencies should imply a tendency towards economic integration, while the deflation between 1929 and 1936 should show up in domestic disintegration.

3 Economic Model Framework and the Polish Wheat Market

As mentioned in the introduction we measure economic integration by the LOP. In line with the transaction cost view it implies that arbitrage processes induces price adjustment whenever the price differences exceed the transaction costs. However, due to these costs arbitrage does not equalize the prices completely but only reduces the price differences to the amount of the transaction costs. We now formalize this explanation more precisely.

Let us consider two market places i and j and let N_{ij} denote some export level of a good from place i to j. Furthermore, assume for a moment that the transaction costs mainly consist of transport costs and take the iceberg-form which is used in the recent literature of economic geography. Accordingly, if P_{jt} is the price of the good in location j at time t, then $e^{-\tau}P_{jt}$ is the per-unit revenue when the good is sold in location j where τ is a cost parameter. Hence, $(1-e^{-\tau})P_{jt}$ are the transport costs which "melt away" a portion of the revenue. Intuitively, τ depends positively on the geographical distance between the locations i and j. When border effects are present it also differs depending on whether the locations lie in the same partition area or not. Moreover, one can extend the interpretation of τ to other kinds of transaction costs than transport costs which may relate to social, cultural and technical issues. Finally, let P_{it} be the price in location i, then arbitrage from i to j is only profitable if

$$\begin{aligned} P_{jt}N_{ij}e^{-\tau} &> P_{it}N_{ij} \\ P_{jt}e^{-\tau} &> P_{it} \\ P_{jt}/P_{it} &> e^{\tau} \\ \log(P_{jt}/P_{it}) &> \tau \\ \log(P_{jt}) - \log(P_{it}) = p_{jt} - p_{it} &> \tau \end{aligned} \tag{3.1}$$

Hence, arbitrage from i to j takes place when the log-price difference $p_{jt} - p_{it}$ is larger than the cost parameter τ. Equivalently, one trades from location j to i only if $\log(P_{jt}) - \log(P_{it}) = p_{jt} - p_{it} < -\tau$. Thus, we obtain $[-\tau; \tau]$ as a band of no arbitrage owing to transaction costs. In other words, within this band no trade occurs in order to reduce price differences between the two markets since transaction costs exceed possible arbitrage profits. Obviously, the size of the band increases with τ and may e.g. increase with geographical distance.

Suppose now, that the aggregate price level P_t changes over time and that the nominal transaction costs are fixed at the same time. Then, exporters should care about real rather than nominal profits. This can be seen by defining the transport costs more precisely as $(1 - e^{-\tau})P_{jt} = (1 - e^{-\tau^*/P_t})P_{jt}$. As long as the nominal cost parameter τ^* moves along with the aggregate price level P_t the real transport costs remain the same in proportional terms and it does not matter whether we consider nominal or real prices. However, if τ^* is fixed, the real (proportional) transport costs fall if the aggregate price level increases and they increase if the aggregate price level decreases. Accordingly, we obtain $[-\tau^*/P_t; \tau^*/P_t]$ as a band of no arbitrage. Hence, the size of the band changes in line with the aggregate price level if τ^* is fixed. The band is smaller in times of inflation and larger in times of deflation meaning that inflation fosters and deflation hinders market integration. Thus, $[-\tau^*/P_t; \tau^*/P_t]$ is the relevant transaction cost band as long as the nominal fixed transport costs caused by the unchanged Polish railway tariffs are a crucial part of the whole transaction costs.

We now comment on the characteristics of the Polish wheat market in relation to our economic model framework. Like for other kinds of grain, the cultivation of wheat was unevenly distributed in Poland. The areas around Poznań and Lwów were excess producers and the main net-exporters of wheat. In contrast, the region around Kraków, the south-west and north east parts of Poland were net-importers. Hence, due to these regional differences in production trade of wheat was necessary and took place. Furthermore, the number of grain silos was rather limited and there existed no dense network of silos which limited the ability to store grain over a longer time.

Most of the wheat grinding has been done in small mills. The number of these small mills was rather high and they were evenly spread around the whole country. So, here we can speak of a dense, decentralized network of mills from which the flour was shipped to the cities. Therefore, one may assume that the production stage of flour relied less on freight transport on railways since the distances between the cities and the mills are lower than between the wheat cultivating areas and the mills. Nevertheless, the effect of a changing aggregate price level in the presence of fixed nominal transport costs should be present. This results from the fact that the general price level has still an impact on the millers' and grain traders' ability to exploit spatial price differences in the wheat market.

The description of the historical background has shown that the peasants strongly suffered during the deflationary period. The dramatic fall in the grain prices enforced most of the Polish farmers to sell their grain simply in order to survive. Their financial and technical ability to store grain in large quantities to sell it later at possibly higher prices was clearly limited. The grain traders also suffered from the price decreases. Many went bankrupt and the market concentration increased. These developments created an incentive to establish cartels in order to stop the price fall. However, only with respect to Warsaw there exists some indications that the "Spolka Akcyjna Handlu Ziemioplodami" [Grain Trading Plc.], one of the big trading corporations, tried to organize an agreement in the Warsaw milling industry. But, we do not have direct evidence, whether they succeeded or not (see Srokowski 1939, p. 329 and Śliwa 1935). Nevertheless, we still keep the assumption that the wheat flour prices are the outcome of a competitive market such that arbitrage can adjust the prices.

4 Econometric Framework and Methods

The LOP taking transaction costs into account is usually translated into a threshold cointegration model (see Lo & Zivot 2001). If the strict LOP holds, then the prices of the same good do not differ lastingly at two spatially separated market places which implies that the two price series should be cointegrated: whenever the prices deviate from each other adjustment processes ensure that they turn back to the price parity equilibrium. Since we use variables in logarithms according to our economic model framework, the LOP 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.

However, arbitrage does not occur if the price difference is smaller than the transaction costs. Accordingly, the prices do not adjust within the band of no arbitrage so that the log-prices are not cointegrated. By contrast, outside the band we expect adjustment and, thus, cointegration. These considerations lead to a threshold cointegration model. Referring to the log-price series $p_{1,t}$ and $p_{2,t}$, the transaction cost view of the LOP implies the following symmetric three-regime BAND-threshold autoregressive (BAND-TAR (3)) model:[4]

$$\Delta z_t = \begin{cases} \phi(z_{t-1} - \tau) + \eta_t, & \text{if } z_{t-1} > \tau, \\ \eta_t, & \text{if } -\tau \leq z_{t-1} \leq \tau, \\ \phi(z_{t-1} + \tau) + \eta_t, & \text{if } z_{t-1} < -\tau, \end{cases} \tag{4.2}$$

where $z_t = p_{1,t} - p_{2,t}$ is the log-price difference at time t and $\eta_t \sim$ i.i.d. $(0, \sigma)$. The symmetric threshold band or regime $[-\tau,\tau]$ relates to the band of no arbitrage. Hence, z_t behaves like a random walk within this regime. Its limits are described by the so-called thresholds which coincide with the transport cost parameter and are also labelled as τ. In contrast, in the outer regimes, for which we have $|z_t| > \tau$, economic forces push the prices together implying $-2 < \phi < 0$. This BAND-TAR(3) model imposes the restrictions $\mu_3 = -\phi\tau$ and $\mu_1 = \phi\tau$ on the constants in the outer regimes which guarantee that the prices only adjust to the edge of the transaction band. Moreover, the transaction cost view suggests symmetry regarding the adjustment coefficient ϕ and the threshold τ since arbitrage should be induced in the same way no matter where the prices are higher. If these restrictions are not imposed we obtain the more general TAR(3) model

$$\Delta z_t = \begin{cases} \phi_3 z_{t-1} + \mu_3 + \eta_t, & \text{if } z_{t-1} > \tau_3, \\ \phi_2 z_{t-1} + \mu_2 + \eta_t, & \text{if } \tau_1 \leq z_{t-1} \leq \tau_3, \\ \phi_1 z_{t-1} + \mu_1 + \eta_t, & \text{if } z_{t-1} < \tau_1, \end{cases} \tag{4.3}$$

of which (4.2) is a special case. It is also possible to incorporate lags of Δz_t into (4.3).

Note, that (4.2) is a univariate model with respect to the log-price difference $z_t = p_{1,t} - p_{2,t}$ which is the cointegrating residual regarding the cointegrating vector $(1, -1)$. Lo & Zivot

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

(2001) propose to use a BAND-threshold vector error correction model (BAND-TVECM) instead of a univariate BAND-TAR model. A BAND-TVECM describes the whole dynamics regarding the time series and allows for asymmetries in the adjustment of individual prices to disequilibria. Lo & Zivot (2001) evaluate the relative performance of multivariate and univariate procedures within a threshold cointegration analysis by means of an extensive Monte Carlo study. However, their results do not indicate a general advantage for multivariate procedures. Therefore, we follow a pragmatic approach and apply both univariate and multivariate methods whenever there are reasonable procedures available that may help in answering our questions of interest. We will refer to the results of Lo & Zivot (2001) in the following when explaining the different econometric procedures we have used.

So far we have just considered the simple log-price difference z_t. The presentation of the data in the next section will show that the single series may be characterized by a broken linear trend due to the succession of an inflationary and a deflationary period. Furthermore, the price series may have different levels. The question is whether these deterministic components affect the log-price differences in the sense that we have to include them into the price relationship in order to obtain stationarity. We can still work with the log-price differences as long as the single contain the same deterministics, i.e. the magnitudes of the constants, the linear trends and also of broken components are the same. Then, they cancel out when subtracting the series. But if the deterministics differ between the series, then we have to 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 has important economic interpretations. For example, a constant in z_t^* means that the prices in one city are significantly larger than in the other one. Such systematic differences could be due to different local selling and buying costs which in turn may be caused e.g. by different wage and rent costs. This indicates that certain markets, like e.g. the labor market, are not perfectly integrated on a national or regional level or are characterized by rather high transaction costs. Even if the wheat flour market we look at is integrated, the existence of transaction costs in these other markets prevents economic agents from avoiding the more expensive market places completely. Therefore, we may observe systematic higher prices in one of the respective locations. The discussion is rather important since we use retail prices in cities which may be quite strongly affected by regionally differing cost components like wages and rents.

If a trend enters the cointegrating relationship the systematic price differences increase over time. Accordingly, a broken deterministic component suggests that the change from inflation to deflation caused a shift in the price pattern possibly due to the impact of the changing aggregate price level on market integration or cost levels. Hence, deterministic terms in the log-price relationship indicate a general lack of market integration. Therefore, we address this issue in the empirical analysis although we are not able to asses to which extend the deterministic price differences have to be attributed to the characteristics of the wheat flour market. In any case, by allowing for deterministic terms we consider a relative version of the LOP in the sense that adjustments still occurs, but only in line with the extended price relationship and not towards the the price parity. Note, that the discussed systematic price differences have to be distinguished from the effects of transaction costs which prevent arbitrage from equalizing the prices completely. Transaction costs refer to the occurrence of adjustment but the systematic price differences affect the equilibrium toward which adjustment takes place.

As mentioned in the introduction we perform the threshold cointegration analysis in three steps according to Lo & Zivot (2001) and Balke & Fomby (1997). First we test for cointegration, then for threshold nonlinearity, and finally the threshold models are estimated provided that we found cointegration and nonlinearity.

To test for cointegration we apply a generalization of the multivariate Johansen testing procedure which allows for broken linear trends and levels. This generalization has been proposed by Johansen, Mosconi & Nielsen (2000). It does not only enables us to test for cointegration in a more general setup of deterministic terms but also allows us to test whether certain deterministic components are present and to which extent they affect the price cointegration relationship according to the foregoing discussion. Additionally, we can test within the Johansen procedure 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. This information is also important for the further econometric analysis since some of the procedures require a known cointegrating vector.

Assuming one break in the deterministic terms at time $t = T_1$, the Johansen procedure is based on a maximum likelihood estimation of the linear n-dimensional VECM model

$$\begin{aligned}\Delta y_t = \alpha(\beta' y_{t-1} - \theta_1(t-1)D_{1,t} - \theta_2(t-1)D_{2,t}) + \nu_1 D_{1,t} + \nu_2 D_{2,t} + \gamma_2 d_{2,t} + \varepsilon_t,\\ t = p+1, p+2, \ldots, T,\end{aligned} \tag{4.4}$$

where $D_{1,t}$ is one for all observations before T_1 and zero otherwise, $D_{2,t} = 1 - D_{1,t}$, $d_{2,t}$ is one for $t = T_1$ and zero otherwise. Hence, these variables describe the two regimes before and after the break in the deterministic components. Moreover, θ_1, θ_2, ν_1 and ν_2 are $(n \times 1)$ parameter vectors related to the linear trends and constants of the two regimes, and $\varepsilon_t \sim N(0, \Omega)$. The Johansen procedure tests for the rank r of the matrix $\Pi = \alpha\beta'$, where α $(n \times r)$ is the matrix of adjustment coefficients and the matrix β $(n \times r)$ contains the coefficients of the cointegrating vectors related to the variables y_t. Hence, the rank r determines the number of cointegration relations. 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 values of the test can be computed by using a response surface given in Johansen et al. (2000). To simplify the exposition we have ignored any short-run dynamics, i.e. no lags of Δy_t and $d_{2,t}$ are considered.

In (4.4) the trends are included in the cointegration relations whereas the constants are not. This is the appropriate model representation if different linear trends exist in the two regimes and if a quadratic trend is ruled out. Provided we found a cointegrating rank of one, we can perform restriction tests on the corresponding parameters θ_1, θ_2, ν_1, and ν_2 to study which deterministic terms are present and enter the cointegrating relationship. Similarly, we test whether β can be restricted to $(1, -1)$. All these restriction tests are asymptotically $\chi^2(k)$ distributed where k 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). We will be more precise on the sequence of tests when we describe the empirical results in the next section.

Since the Johansen test is a linear cointegration test it may have low power if threshold cointegration is the appropriate alternative. Therefore, we also use a procedure by Berben & van Dijk (1999) (BVD test) which tests the null of a unit root against a stationary two-regime TAR model. The test has a sup-F-type form comparing the sum of squared residuals under the null and the alternative hypothesis. Since the threshold parameter is not identified under the null hypothesis of linearity critical values have to be determined by bootstrap methods for each single case.

Because the BVD test is a unit root test we apply it to the residuals of the cointegrating relationship assuming a known vector $(1, -1)$. In case of an estimated vector the asymptotic distribution of the test statistic may not hold. Note, that we can only pretend to know the vector even if the respective Johansen restriction test does not reject the vector $(1, -1)$. Nevertheless, we use the test because it has the highest small sample power in the Monte Carlo study by Lo & Zivot (2001) and is clearly superior to the unit root test by Enders & Granger (1998) which is based on a biased estimate of the threshold under the alternative.[5] Although the BVD test is designed for two-regime TAR alternatives the high power was obtained for three-regime TAR data generating processes (DGPs). Lo & Zivot (2001) explain this outcome by results of Bai (1997) which say that the threshold estimate from a two-regime model is consistent for one of the two thresholds in a three regime model.

If cointegration is found the next step is to test whether the dynamics of the data can be described by a threshold model. Rejection of the null hypothesis by the BVD test would already suggest threshold effects according to the alternative hypothesis. However, the BVD test may also have power against linear cointegration and we want to test for three-regime threshold dynamics as well. Therefore we apply explicit threshold nonlinearity tests.

We first use the univariate and multivariate tests suggested by Tsay (1989, 1998). The idea of these procedures is to arrange the data according to the value of the threshold variable (in our case z_{t-1}) and to perform an autoregression based on these arranged data. The rearrangement does not change the dynamic relationship between the dependent variable and its lags but if the data follow a threshold model, the thresholds translate to structural breaks in the arranged data. The statistics testing for these breaks are asymptotically F (univariate test) and χ^2 (multivariate test) distributed. 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 more small sample power if it is the appropriate alternative. Based on nested hypotheses Hansen (1997, 1999) proposes to test the null of a univariate linear AR model against a stationary two-regime TAR model and a three regime TAR model respectively. The procedures have a sup-F-type form like the BVD test and critical values have to be computed by bootstrap methods as well. We use the bootstrap for homoscedastic error terms. In line with the argumentation for the BVD test both versions testing against two- and three-regime alternatives have comparable small sample power in the simulation study of Lo & Zivot (2001).

[5]Enders & Siklos (2001) have found that their cointegration test allowing for an unknown cointegrating vector has rather low power for the TAR model. Therefore, we do not apply their procedure.

As pointed out by Hansen & Seo (2001) the univariate tests of Tsay (1989) and Hansen (1997, 1999) are only known to be valid if the cointegrating vector is known. The corresponding arguments regarding the BVD test made above apply here again. That is why, we also use the multivariate SupLM test by Hansen & Seo (2001) which allows for an unknown cointegrating vector. However, the results of their procedure do not give additional insights. Therefore we do not comment on this test here in detail.

If the tests indicate threshold nonlinearity one can proceed to estimate the threshold models. A reasonable strategy would be to estimate first the unrestricted model (4.3) for the cointegrating residual z_t and then to test for the restrictions on the model parameters implied by the transaction cost view. Unfortunately, the results of Lo & 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 & Zivot (2001) conclude that these procedures are essentially useless. So, we are left with simply comparing the estimation results of the unrestricted and restricted threshold models (4.3) and (4.2).

We estimate TAR models via sequential conditional least squares methods. First, a two-regime model is estimated. For that purpose the possible values for the threshold are restricted to the values of threshold variable z_{t-1}. Then, the model is estimated for each possible threshold value and the value minimizing the respective sum of squared residuals (SSR) is taken as the estimate. Afterwards, a three-regime model is estimated in the same way given the first threshold estimate which is consistent for one of two thresholds (see Bai 1997). The pair of threshold values minimizing the SSR are taken as the estimates and the estimates of the other parameters are automatically obtained by applying the threshold estimates. Within the estimation procedure it is assured that each regime contains a minimum number of observations. Following the literature, we let the minimum number to be equal to 10% of the total number of observations. The reader is referred to Hansen (1999) and Lo & Zivot (2001) for more details on the estimation of TAR models.

Finally, we describe how we evaluate the effects of the old borders and the changing aggregate price level within the econometric framework introduced.

To analyze the impact of the former partition borders we can distinguish between the four within-border city pairs Warsaw-Lodz, Warsaw-Wilno, Wilno-Lodz (all former Russian part) and Kraków-Lwów (Austrian part) and the remaining eleven across-border pairs. If their exist systematic border effects we first expect to find less evidence for cointegration between the flour prices of the across-border compared to the within-border pairs. Less evidence refers both to a relatively lower number of cointegrated pairs and to the significance level of the respective test statistics. Obviously, a weaker indication of cointegration means that the old borders hinder price adjustment. Secondly, the adjustment coefficient ϕ of the outer regimes in (4.2) should be higher for across-border pairs if the former borders matter. Finally, in case of borders effects, the sizes of the threshold bands for across-border pairs should be larger than for within-border pairs when corrected for the distance between the cities. As mentioned, inference about the model parameters is heavily size distorted in small samples. Therefore we cannot base the comparison of the estimates for ϕ on reliable test procedures. The same applies regarding the size of the threshold band because no corresponding asymptotic theory for threshold estimates has been derived in the three-regime framework.[6]

In the previous section we have explained that $[-\tau^*/P_t; \tau^*/P_t]$ is the relevant band of no arbitrage when taking account of the aggregate price level P_t and the nominal transport cost parameter τ^*. If the nominal fixed railway costs are the main part of the whole transaction costs, the band of no arbitrage changes with the aggregate price level since τ^* is fixed. This, however, means that we are supposed to have a changing threshold band, i.e. a model with thresholds changing over time. Unfortunately, the TAR approach assumes fixed thresholds. Therefore, we suggest to use the log-price series multiplied with the aggregate price level P_t, i.e. the series $p_{1t}P_t$ and $p_{2t}P_t$, instead of the simple log-prices in order to generate a TAR model with a stable threshold band.

To see this, remember that in line with the band $[-\tau^*/P_t; \tau^*/P_t]$ no adjustment in the log-prices takes place whenever $|z_{t-1} = p_{1,t-1} - p_{2,t-1}| < \tau^*/P_t$. Instead of comparing the threshold variable z_{t-1} with τ^*/P_t we multiply z_{t-1} by P_t and compare the new threshold variable $z_{t-1}P_t$ with τ^* which, by assumption, is fixed. Hence, instead of building threshold models with respect to $p_{1,t}$ and $p_{2,t}$ we consider $p_{1t}P_t$ and $p_{2t}P_t$.[7]

[6]Only for two-regime TAR models Hansen (1997) has suggested a likelihood ratio statistic for testing hypothesis concerning the threshold with an asymptotic distribution free of nuisance parameters.

[7]If the nominal fixed costs are a only a minor part of the whole transaction costs, then τ^* changes over time; but, depending on the share of the fixed costs, not as strong as the aggregate price level P_t. Thus, we obtain a TAR model with changing thresholds for both the adjusted and the simple log-prices.

Intuitively, the changing thresholds are modelled by the multiplicative transformation through the fact that disequilibria between $p_{1,t}$ and $p_{2,t}$ are scaled by the price level. In times of high values of P_t (during inflation) price differences are inflated in relative terms. Hence, a certain observation is more likely to lie in one of the outer regimes so that adjustment according to the parameter ϕ in (4.2) takes place. If the observation belongs already to the outer regimes, adjustment is stronger due to the inflated price differences. This simulates the smaller threshold band in times of increasing price levels and corresponds to the fostering power of inflation with respect to market integration in case of nominal fixed transaction costs. An opposite explanation can be given for a falling aggregate price level which deflated the price differences in relative terms.

To figure out whether the aggregate price level plays the assumed role in the integration process we propose to run the threshold analysis for both the simple log-price and price level adjusted series. Concerning the cointegration analysis we expect stronger evidence for cointegration using the adjusted series since the Johansen and BVD tests should have more small sample power if they are applied to data following a model with fixed thresholds. This can be assumed because the tests' alternative hypotheses are closer to a framework of stable thresholds.

When applying the threshold nonlinearity tests of Tsay (1989, 1998) we assume to obtain similar results since these tests are independent of the specific threshold structure. In fact, changing thresholds translate to multiple breaks in the arranged autoregression. Therefore, the Tsay tests should not have lower power when using the simple log-price series in contrast to the Hansen tests which consider a specific TAR model with fixed thresholds under the alternative. Note, that the BAND-TAR(3) model implied by the LOP is nested in an unrestricted TAR(3) model. Therefore, we assume that especially the Hansen test with this TAR(3) alternative should reject the null of linearity more easily when the adjusted series are used. With respect to the estimation of the threshold models we expect much higher estimation uncertainty, i.e. higher standard errors of the estimated parameters, when using the unadjusted log-prices. Furthermore, the estimated parameters should be less in line with the model restrictions implied by the transaction cost view of the LOP. As for the analysis of the border effects the evaluation of the model estimates cannot be based on reliable testing procedures.

5 Empirical Results

5.1 Data and Preliminary Analysis

We use monthly retail prices per kg of wheat flour expressed in Groszy (= 0.01 Zloty) from 1924:01-1937:04, so we have 160 observations. The start and ending dates are determined by data availability. The prices were reported to the GUS in the last week of the respective month. The data are taken from the GUS' publications *Rocznik Statystyczny* [Statistical Yearbook] and *Statystyka Cen* [Price Statistics], and from Koniunktura gospodarcza Polski which is a publication of Instytut Badania Koniunktur Gospodarczych i Cen [Institute for Business Cycle and Price Research].

The log-price series named *logwar*, *logwil*, *loglodz*, *logpos*, *logkrak*, and *loglwów* are shown in Figure 3. We clearly see a price increase until 1927 followed by a short period of stabilization. Then, starting from 1929 on, in line with the great depression, the prices fell dramatically until 1936 interrupted by a an increase in 1933. Finally, in 1937 prices went up again. The general movement of all time series is similar. However, they may have a different level. For example, the prices in Lwów seem to be the lowest ones for most of the observations.

Furthermore, we use the food-price index (FPI) (1928=100) to perform the adjustment of the log-prices in order to consider the effects of the changing aggregate price level. As mentioned in Section 2 the FPI is the most complete price index available for interwar Poland on a monthly basis. Obviously, we would prefer to use a more general price index to capture the development of the aggregate price level. In this connection, one has to note that the FPI overestimates the deflation in the 1930s since the prices of non-agricultural goods fell less owing to cartel agreements. Moreover, the share of wheat in the FPI, which was slightly less than 5%, is much larger than in a more general consumer price index.

Figure 4 displays the development of the FPI-adjusted log-price series *fpiwar*, *fpiwil*, *fpilodz*, *fpipos*, *fpikrak*, and *fpilwów* which are obtained by multiplying the log-price series with the FPI for each period t. In the following these series are named as *fpi*-series. Obviously, they move rather similar and their general pattern is comparable to the FPI itself which is shown in Figure 2. Clearly, the upward and downward movement in the adjusted prices is much stronger than for the log-prices. Thus, broken deterministic components are likely to be more relevant for the former set of series.

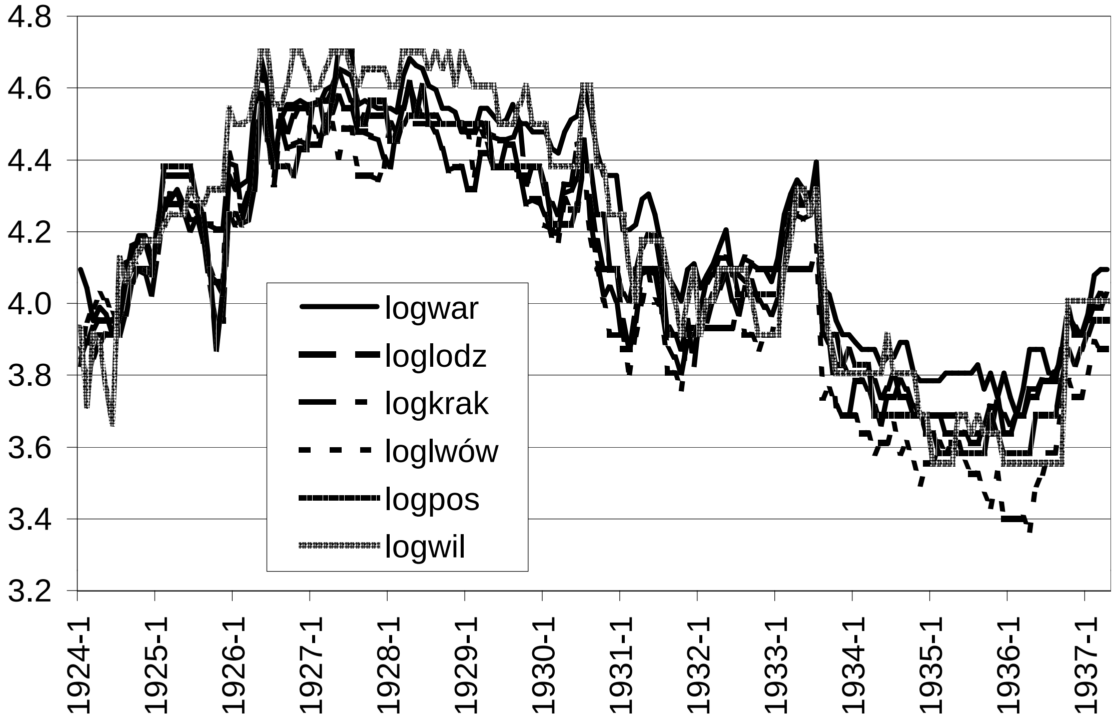

Figure 3. Food price index 1921-1923 (1921=100)

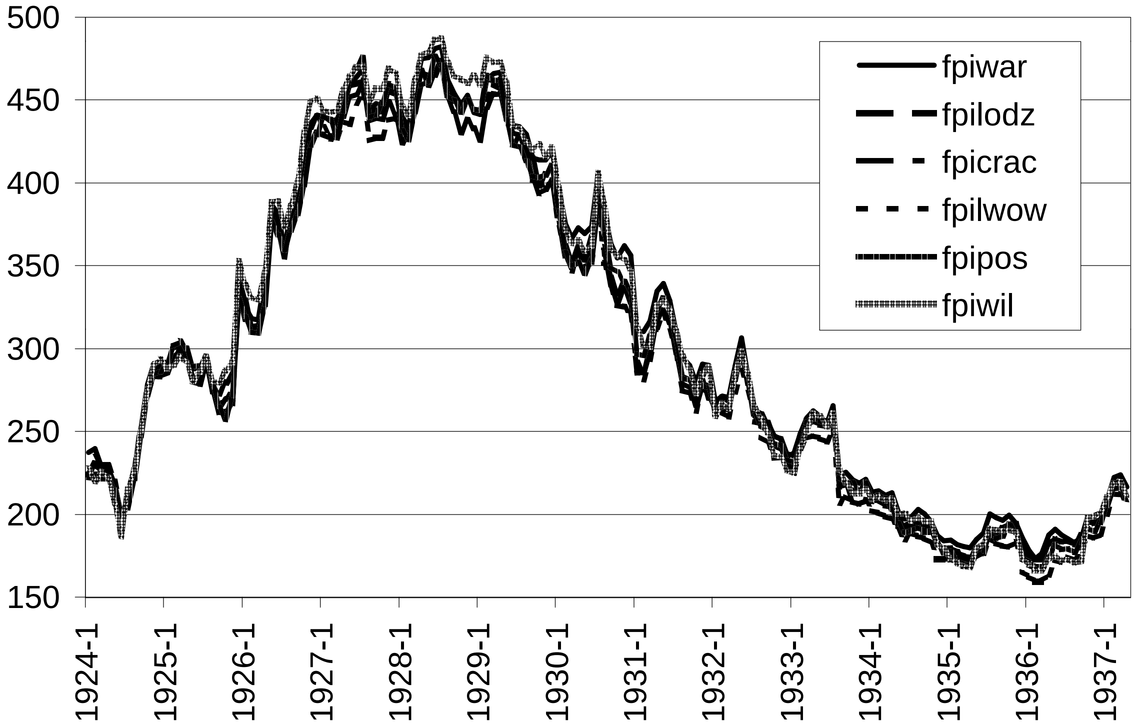

Figure 4. Food price index 1924-1937 (1928=100)

The results presented in the following are obtained by using different econometric software packages and computer programs. The unit root analysis has been done with EViews 4.1 and the cointegration analysis with PcFiml 9.10 (see Doornik & Hendry 1997). To compute the BVD and Tsay nonlinearity test statistics and their p-values and to perform the estimation of the BAND-TAR model (4.2) we use own GAUSS programs. We have applied GAUSS programs from Bruce Hansen's web page (http://www.ssc.wisc.edu/∼bhansen/progs/progs_threshold.html) to compute the test statistics for the procedures by Hansen (1997, 1999) and their respective bootstrap p-values as well as to estimate the unrestricted TAR model (4.3).

Before continuing with the empirical analysis we want to comment on seasonality effects. This issue is rather important since different seasonal patterns in the price series would raise doubts about market integration. First, we apply a test by Canova & Hansen (1995) to analyse whether deterministic or stochastic seasonality is present. The test clearly suggests that seasonality is deterministic for all series. Then, we proceed to estimate univariate AR models including seasonal dummies for the first differences of the *log* and *fpi*-series in order to test for significance of these dummy variables. Regarding the log-prices the dummy variables are jointly significant only for Warsaw (10% level). This surprising result may be due to the fact that flour can be gained from both summer and winter wheat which have the same degree of grinding. Therefore, they can be regarded as perfect substitutes. Hence, only smaller storage capacities are required to eliminate seasonal price differences. Accordingly, the log-price differences of all 15 city pairs are not affected by deterministic seasonality.

However, the FPI has a significant seasonal pattern and this carries over to the *fpi*-series owing to the multiplication with the FPI. But only the dummy variables representing the months January and August are significant. Nevertheless, the price differences for the pairs *fpiwil*-*fpilodz*, *fpiwar*-*fpilodz*, and *fpilodz*-*fpicrac* contain significant seasonal effects. As a response, we also allow for seasonal dummies in the unit-root and cointegration analyses and adjust the *fpi*-series by regressing the cointegrating residuals on seasonal dummies before we estimate the threshold models. This modelling approach has not changed the general findings. For this reason and because we do not know whether the nonlinearity tests are still valid if seasonal dummies are added we only present the results obtained without the dummy variables. Finally, it seems that different seasonal patterns are mainly due to

Table 3. Unit Root Test statistics

	Level Series	First Differences		Level Series	First Differences
fpiwar	−2.30(2)	−9.86(1)***	*logwar*	−2.26(3)	−8.61(2)***
fpiwil	−2.58(1)	−9.36(1)***	*logwil*	−2.97(0)	−13.80(0)***
fpipos	−2.53(5)	−5.48(4)***	*logpos*	−3.64(1)	−9.36(1)***
fpilodz	−2.47(5)	−5.32(4)***	*loglodz*	−3.39(1)	−8.71(1)***
fpicrac	−1.94(3)	−7.50(3)***	*logcrac*	−2.55(0)	−12.25(0)***
fpilwów	−2.41(2)	−9.72(1)***	*loglwów*	−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 for the first differences are −3.74 (5%) and −4.34 (1%). They are taken from the Tables IV.B and VI.B in Perron (1989) respectively and relate to the relative break point $\lambda = 65/160 = 0.4$.

data transformation. Therefore, we do not relate them to a lack of market integration.

In Table 3 the results of the unit root analysis are summarized. 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 in Perron & Vogelsang (1993). This procedure is a generalization 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 the constant (Model C in Perron (1989)) and for the first differences the version with a break in the constant only (Model A in Perron (1989)). As the break date we choose the observation May 1929 since from this month on the FPI started to fall so that May 1929 is the turning point from the inflationary to the deflationary period. A Chow breakpoint test confirmed this break date. Accordingly, we have 64 and 96 observations for the two periods respectively. Obviously, all time series can be regarded as integrated of order one since the null hypothesis of a unit root is not rejected for the level series but rejected for the first differences.

5.2 Results of Cointegration Analysis

The results of the generalized Johansen procedure for the *fpi*-series are given in Table 4. The misspecification test for vector autocorrelation described in Doornik & Hendry (1997) suggests no significant autocorrelation for the corresponding vector autocorrelation (VAR) models. We see that all city pairs have a cointegrating rank of one but the pairs $fpiwar$ −

$fpipos$ and $fpiwar - fpikrak$ cointegrate at a 10% level only. In a next step, we test whether the cointegrating vector β can be restricted to (1, -1). The results in the 4th column of Table 4 show that this assumption cannot be rejected for almost all of the city pairs applying a 5% significance level. Hence, we can conclude that the price differences of the respective city pairs form a stationary relationships as the LOP implies.

But which deterministic terms enter this relationship? To answer this question we have first tested whether the trend coefficients θ_1 and θ_2 in the cointegrating relationship can be both set to zero. With the exception of $fpilodz - fpilwów$ this restriction is not rejected. Then, we checked whether linear trends can be completely excluded from the model implying that the constants can be restricted to the cointegrating relationship. Technically, this restriction translates to the hypotheses $\nu_i = \Pi\mu_i$ $(i = 1, 2)$. This is rejected for 8 of the 15 city pairs and we have borderline cases close to the 10% significance level for most of the other pairs. Furthermore, when testing separately for the inflationary or deflationary periods we find that a linear trend is present in at least one of the periods. Additionally, we figure out that the trend components for both periods are different for all city pairs. Therefore, we proceed with the assumption that the VECMs for all pairs contain a broken linear trend. However, the trend components do not enter the cointegrating relations since the price differences eliminate them, i.e. the broken trend is orthogonal to the cointegration space. Only for the pair Lodz-Lwów a broken trend is present in the cointegrating relation.

The fact that the trend components are orthogonal to the cointegrating space requires that the constants are outside the cointegrating relationship in order to generate a broken linear trend in the level of the data.[8] In general it is possible to decompose (the parameter of) the constant into parts inside and outside the cointegrating relationship (compare Johansen (1994, 1995) for a more precise definition). However, we cannot test for significance of the component inside the relationship since it is not identified. Therefore, we just state the means of the price differences. Since the restriction tests suggest broken components we present the means for both the inflationary and deflationary period. Obviously, the means of most of the pairs differ between the periods. Relating the largest observed constant of 13.10 to the average value of all fpi-series in the sample, which is 301, it follows that the systematic

[8]Note, that the VECM (4.4) is written in first differences so that an unrestricted constant can generate a linear trend in the data. Accordingly, a linear trend is ruled out if a constant is restricted to the cointegrating relation which is stated in levels.

Table 4. Results of Generalized Johansen Procedure for fpi-series

City pair (k)	AR(1-5) F(20,290-8k) p-value	$H_0(r_0)$	Test statistic (p-value)	$\beta = (1, -1)$ $\chi^2(1)$ p-value	$\theta_1 = \theta_2 = 0$ $\chi^2(2)$ p-value	$(\nu_1 : \nu_2) = \Pi(\mu_1 : \mu_2)$ p-value	mean of price-diff. 1924:01-1929:04	1929:05-1937:04
$fpiwar - fpiwil$ (1)	0.101	$r_0 = 0$	52.42(0.000)***	0.466	0.474	0.032**	−5.42	4.82
		$r_0 = 1$	9.72(0.362)					
$fpiwar - fpipos$ (3)	0.097	$r_0 = 0$	34.41(0.057)*	0.120	0.463	0.112	4.47	8.62
		$r_0 = 1$	8.17(0.508)					
$fpiwar - fpilodz$ (1)	0.178	$r_0 = 0$	42.02(0.007)***	0.434	0.224	0.028**	7.68	9.49
		$r_0 = 1$	10.29(0.316)					
$fpiwar - fpikrak$ (2)	0.190	$r_0 = 0$	32.92(0.081)*	0.231	0.847	0.226	2.41	6.80
		$r_0 = 1$	11.71(0.219)					
$fpiwar - fpilwów$ (1)	0.108	$r_0 = 0$	43.32(0.005)***	0.075*	0.638	0.057*	4.71	14.97
		$r_0 = 1$	9.98(0.340)					
$fpiwil - fpipos$ (2)	0.153	$r_0 = 0$	44.77(0.003)***	0.022**	0.680	0.222	9.83	3.80
		$r_0 = 1$	10.34(0.312)					
$fpiwil - fpilodz$ (2)	0.110	$r_0 = 0$	40.45(0.012)**	0.195	0.497	0.090*	13.10	4.67
		$r_0 = 1$	9.46(0.385)					
$fpiwil - fpikrak$ (2)	0.277	$r_0 = 0$	39.12(0.017)**	0.055*	0.929	0.171	7.83	1.98
		$r_0 = 1$	9.39(0.391)					
$fpiwil - fpilwów$ (2)	0.109	$r_0 = 0$	38.45(0.020)**	0.030**	0.630	0.120	10.13	10.15
		$r_0 = 1$	9.59(0.373)					

Table 4. cont'd. Results of Generalized Johansen Procedure for fpi-series

City pair (k)	AR(1-5) F(20,290-8k) p-value	$H_0(r_0)$	Test statistic (p-value)	$\beta = (1, -1)$ $\chi^2(1)$ p-value	$\theta_1 = \theta_2 = 0$ $\chi^2(2)$ p-value	$(\nu_1 : \nu_2) = \Pi(\mu_1 : \mu_2)$ p-value	mean of price-diff. 1924:01-1929:04	1929:05-1937:04
$fpipos - fpilodz$ (3)	0.110	$r_0 = 0$	51.10(0.000)***	0.321	0.122	0.020**	3.27	0.87
		$r_0 = 1$	12.06(0.199)					
$fpipos - fpikrak$ (3)	0.085	$r_0 = 0$	44.20(0.004)***	0.430	0.716	0.114	−2.00	−1.82
		$r_0 = 1$	11.74(0.217)					
$fpipos - fpilwów$ (3)	0.181	$r_0 = 0$	36.31(0.036)**	0.894	0.456	0.117	0.30	6.35
		$r_0 = 1$	11.98(0.203)					
$fpilodz - fpikrak$ (1)	0.087	$r_0 = 0$	53.63(0.000)***	0.446	0.159	0.016**	−5.27	−2.69
		$r_0 = 1$	10.25(0.324)					
$fpilodz - fpilwów$ (1)	0.219	$r_0 = 0$	47.85(0.001)***	0.142	0.032**	0.005***	−2.97	5.48
		$r_0 = 1$	10.19(0.324)					
$fpikrak - fpilwów$ (1)	0.132	$r_0 = 0$	46.87(0.002)***	0.304	0.209	0.017**	2.30	8.17
		$r_0 = 1$	10.40(0.307)					

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 & Hendry 1997). The p-values for the cointegration test statistics are obtained from the response surface given in Johansen et al. (2000) for a relative break point $\lambda = 65/160 = 0.4$. ***, **, ** denote significance at the 1%, 5%, and 10% level respectively.

price differences between two cities can amount up to more than 4% of this average. Since the series are transformed in a nonlinear way it is not trivial to translate this figure to the original price data. Nevertheless, we think that broken constants are relevant for most of the price relationships.

Summarizing, we can conclude that there is evidence for the validity of a relative version of the LOP since the price differences adjust appropriately to price disequilibria. However, the price differences seem not eliminate the deterministic terms in the cointegrating relations. Furthermore, no systematic differences between the across-border and within-border pairs can be observed, although the two pairs for which we found cointegration only at the 10% level are across-border pairs.

The results for the *log*-series are summarized in Table 5. As for the *fpi*-series we find strong evidence for cointegration with respect to the price-differences. The evidence seems to be even slightly stronger in contrast to our prior about the effect of a changing aggregate price level in the presence of nominal fixed transaction costs. With the exception of five city pairs, always including Lodz or Lwów, we can exclude the trend from the cointegrating relationship and also from the whole model. Thus, it seems that the series of these pairs are not affected by a trend at all.[9] However, the constants in the cointegration relations cannot be restricted to zero with the exception of $logpos - logcrac$. But the constants are the same for both periods except for the pairs including Warsaw for which the restricted constants differ. The same applies with respect to the pairs including trend components in the price-relationship. For $logwar - loglodz$ and $logwar - loglwów$ the trend is broken and for the other three pairs the trend is the same for both the inflationary and deflationary period. The latter results regarding the deterministic terms are not reported here in detail. We only state the values and standard errors of the restricted constants and the simple mean of the log-price differences in case of unrestricted constants in the last column of Table 5. The largest reported mean of 0.137 is equal 3.3% of the average value of all *log*-series.

Finally, we apply the unit root test by Berben & van Dijk (1999) to the price-differences for both the *fpi*- and *log*-series since a cointegrating vector (1, -1) has not been rejected

[9]This result is somewhat surprising with respect to $logwil - loglodz$, $logpos - loglodz$, and $logpos - loglwów$. If a trend is caused by the Lodz and Lwów series, as the outcomes of the restriction tests regarding $\theta_1 = \theta_2 = 0$ suggest, then also these pairs should require the inclusion of a linear trend into the model. However, the p-values with respect to $logwil - loglodz$ and $logpos - loglodz$ are borderline cases (compare 6th column). Probably, the tests may suffer from low small sample power.

Table 5. Results of Generalized Johansen Procedure for *log*-series

City pair (k)	AR(1-5) F(20,290-8k) p-value	$H_0(r_0)$	Test statistic (p-value)	$\beta = (1, -1)$ $\chi^2(1)$ p-value	$\theta_1 = \theta_2 = 0$ $\chi^2(2)$ p-value	$(\nu_1 : \nu_2) = \Pi(\mu_1 : \mu_2)$ p-value	mean of price-diff. 1924:01-1929:04	mean of price-diff. 1929:05-1937:04
logwar − *logwil* (1)	0.225	$r_0 = 0$	48.72(0.001)***	0.066*	0.921	0.791	−0.072	0.095
		$r_0 = 1$	7.13(0.612)				(0.027)	(0.022)
logwar − *logpos* (2)	0.134	$r_0 = 0$	40.72(0.011)**	0.941	0.718	0.847	0.137	0.034
		$r_0 = 1$	11.41(0.233)				(0.016)	(0.020)
logwar − *loglodz* (1)	0.111	$r_0 = 0$	50.04(0.001)***	0.186	0.009***	0.033**	0.087	0.138
		$r_0 = 1$	11.04(0.260)					
logwar − *logkrak* (1)	0.501	$r_0 = 0$	49.27(0.001)***	0.614	0.315	0.373	0.026	0.096
		$r_0 = 1$	10.29(0.311)				(0.019)	(0.015)
logwar − *loglwów* (1)	0.562	$r_0 = 0$	49.61(0.001)***	0.962	0.030**	0.076*	0.048	0.236
		$r_0 = 1$	9.84(0.347)					
logwil − *logpos* (1)	0.106	$r_0 = 0$	39.35(0.013)**	0.622	0.771	0.614	0.072	
		$r_0 = 1$	7.55(0.567)				(0.023)	
logwil − *loglodz* (1)	0.406	$r_0 = 0$	51.02(0.000)***	0.509	0.170	0.319	0.090	
		$r_0 = 1$	9.79(0.351)				(0.028)	
logwil − *logkrak* (1)	0.717	$r_0 = 0$	65.43(0.000)***	0.167	0.503	0.517	0.042	
		$r_0 = 1$	10.59(0.288)				(0.019)	
logwil − *loglwów* (1)	0.319	$r_0 = 0$	63.72(0.000)***	0.111	0.041**	0.090*	0.132	
		$r_0 = 1$	11.28(0.241)					

Table 5. cont'd. Results of Generalized Johansen Procedure for *log*-series

City pair (k)	AR(1-5) F(20,290-8k) p-value	$H_0(r_0)$	Test statistic (p-value)	$\beta = (1, -1)$ $\chi^2(1)$ p-value	$\theta_1 = \theta_2 = 0$ $\chi^2(2)$ p-value	$(\nu_1 : \nu_2) = \Pi(\mu_1 : \mu_2)$ p-value	mean of price-diff. 1924:01-1929:04 1929:05-1937:04
logpos − *loglodz* (2)	0.264	$r_0 = 0$	46.55(0.002)***	0.643	0.112	0.249	0.019
		$r_0 = 1$	15.27(0.074)*				(0.015)
logpos − *logkrak* (2)	0.581	$r_0 = 0$	45.60(0.003)***	0.288	0.333	0.519	−0.029
		$r_0 = 1$	13.43(0.131)				(0.012)
logpos − *loglwów* (2)	0.537	$r_0 = 0$	39.93(0.013)**	0.710	0.460	0.673	0.066
		$r_0 = 1$	14.90(0.083)*				(0.022)
loglodz − *logkrak* (1)	0.365	$r_0 = 0$	51.72(0.000)***	0.477	0.136	0.235	−0.049
		$r_0 = 1$	9.06(0.416)				(0.009)
loglodz − *loglwów* (1)	0.348	$r_0 = 0$	48.16(0.001)***	0.313	0.001***	0.006***	0.043
		$r_0 = 1$	9.55(0.372)				
logkrak − *loglwów* (1)	0.841	$r_0 = 0$	45.10(0.003)***	0.567	0.025**	0.053*	0.091
		$r_0 = 1$	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 & Hendry 1997). The p-values for the cointegration test statistics are obtained from the response surface given in Johansen et al. (2000) for a relative break point $\lambda = 65/160 = 0.4$. ***, **, ** denote significance at the 1%, 5%, and 10% level respectively. The means and the standard errors given in parenthesis in the last two columns refer to the estimation of the respective VECM if the constants could be restricted to the cointegrating relationship. For the other pairs only the simple means of price differences are reported. If the restriction tests do not indicate broken deterministic component the mean refers to the whole period.

Table 6. Results of Berben-van Dijk Test

City pairs (k)	Test statistics (p-value)	City pairs (k)	Test statistics (p-value)
$fpiwar - fpiwil$ (0)	14.06(0.000)***	$logwar - logwil$ (0)	13.35(0.002)***
$fpiwar - fpipos$ (0)	13.48(0.002)***	$logwar - logpos$ (1)	8.88(0.005)***
$fpiwar - fpilodz$ (0)	6.32(0.059)*	$logwar - loglodz$ (0)	9.32(0.009)***
$fpiwar - fpikrak$ (1)	5.49(0.085)*	$logwar - logkrak$ (1)	6.38(0.050)**
$fpiwar - fpilwów$ (1)	5.01(0.120)	$logwar - loglwów$ (1)	5.40(0.095)*
$fpiwil - fpipos$ (0)	11.62(0.001)***	$logwil - logpos$ (0)	17.23(0.001)***
$fpiwil - fpilodz$ (0)	8.87(0.007)***	$logwil - loglodz$ (0)	8.94(0.009)***
$fpiwil - fpikrak$ (1)	11.83(0.002)***	$logwil - logkrak$ (0)	11.70(0.000)***
$fpiwil - fpilwów$ (1)	8.04(0.014)**	$logwil - loglwów$ (1)	18.35(0.003)***
$fpipos - fpilodz$ (0)	15.83(0.000)***	$logpos - loglodz$ (0)	14.16(0.000)***
$fpipos - fpikrak$ (0)	20.03(0.000)***	$logpos - logkrak$ (1)	11.22(0.001)***
$fpipos - fpilwów$ (0)	13.38(0.001)***	$logpos - loglwów$ (1)	8.80(0.006)**
$fpilodz - fpikrak$ (0)	15.51(0.000)***	$loglodz - logkrak$ (0)	17.97(0.000)***
$fpilodz - fpilwów$ (0)	9.06(0.006)***	$loglodz - loglwów$ (1)	4.62(0.153)
$fpikrak - fpilwów$ (7)	13.25(0.001)**	$logkrak - loglwów$ (1)	6.23(0.059)**

Note: The number of lagged differences included in the unit-root regressions is stated in parenthesis. The p-values have been computed by applying a bootstrap procedure (see Berben & van Dijk 1999) and ***, **, ** denote significance at the 1%, 5%, and 10% level respectively.

for most of the city pairs. The results are given in Table 6. Note that the test only allows for a constant but not for linear trends or broken components. This may explain why the test does not reject the unit root null hypothesis with respect to $fpiwar - fpilwów$ and $loglodz - loglwów$ for which the restriction tests suggest to include a broken constant or a linear trend into the price relationship. However, this result can also indicate that for these pairs threshold nonlinearity does not describe the adjustment dynamics properly since the Johansen test finds cointegration for these pairs and the BVD test can reject nonstationarity for the other pairs in favour of a stationary threshold alternative. In any case, the BVD test confirms the outcomes of the generalized Johansen procedure in the sense that almost all price differences are stationary at a 5% or 1% significance level. Furthermore, the *fpi*- and *log*-series produce similar results in general although the significance levels may differ for some pairs. Again, we do not find important differences between the across-border and within-border pairs. On the one hand the two city pairs for which we cannot reject nonstationarity are across border pairs. On the other hand, for the within-border pair $fpiwar - fpilodz$ nonstationary is rejected at a 10% only.

5.3 Results of Threshold Nonlinearity Tests

Since the cointegration analysis clearly supports stationarity of the price differences we proceed to test for threshold effects by applying specific threshold nonlinearity tests. We will relate their results to the outcome of the BVD test which has already suggested threshold nonlinearity according to its alternative hypothesis.

As described in Section 4 we apply four procedures to test for nonlinearity. The multivariate and univariate Tsay tests (Tsay-M, Tsay-U) and the univariate procedures by Hansen (1997, 1999) which test linearity against a two- and three-regime TAR model respectively (Hansen-U12, Hansen-U13). We have to note that none of the procedures allows for broken deterministic components or a linear trend. The omission of a trend may not be problematic for the univariate procedures which refer to the cointegrating residuals only. The cointegration analysis has shown that the trend can be excluded from the price-relationship for most of the city pairs. Nevertheless, one has to be careful in interpreting the tests' outcomes. We do not adjust the cointegrating residual for broken deterministic terms or a linear trend because this may make the tests invalid.[10]

The results in Table 7 show that threshold nonlinearity does not describe the dynamics of the fpi-series in general. Depending on the test, between four and seven pairs exhibit threshold effects. The Tsay tests report the strongest evidence, possibly because they are nonparametric(see Lo & Zivot 2001 for a similar finding). Therefore, they may be less sensitive to model misspecification which could emerge e.g. from omitted deterministic terms. Obviously, the tests differ with respect to specific city pairs. For some pairs nonlinearity is only found by the univariate tests and for other pairs only by the Tsay tests. If the univariate tests suggest threshold effects regarding the cointegrating residual one can expect that these effects are also detected by the multivariate Tsay procedure. However, this is not the case. In contrast, when the multivariate procedure indicates nonlinearity the univariate

[10] In general, the test by Hansen & Seo (2001) which takes account of a linear trend does not give different results than the other procedures.

Table 7. Results of Threshold Nonlinearity Tests for fpi-series

City pair	Tsay-M $\chi^2(4k)$ p-value (k)	Tsay-U $F(2,141)$ p-value	Hansen-U12 Bootstrap p-value	Hansen-U13 Bootstrap p-value
$fpiwar - fpiwil$	0.007(1)***	0.262	0.019*	0.042**
$fpiwar - fpipos$	0.192(3)	0.427	0.257	0.138
$fpiwar - fpilodz$	0.808(1)	0.498	0.104	0.180
$fpiwar - fpikrak$	0.225(2)	0.785	0.530	0.294
$fpiwar - fpilwów$	0.200(1)	0.030**	0.240	0.422
$fpiwil - fpipos$	0.326(2)	0.018**	0.023**	0.032**
$fpiwil - fpilodz$	0.377(2)	0.009***	0.014***	0.010***
$fpiwil - fpikrak$	0.081(2)*	0.003***	0.171	0.149
$fpiwil - fpilwów$	0.915(2)	0.133	0.519	0.344
$fpipos - fpilodz$	0.123(3)	0.224	0.134	0.312
$fpipos - fpikrak$	0.036(3)**	0.838	0.086*	0.037**
$fpipos - fpilwów$	0.002(3)***	0.373	0.314	0.217
$fpilodz - fpikrak$	0.031(1)**	0.998	0.950	0.990
$fpilodz - fpilwów$	0.028(1)**	0.069*	0.340	0.404
$fpikrak - fpilwów$	0.081(1)*	0.041**	0.006***	0.084*

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 vector autoregressions for Tsay-M is stated in parentheses behind the p-value. ***, **,** denote significance at the 1%, 5%, and 10% level respectively.

tests may not reject linearity if the short-run dynamics is affected by nonlinearity but not the cointegrating residual.

The pair $fpikrak - fpilwów$ is the only one for which all tests find threshold effects. Considering exclusively the univariate tests there seems to be robust threshold evidence for $fpiwil - fpipos$ and $fpiwil - fpilodz$. Contrary, with respect to $fpiwar - fpipos$, $fpiwar - fpilodz$, $fpiwar - fpikrak$, $fpiwil - fpilwow$, and $fpipos - fpilodz$ no test detects nonlinearity, although we can observe some borderline cases close the 10% level.

With respect to the *log*-series between two and five city pairs exhibit threshold effects according to the tests. The reason for these lower numbers is that nonlinearity is not found

Table 8. Results of Threshold Nonlinearity Tests for *log*-series

City pair	Tsay-M $\chi^2(4k)$ p-value (k)	Tsay-U $F(2,141)$ p-value	Hansen-U12 Bootstrap p-value	Hansen-U13 Bootstrap p-value
$logwar - logwil$	0.007(1)***	0.707	0.112	0.224
$logwar - logpos$	0.008(2)***	0.404	0.193	0.412
$logwar - loglodz$	0.556(1)	0.203	0.138	0.190
$logwar - logkrak$	0.159(1)	0.984	0.989	0.712
$logwar - loglwów$	0.161(1)	0.069*	0.081*	0.028**
$logwil - logpos$	0.000(1)***	0.001***	0.075*	0.157
$logwil - loglodz$	0.012(1)**	0.000***	0.329	0.579
$logwil - logkrak$	0.008(1)***	0.004***	0.026**	0.035**
$logwil - loglwów$	0.315(1)	0.008***	0.039**	0.395
$logpos - loglodz$	0.647(2)	0.500	0.067*	0.243
$logpos - logkrak$	0.782(2)	0.688	0.654	0.323
$logpos - loglwów$	0.248(2)	0.348	0.118	0.327
$loglodz - logkrak$	0.109(1)	0.295	0.774	0.275
$loglodz - loglwów$	0.345(1)	0.250	0.586	0.852
$logkrak - loglwów$	0.371(1)	0.768	0.212	0.280

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 vector autoregressions for Tsay-M is stated in parentheses behind the p-value. ***, **,** denote significance at the 1%, 5%, and 10% level respectively.

for the pairs Poznań-Kraków and Kraków-Lwów in contrast to the fpi-series. Thus, robust evidence is only related to some of the pairs including Wilno. Note, that the indication for threshold effects is especially low when using the univariate Hansen test with a three-regime TAR alternative. As discussed in Section 4 we expect such an outcome if the nominal transaction costs are fixed since fixed costs induce a changing threshold band.

The results of the threshold nonlinearity tests are not in line with the findings of the BVD test which has rejected nonstationarity of the price-difference in favour of a stationary two-regime TAR model for both the fpi- and log-series. This could be explained by the already mentioned fact that the BVD test has also power against a stationary linear AR model since this model is nested in the two-regime TAR model considered under the alternative. However, there exists a couple of reasons why the specific nonlinearity test may fail to detect threshold effects. First of all, the procedures have clearly lower small sample power in the simulation study by Lo & Zivot (2001) for DGPs generated from a BAND-TAR model compared to a general TAR model like (4.3). Furthermore, it is not clear how the small sample power of the BVD test on the one side and the nonlinearity tests on the other side is affected by the omission of certain important deterministic terms like broken components. Finally, in contrast to the BVD procedure the univariate nonlinearity tests assume stationarity under the null hypothesis. If the price differences are near unit root processes or exhibit deterministic nonstationarity due to omitted linear trend terms the nonlinearity tests may also be size distorted (compare Lo & Zivot 2001).

Hence, it is not obvious on which procedures to rely. Nevertheless, we have evidence for threshold nonlinearity at least with respect to some of the pairs. Therefore, we proceed with estimating the TAR models. Referring to the fpi-series the Hansen tests do not indicate that a three-regime TAR model is less appropriate than a two-regime model. That is why we follow the economic considerations and estimate the unrestricted and restricted three-regime models (4.3) and (4.2). The estimation results can also give us more insights on the relevance of threshold effects since similar parameter estimates for the different regimes or a small number of observations in the threshold band may doubt to model the dynamics of price adjustments by threshold nonlinearity.

5.4 Estimation of Threshold Cointegration Models

Tables 9 and 10 display the estimation results for the unrestricted TAR(3) model (4.3) and the linear AR(1) model. We use mean-adjusted price-differences since the cointegration analysis has confirmed a cointegrating vector $(1, -1)$ for almost all pairs. In contrast to the threshold nonlinearity tests, we think that it should be admissible to adjust also for other deterministic terms like a trend or broken components in accordance with the outcomes of the general Johansen procedure. But we have not done it yet. Instead of presenting the estimates $\hat{\phi}_i$ we refer to $\hat{\alpha}_i = 1 + \hat{\phi}_i$ $(i = 1, 2, 3)$ which is obtained from an equivalent model like (4.3) for z_t. The coefficients α_i $(i = 1, 2, 3)$ measure the speed of adjustment to price differences so that they are easier to interpret and therefore usually stated. Furthermore, we

Table 9. Estimation of Unrestricted Threshold Model for fpi-series

City pair	Linear model	Thresholds		Lower Regime			Threshold Regime		Upper Regime		
	$\hat{\alpha}_l$ (*s.e.*)	$\hat{\tau}_1$	$\hat{\tau}_2$	$\hat{\alpha}_1$ (*s.e.*)	$\hat{\delta}_1$	obs.	$\hat{\alpha}_2$ (*s.e.*)	obs.	$\hat{\alpha}_3$ (*s.e.*)	$\hat{\delta}_3$	obs.
$fpiwar - fpiwil$	0.785 (0.053)	−4.470	8.329	0.554 (0.119)	−8.090	49	0.613 (0.132)	87	−0.603 (0.433)	9.738	23
$fpiwar - fpipos$	0.713 (0.059)	−8.102	−2.970	0.936 (0.476)	20.026	17	−1.293 (0.654)	27	0.698 (0.083)	0.098	115
$fpiwar - fpilodz$	0.862 (0.038)	2.471	5.936	0.816 (0.076)	−0.845	114	2.546 (0.670)	19	1.252 (0.152)	19.999	26
$fpiwar - fpikrak$	0.791 (0.063)	1.010	3.827	0.694 (0.123)	−1.497	105	−0.946 (0.807)	22	0.704 (0.191)	4.849	32
$fpiwar - fpilwów$	0.863 (0.048)	−8.880	−0.531	0.801 (0.340)	−5.089	27	1.646 (0.243)	42	0.768 (0.113)	2.295	90
$fpiwil - fpipos$	0.777 (0.056)	−12.099	0.638	−1.246 (0.934)	−12.632	17	0.549 (0.181)	77	0.426 (0.123)	7.466	65
$fpiwil - fpilodz$	0.844 (0.046)	3.337	8.428	0.837 (0.116)	−1.802	101	−1.672 (0.968)	19	0.283 (0.193)	12.589	39
$fpiwil - fpikrak$	0.758 (0.062)	−10.728	11.148	−0.944 (0.747)	−13.136	17	0.943 (0.073)	126	0.472 (0.369)	0.747	16
$fpiwil - fpilwów$	0.770 (0.064)	−11.494	9.454	0.067 (0.496)	−13.167	18	0.974 (0.104)	119	1.085 (0.285)	80.186	22

Table 9. cont'd. Estimation of Unrestricted Threshold Model for fpi-series

City pair	Linear model	Thresholds		Lower Regime			Threshold Regime		Upper Regime		
	$\hat{\alpha}_l$ (s.e.)	$\hat{\tau}_1$	$\hat{\tau}_2$	$\hat{\alpha}_1$ (s.e.)	$\hat{\delta}_1$	obs.	$\hat{\alpha}_2$ (s.e.)	obs.	$\hat{\alpha}_3$ (s.e.)	$\hat{\delta}_3$	obs.
$fpipos - fpilodz$	0.685 (0.068)	−6.856	9.063	0.333 (0.567)	−1.072	18	0.833 (0.078)	124	−0.214 (0.549)	9.116	17
$fpipos - fpicrac$	0.598 (0.074)	3.106	6.924	0.641 (0.104)	0.359	121	2.648 (0.797)	19	−0.604 (0.399)	8.156	19
$fpipos - fpilwów$	0.734 (0.055)	1.938	5.317	0.719 (0.092)	−1.602	94	−0.692 (1.261)	29	0.852 (0.160)	−17.719	36
$fpilodz - fpikrak$	0.682 (0.071)	−4.739	2.643	0.490 (0.310)	−4.117	19	0.503 (0.150)	97	0.169 (0.198)	3.107	43
$fpilodz - fpilwów$	0.794 (0.051)	−5.418	0.458	0.292 (0.225)	−7.802	38	−0.003 (0.461)	41	0.727 (0.144)	2.032	80
$fpicrac - fpilwów$	0.720 (0.061)	−3.820	3.765	0.484 (0.228)	−5.480	47	0.646 (0.178)	72	0.520 (0.224)	0.584	40

Note: The estimation refers to model (4.3) with $\alpha_i = 1 + \phi_i$ $(i = 1, 2, 3)$ and $\delta_i = \mu_i/(1 - \alpha_i)$ $(i = 1, 3)$. The linear AR(1) model is estimated with a constant but its estimates are not reported since they are very close to zero due to mean adjustment of the price-differences. The given standard errors are computed according to White (1980) to be robust of unknown heteroscedasticity.

give the estimator $\hat{\delta}_i = \hat{\mu}_i/(1-\hat{\alpha}_i)$ $(i = 1, 3)$ in line with the restrictions on the constants in the BAND-TAR model (4.2).[11] The estimators $\hat{\delta}_1$ and $\hat{\delta}_3$ should be close to the estimated thresholds $\hat{\tau}_1$ and $\hat{\tau}_2$ if the restrictions hold.[12]

Obviously, the point estimates for both the *fpi*- and *log*-series are far away from the implications of the transaction cost view of the LOP. Many threshold bands do not include the zero or the autoregressive coefficients of the middle regime are smaller than one of the two outer regimes' coefficients. These observations mean that price adjustment is stronger closer to the price parity than further away. However, the standard errors are rather high, especially for regimes with a low number of observations.[13] There are no apparent differences in the standard errors between the *fpi*- and *log*-series in contrast to our expectation on the effects of changing thresholds. It seems that there are not enough observations to estimate the six regime parameters and the two thresholds more reliable.[14] Interestingly, for a couple of pairs $\hat{\delta}_1$ and $\hat{\delta}_2$ are close to threshold estimates confirming that adjustment only takes place till the edge of the threshold band. Furthermore, the point estimates do not contradict a three regime TAR model in a technical sense for most of the pairs. The only exceptions are $fpiwar - fpiwil$, $fpiwil - fpilwów$, $fpilodz - fpikrak$, $logwar - loglodz$, $logwil - loglwów$, $loglodz - loglwów$, and $logcrac - loglwów$ for which either $\hat{\alpha}_1$ or $\hat{\alpha}_3$ are close to $\hat{\alpha}_2$ suggesting a two-regime TAR model. For $logwar - loglodz$ one could even think of a linear model. Assessing all the previous comments we think that it is worthwhile to estimate the restricted BAND-TAR model (4.2) although we cannot test for the validity of the respective restrictions by applying reliable procedures.

The restricted estimates for the *fpi*-series in Table 11 show that $fpiwar - fpipos$, $fpiwar - fpilodz$, $fpiwar - fpicrac$, and $fpilodz - fpikrak$ have only the minimum number of observations in the middle regime. Firstly, this could indicate that the restrictions imposed within the three-regime framework are not correct such that the middle regime is estimated as small as possible. Secondly, we may also interpret the outcome as a sign that a two-regime or even a linear model is more appropriate for the mentioned pairs. For all these pairs the estimated AR parameter $\hat{\alpha}$ of the outer regimes is very similar to $\hat{\alpha}_l$ from the linear

[11] We do not state $\hat{\delta}_2$ since this representation is only meaningful for $|\alpha| < 1$. Accordingly, one should be careful in interpreting $\hat{\delta}_i$ if the estimator for α_i $(i = 1, 3)$ is close to or larger than one in absolute terms.

[12] Since $\hat{\delta}_1$ and $\hat{\delta}_3$ are ratios of dependent random variables we do not present their standard errors.

[13] Note that 16 is the minimum number of observations per regime in our case.

[14] Compare Lo & Zivot (2001) for similar findings in their empirical application and simulation study.

Table 10. Estimation of Unrestricted Threshold Model for *log*-series

City pair	Linear model	Thresholds		Lower Regime			Threshold Regime		Upper Regime		
	$\hat{\alpha}_l$ (*s.e.*)	$\hat{\tau}_1$	$\hat{\tau}_2$	$\hat{\alpha}_1$ (*s.e.*)	$\hat{\delta}_1$	obs.	$\hat{\alpha}_2$ (*s.e.*)	obs.	$\hat{\alpha}_3$ (*s.e.*)	$\hat{\delta}_3$	obs.
$logwar - logwil$	0.774 (0.065)	0.066	0.115	0.772 (0.069)	−0.024	108	−1.288 (0.857)	26	−0.035 (0.431)	0.152	25
$logwar - logpos$	0.719 (0.056)	−0.084	0.008	1.513 (0.356)	−0.225	26	1.170 (0.329)	59	0.847 (0.131)	−0.107	74
$logwar - loglodz$	0.816 (0.047)	−0.045	0.095	1.028 (0.169)	−1.018	40	0.995 (0.090)	102	1.449 (0.218)	0.230	17
$logwar - logkrak$	0.760 (0.066)	0.028	0.080	0.744 (0.117)	−0.009	109	2.457 (0.715)	25	0.954 (0.313)	−0.843	25
logwar − loglwów	0.876 (0.042)	0.049	0.112	0.992 (0.058)	1.204	98	−1.260 (0.696)	36	0.445 (0.186)	0.137	25
$logwil - logpos$	0.717 (0.080)	−0.117	0.033	−0.213 (0.686)	−0.137	25	0.668 (0.185)	65	0.512 (0.103)	0.075	69
$logwil - loglodz$	0.803 (0.064)	−0.108	0.135	0.357 (0.352)	−0.168	32	0.721 (0.120)	102	0.212 (0.274)	0.163	25
$logwil - logkrak$	0.720 (0.082)	−0.101	0.113	−0.248 (0.521)	−0.144	29	1.023 (0.099)	107	0.433 (0.330)	0.007	23
logwil − loglwów	0.680 (0.096)	−0.091	0.064	0.135 (0.412)	−0.159	28	0.443 (0.239)	84	0.464 (0.196)	0.055	47

Table 10. cont'd. Estimation of Unrestricted Threshold Model for *log*-series

City pair	Linear model	Thresholds		Lower Regime			Threshold Regime		Upper Regime		
	$\hat{\alpha}_l$ (*s.e.*)	$\hat{\tau}_1$	$\hat{\tau}_2$	$\hat{\alpha}_1$ (*s.e.*)	$\hat{\delta}_1$	obs.	$\hat{\alpha}_2$ (*s.e.*)	obs.	$\hat{\alpha}_3$ (*s.e.*)	$\hat{\delta}_3$	obs.
logpos − loglodz	0.706 (0.058)	−0.055	0.105	0.195 (0.193)	−0.068	50	0.663 (0.128)	91	−1.008 (0.352)	0.124	18
logpos − logcrac	0.627 (0.065)	−0.007	0.073	0.478 (0.105)	−0.026	84	−0.462 (0.365)	50	0.068 (0.507)	0.085	25
logpos − loglwów	0.755 (0.046)	−0.076	0.052	0.465 (0.166)	−0.101	38	0.632 (0.206)	67	0.110 (0.187)	0.089	54
loglodz − logkrak	0.644 (0.073)	−0.014	0.002	0.780 (0.182)	0.046	59	−5.668 (2.607)	21	0.574 (0.136)	0.007	79
loglodz − loglwów	0.843 (0.045)	−0.112	0.011	−0.117 (0.382)	−0.131	25	0.932 (0.203)	64	0.986 (0.126)	−1.220	70
logcrac − loglwów	0.805 (0.049)	−0.052	0.043	0.533 (0.215)	−0.084	49	0.524 (0.204)	59	1.007 (0.261)	4.272	51

Note: The estimation refers to model (4.3) with $\alpha_i = 1 + \phi_i$ $(i = 1, 2, 3)$ and $\delta_i = \mu_i/(1 - \alpha_i)$ $(i = 1, 3)$. The linear AR(1) model is estimated with a constant but its estimates are not reported since they are very close to zero due to mean adjustment of the price-differences. The given standard errors are computed according to White (1980) to be robust of unknown heteroscedasticity.

Table 11. Estimation of BAND-Threshold Model for fpi-series

City pair	Linear model $\hat{\alpha}_l$ $(s.e.)$	BAND-Threshold model $\hat{\tau}$	$\hat{\alpha}$ $(s.e.)$	Observations per Regime Lower	Thresh.	Upper
$fpiwar - fpiwil$	0.785 (0.053)	5.732	0.542 (0.104)	37	82	40
$fpiwar - fpipos$	0.713 (0.059)	0.880	0.681 (0.066)	72	16	71
$fpiwar - fpilodz$	0.862 (0.038)	0.741	0.850 (0.042)	84	16	59
$fpiwar - fpikrak$	0.791 (0.063)	0.516	0.777 (0.068)	81	16	62
$fpiwar - fpilwów$	0.863 (0.048)	7.591	0.686 (0.121)	30	102	27
$fpiwil - fpipos$	0.777 (0.056)	8.220	0.303 (0.151)	33	100	26
$fpiwil - fpilodz$	0.844 (0.046)	13.560	0.138 (0.200)	16	123	20
$fpiwil - fpikrak$	0.758 (0.062)	8.491	0.150 (0.167)	26	102	31
$fpiwil - fpilwów$	0.770 (0.064)	8.444	0.472 (0.145)	27	107	25
$fpipos - fpilodz$	0.685 (0.068)	4.888	0.278 (0.157)	39	81	39
$fpipos - fpicrac$	0.598 (0.074)	0.652	0.559 (0.081)	72	26	61
$fpipos - fpilwów$	0.734 (0.055)	3.182	0.614 (0.077)	54	50	55
$fpilodz - fpikrak$	0.682 (0.071)	0.655	0.642 (0.083)	62	16	81
$fpilodz - fpilwów$	0.794 (0.051)	8.225	0.227 (0.195)	22	121	16
$fpicrac - fpilwów$	0.720 (0.061)	3.148	0.523 (0.096)	48	64	47

Note: The estimation refers to model (4.2) with $\alpha = 1 + \phi$. The given standard errors are computed according to White (1980) to be robust of unknown heteroscedasticity.

model. Hence, taking out the observations from the middle regime does not have an important effect when estimating α. Thus, the true AR parameter for the middle regime might not be very different from the ones of the outer regimes and the linear model. For these two reasons, we favour to interpret the results with respect to these city pairs as an indication of an inappropriate model specification instead of having obtained a reliable estimate of a small threshold or transaction cost band. This view is also supported by the fact that none of the nonlinearity tests suggests threshold effects for $fpiwar - fpipos$, $fpiwar - fpilodz$, and $fpiwar - fpicrac$. Regarding $fpilodz - fpikrak$ only the multivariate Tsay test implies threshold nonlinearity but all univariate test do not reject linearity with p-values higher than 0.90. Additionally, the unrestricted estimates doubt a three-regime setup.

The other pairs show estimation results in line with our economic considerations. First, the threshold regimes contain a high number of observations as in other empirical studies on threshold models for single commodities' prices demonstrating the relevance of the transactions cost band (compare e.g. Goodwin, Grennes & Craig 2002 and Goodwin & Piggott 2001). Since the fpi-series are transformed in a nonlinear way it is difficult to give an estimation of the sizes of the bands in terms of the original prices. But regarding the scale of the fpi-series the sizes of the estimated bands are between 0.4% and 9.0% of the average value of all series ignoring the pairs with the minimum number of observations in the middle regime. Hence, the transaction cost bands may be small for some pairs relative to the wheat flour prices and therefore difficult to detect; but they are important with respect to the observed price differences for two market places which have an average value of 7.54 when measured in fpi-series units. Furthermore, we obtain similar figures for the log-series.

Second, we observe that the estimated AR parameters $\hat{\alpha}$ in the outer regimes are always lower than $\hat{\alpha}_l$ from the linear model as it is usually expected. The AR parameters in the outer regimes are supposed to be smaller since the observations in the threshold-band, which follow probably a higher AR coefficient, are ignored for estimation. In other words, one focusses only on the observations for which adjustment takes place. The smaller estimates from the threshold models are often more in line with the economic expectation on the speed of adjustment processes. A smaller α means that the adjustment in the prices due to disequilibria is faster. The speed of adjustment is usually measured by the so-called half-life $\ln(0.5)/\ln(\alpha)$ which states the number of periods required to reduce one-half of a deviation from the price-parity.[15] However, the reduction in the value of $\hat{\alpha}$ is rather different for the single pairs. In the linear model $\hat{\alpha}_l$ is between 0.598 and 0.863 and reduces to 0.138-0.686 when threshold effects are allowed for. Thus, the half-lives reduce from 1.3-4.7 months to 0.3-1.8 months so that adjustment induced by deviations from the stationary price

[15] For a thoroughly discussion on the economic interpretation of and the relationship between estimates of the AR coefficient in linear and threshold models see Obstfeld & Taylor (1997) and Taylor (2001).

Table 12. Estimation of BAND-Threshold Model for *log*-series

City pair	Linear model $\hat{\alpha}_l$ (*s.e.*)	BAND-Threshold model $\hat{\tau}$	$\hat{\alpha}$ (*s.e.*)	Observations per Regime Lower	Thresh.	Upper
$logwar - logwil$	0.774 (0.065)	0.101	0.482 (0.158)	32	97	30
$logwar - logpos$	0.719 (0.056)	0.006	0.706 (0.060)	66	16	77
$logwar - loglodz$	0.816 (0.047)	0.007	0.804 (0.051)	84	17	58
$logwar - logkrak$	0.760 (0.066)	0.025	0.691 (0.089)	75	33	51
$logwar - loglwów$	0.876 (0.042)	0.140	0.519 (0.124)	26	116	17
$logwil - logpos$	0.717 (0.080)	0.079	0.413 (0.165)	43	72	44
$logwil - loglodz$	0.803 (0.064)	0.165	0.307 (0.248)	20	123	16
$logwil - logkrak$	0.720 (0.082)	0.087	0.393 (0.178)	33	88	38
$logwil - loglwów$	0.680 (0.096)	0.103	0.375 (0.222)	26	108	25
$logpos - loglodz$	0.706 (0.058)	0.068	0.256 (0.126)	40	83	36
$logpos - logcrac$	0.627 (0.065)	0.020	0.522 (0.078)	66	32	61
$logpos - loglwów$	0.755 (0.046)	0.069	0.509 (0.075)	40	74	45
$loglodz - logkrak$	0.644 (0.073)	0.005	0.620 (0.081)	69	16	74
$loglodz - loglwów$	0.843 (0.045)	0.016	0.827 (0.052)	76	16	67
$logcrac - loglwów$	0.805 (0.049)	0.011	0.790 (0.054)	68	16	75

Note: The estimation refers to model (4.2) with $\alpha = 1 + \phi$. The given standard errors are computed according to White (1980) to be robust of unknown heteroscedasticity.

relationship is much faster in the threshold models. Again, we have similar outcomes for the *log*-series. Hence, we can conclude that threshold nonlinearity and, thus, transaction costs are important in understanding the dynamics of price adjustments for most of the city pairs.

Finally, we want to comment on the effects of the old borders and of the changing aggregate price level in the presence of nominal fixed transaction costs.

As in case of the cointegration analysis we do not observe systematic differences between within-border and across-border city pairs. The estimates of α with respect to the within-pairs $fpiwar - fpiwil$, $fpiwar - fpilodz$, $fpiwil - fpilodz$, and $fpilodz - fpilwów$ range from 0.138 to 0.850 and cover the whole spread of all $\hat{\alpha}$'s. Moreover, we have regressed the sizes of the threshold bands on the geographical distance between the city pairs and on a dummy variable which is set to one for within-border pairs and is zero otherwise. The estimated coefficient for the dummy variable is not significant neither when including nor when ignoring the pairs with a minimum number of observations in the threshold regime. Hence, the former partition borders have no significant effect on the sizes of the bands. However, distance has a significant impact on the bands' size if all pairs are considered in the regression. The same conclusions can be derived for the *log*-series.

The *fpi*- and *log*-series do not produce important differences as in the cointegration analysis. The estimates for α and the corresponding standard errors are not different in general. The latter is in contrast to our expectation of higher standard errors when the *log*-series are used since they imply an unstable threshold band in case of a changing aggregate price level and nominal fixed transaction costs. However, we find different estimation results for some of the city pairs. On the one hand, this may indicate that the relevance of the price level's impact differs among the pairs. On the other hand these difference may be simply due to technical effects of the nonlinear data transformation regarding the *fpi*-series. These effects can e.g. result from the involvement of deterministic components we have not taken care of yet since we have only used mean-adjusted price differences.

6 Summary and Concluding Remarks

In this paper we have studied the topic of economic integration in interwar Poland by performing a threshold cointegration analysis of the LOP for the Polish wheat market between 1924 and 1937. We consider the transaction cost view of the LOP which says that the prices of the same commodity at two spatially separated market places should converge whenever the price difference exceeds the transaction costs. Our analysis is based on monthly retail prices for wheat flour of six of the biggest Polish cities from which we can construct 15 city pairs for a price comparison.

Within our study we have examined two different issues that are probably relevant for economic integration in interwar Poland. The first issue refers to the fact that Poland comprised three rather different areas belonging to Russia, Germany, and the former Habsburg monarchy after its reunification in 1919. Hence, we ask whether border effects exist within one country. Secondly, we have aimed to analyze the importance of the strongly changing aggregate price level. Since the nominal railway tariffs have been fixed in the sample period we expect that the real transaction costs change with the price level meaning that inflation fosters and deflation hinders economic integration. To capture these effects we suggest to multiply the log-prices with the aggregate price level. The results for the adjusted series should differ from the ones for the log-series since the use of the latter ones implies a model with a changing threshold band.

Our main findings are the following. The price differences form a stationary relationship for almost all city pairs implying that the prices adjust to deviations from the price parity. However, the prices do not converge completely since deterministic terms like constants, linear trends, or broken components enter the price relationships. This suggests imperfections in integration either in the wheat or other markets of the Polish economy. Although the results on threshold nonlinearity are not clear cut regarding all city pairs we observe that threshold effects are important for the dynamics of the price adjustments in the sense that there exist relevant transaction costs for arbitrage processes. Hence, we have found evidence for a relative version of the LOP including transaction costs. That is, the interwar economy can be regarded as integrated but with obvious restrictions.

Furthermore, we can conclude that the old borders did not have an effect on economic integration in Poland. The results of the within-border and across-border city pairs do not differ systematically in the sample period under study. We interpret this outcome as a sign for a successful integration policy in the first years after the reunification in 1919. However, differing results with respect to certain pairs may indicate city specific effects which are interesting to study in more detail.

Regarding the impact of the aggregate price level we can not discriminate between the use of the adjusted series and the log-price series. The results for both sets of series are generally the same. Therefore we suppose that only a certain part of the transaction costs are nominal fixed but not all costs. This implies a changing real transaction cost bands for both sets of series so that the results should not differ importantly. We do not think a sample splitting at the end of the inflationary period produces more helpful outcomes for two reasons. Firstly, the price level still changes within the separated samples which induces again nonstable threshold bands. Therefore, we have the same situation concerning the adjusted and non-adjusted series. Secondly, the price level at the end of the deflationary period is the same as at the beginning of the inflationary period. Hence, only the direction of economic integration due to the evolvement of the price level differs but not the average degree of integration. Thus, both periods are supposed to generate similar findings. To understand the effect of the aggregate price level more precisely this issue should be addressed in a cross-section analysis as in Engel & Rogers (1996). A cross-section study using the price level as a regressor in order to explain price variations is less dependent on the unknown proportion of nominal fixed transaction costs. In fact, it is the unknown proportion that is crucial for our time series approach.

Finally, we expect that a more detailed consideration of deterministic terms when testing for nonlinearity and estimating threshold models may resolve some contradicting and differing results with respect to the threshold effects. An estimation of threshold vector error correction models can give further insights if asymmetric adjustments in the prices are present. Such asymmetries are also relevant regarding city specific effects. All these issues are left for further research.

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