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Economic Integration Across Borders: The Polish Interwar Economy 1921-1937

by

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Abstract

In this paper we study the issue of economic integration across borders for the case of Poland's reunification after the First World War. Using a pooled regression approach and a threshold cointegration framework we find that the Polish interwar economy can be regarded as integrated with some restrictions. Moreover, a significant negative impact of the former partition borders on the level of integration that can be found for the early 1920s vanishes in the middle of the 1920s. This suggests that the integration policy after the reunification of Poland in 1919 was surprisingly successful.

Keywords: Economic integration, Border effects, Law of one price, Poland, Threshold cointegration

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

1 Introduction

What happens to an economy if we change its borders? Since administrative and political borders use to come along with massive barriers to trade, information, and mobility, the removal of such borders should lead to better economic integration. However, the literature on integration due to changing borders remains inconclusive on these issues. Some authors find borders to remain massive barriers to integration even after their removal (e.g. Engel & Rogers 1996), others measure a high degree of integration across borders already before their removal (e.g. Moodley, Kerr & Gordon 2000).

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

Hence, we are in a position to evaluate the process of integration across the former partition borders of Poland. To this end, we draw on two strands in the literature and show that they can complement each other, namely the measurement of border effects in the wake of Engel & Rogers (1996) and the use of threshold cointegration analysis following Balke & Fomby (1997) and Lo & Zivot (2001). Both approaches can be motivated with the law of one price (LOP), adjusted for costs of transport and communication which should hold for a panel of price data in an economically integrated area. But they differ with respect to what dimension of the panel they focus on. Consider the approach of Engel & Rogers (1996). Since arbitrage is expected to keep price differentials between locations within the limits of transport and communication costs, the volatility of these price differentials can be a proxy for the degree of integration between locations. Engel & Rogers (1996) proposed then to regress the cross-section of bilateral price volatilities on distance, location specific effects, and a border dummy variable. Obviously, if one estimates a positively significant coefficient on a border dummy the border matters, insofar as it reduces the degree of integration between locations divided by that border. The idea is appealing for large cross-sections and within our regression specification we can deal with structural instability over time due to its cross-sectional nature. However, it does not fully exploit the evidence on prices. A high volatility of price differentials can be caused by the absence of price adjustment between the locations but also by higher costs of transport and communication. Thus, on the intertemporal dimension the approach cannot distinguish between a situation when price differentials are volatile because prices do not adjust at all or because prices respond with a low speed or adjust only to some high level of transaction costs. Threshold cointegration models as discussed in Balke & Fomby (1997) and Lo & Zivot (2001) allow to make these distinctions and to analyse the dynamics of bilateral price differentials in great detail. This approach, however, gets very labour-intensive for large cross-sections and shares the usual problems of time series analyses with respect to structural instability. Therefore, in order to exploit fully the information contained in our data we will consider the two model frameworks as complementary to each other and apply both to the case of Interwar Poland.

The rest of the paper is organised as follows. In the next section we present the historical background of the study and describe our data set. Section 3 introduces the two econometric model frameworks and discusses their relationship. The empirical results are presented in Section 4. The last section summarises our findings and concludes.

2 Historical Background and Data

Between 1772 and 1795 the noblemen's republic of Poland (Rzeczpospolita Polska) was divided into three parts between the empires of tsarist Russia, the Habsburg monarchy, and the emergent Prussia. As a consequence of the partitions - "the first very great breach in the modern political system of Europe"(Edmund Burke) - Poland disappeared from the map. Only the specific constellation at the end of the First World War, when all three partition powers were severely weakened through war and revolution, opened the way for its restoration. Owing to the long period of partition there were different legislations about virtually all aspects of social, political and economic life. The government, actually, could rely on extensive programs for legal, administrative, and economic unification that had been prepared since 1907 for a future Polish state. However, the agenda was not set by any political or economic "master plan" but rather by the ongoing war that Polish troops fought with the Soviet army in the east (see Landau 1992, Roszkowski 1992).

This war required massive outlays and some mechanism to finance them. Since international credit was not available - the Paris peace conference did not start before January 1919 and Poland was yet to be formally recognized as a state - the government had to choose between the expropriation ("nationalization") of domestic private capital and ways to tax it (Landau & Tomaszewski 1999). The political compromise in 1919 relied on early concessions to the socialists on the one hand and observing private property rights on the other hand. As a consequence, the next step was to create the institutional framework necessary to tax capital and labour: a common currency and a working fiscal administration. The unification of the fiscal administration belonged to the very first institutional changes. While for most of the former Austrian and Russian parts this was 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). A common income tax was decreed in July 1920, but it took several years to implement it on the former Russian territories. Business taxes in turn were introduced and unified on the whole territory until July 1925, following the Russian system of business certifications. Nevertheless, some differences of the tax system - e.g. the real estate tax - remained persistent until 1936 (Weinfeld 1935).

Next, the precondition for any tax system to work was the creation of a common currency area, namely the unification of 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 the Polish Mark it adopted a stepwise strategy to get rid of the competing banknotes (Landau 1992). Some months after the introduction of the Polish Mark as a parallel currency in the different areas, the other currencies were withdrawn. For most of the Polish territory this was realized already in April 1920 with the exception of Upper Silesia (Nov. 1923) (Zbijewski 1931). While such a quick institutional change was an indisputable success, it could not create the necessary revenues to win a war. But it opened the way for the Polish government to effectively tax money holders by inflation. As estimated by Zdziechowski (1925) the money supply increased between 1918 and 1919 by 519%, in the following year by another 929% to reach in 1923 more than 12,000,000% of the level in 1918. Obviously, the temporal gains from seigniorage and the devaluation of the budget deficit were quickly wiped out by the costs of hyperinflation, namely the loss of access to foreign capital. When Prime Minister Władysław Grabski tried to stabilize the currency in 1924, his 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 with the 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 Zloty was fixed par with the Swiss gold franc. After a second wave of devaluations the exchange rate finally stabilized at a sustainable level around May 1926. From now on the government started to defend the parity at any cost.

The war in the east also induced a quick improvement of the transportation system, since it required a network to transport men and material. After rather spontaneous takeovers of the railway networks in the different areas during the last months of the First World War, a railway ministry started its work already in October 1918 and developed a ten-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 centres of the former partition areas connected (Hummel 1939). The speed of the network and its capacity to transport goods was not only a function of the existence of railway connections themselves but also crucially depended on the material used. Table 1 gives an overview for the development of important newly built railway lines and the changes in speed. Since nearly all freight transport over 50km took place on railways with normal gauge (97.6% in 1925 and 98.7% in 1938, see Brzosko 1982, p. 358) this development of the railway network can be expected to have had a strong integrating impact on the economy.

Hence, the most obvious non-tariff barriers to trade and mobility within the new Polish state such as different currencies, different tax systems, and a shortage of transport facilities

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

Date of opening	Connection	Distance	Av. Length of the trip as in 1937
1848	Warsaw-Kraków via Częstochowa	ca.364 km	8.00 hrs
Nov. 1934	Warsaw-Kraków via Radom	ca.320 km	5.20 hrs
1872	Warsaw-Poznań via Torun	ca.376 km	7.00 hrs
Nov. 1921	Warsaw-Poznań via Wrzesnia	ca.304 km	4.45 hrs
1857	Poznań-Kraków via Wroclaw	ca.380 km	n.a.
Nov. 1926	Poznań-Kraków via Wieluń	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); Olszewicz (1938, p. 223).

were considerably reduced if not completely removed until 1926. Moreover, as one of the first steps to unify the new economy a common external tariff was introduced in November 1919. But it took some more time to get rid of internal tariffs and a system of widespread regulations of commodity and factor markets. In part this system was again motivated by the need to furnish 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. For example, there remained a customs frontier between the former Prussian partition area and the rest. This kept grain prices in the Prussian area at an artificially low level, thereby providing cheap supply for the fighting troops (Kozlowski 1989, Landau & Tomaszewski 1999). 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 was removed in mid-1921 and until the end of 1921 most other regulations on the commodity markets had disappeared (Tomaszewski 1966).

To sum up, Poland after 1919 was characterized by a multitude of barriers to trade, information, and mobility which may give rise to border effects in line with the arguments of Engel & Rogers (1996). Between 1919 and 1926 Poland experienced massive efforts to remove these barriers that divided the country for more than one century. But to what extent did this process of institutional change result in better economic integration across the former borders? To analyse this issue we use monthly retail prices from several publications of the Polish Statistical Office (GUS) in Warsaw covering the period from January 1921 to December 1937.[1] All prices were reported to the GUS by the city administrations, for 1921-1925 as monthly averages, for 1926-1937 as prices of the last week in a month. We have evidence over the complete period for the cities of Warsaw, Kraków, Łódź, Lwów, Poznań, and for Wilno from 1924 onwards. This allows us to distinguish between the formerly Russian area (Warsaw, Łódź, Wilno), the formerly Austrian area (Kraków and Lwów) and the formerly German area (Poznań). Due to the currency reform in 1924 that followed the period of hyperinflation, the GUS published all price series until June 1923 in Polish Mark and beginning in January 1924 in the new currency, the Zloty. Therefore, we split our sample at June 1923 and get a first subsample including 30 observations (January 1921-June 1923) for a total of 10 city pairs and a second subsample including 160 observations (January 1924-April 1937) for 15 city pairs.

The GUS provided price series for several basic commodities including coal, soap, vegetables, and four kinds of grain. Some of these markets, however, were at least temporally subject to high levels of concentration with prices set by interregional cartels rather than by competitive arbitrage traders. Therefore, we chose to focus on the market of wheat flour (milled at the same grade) in different cities where the historical records do not show any major market concentration. As most of the wheat grinding has been done in small mills that were evenly spread around the whole country we can speak of a dense, decentralized network of mills from which the flour was shipped to the cities. Hence, we can assume that the prices for wheat flour were the outcome of a competitive market with arbitrage adjusting for large price differentials between different locations.

The log-price series are shown in Figures 1 and 2. Figure 2 refers to the period January 1924-December 1937 which is used for the threshold cointegration analysis. In line with the general price development in Poland the wheat flour prices increase until 1927 followed by a short period of stabilization. Then, in line with the great depression, the prices fell dramatically starting from 1929 onwards. This change from an inflationary to a deflationary environment may lead to possible breaks in the deterministic components of the time series.

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

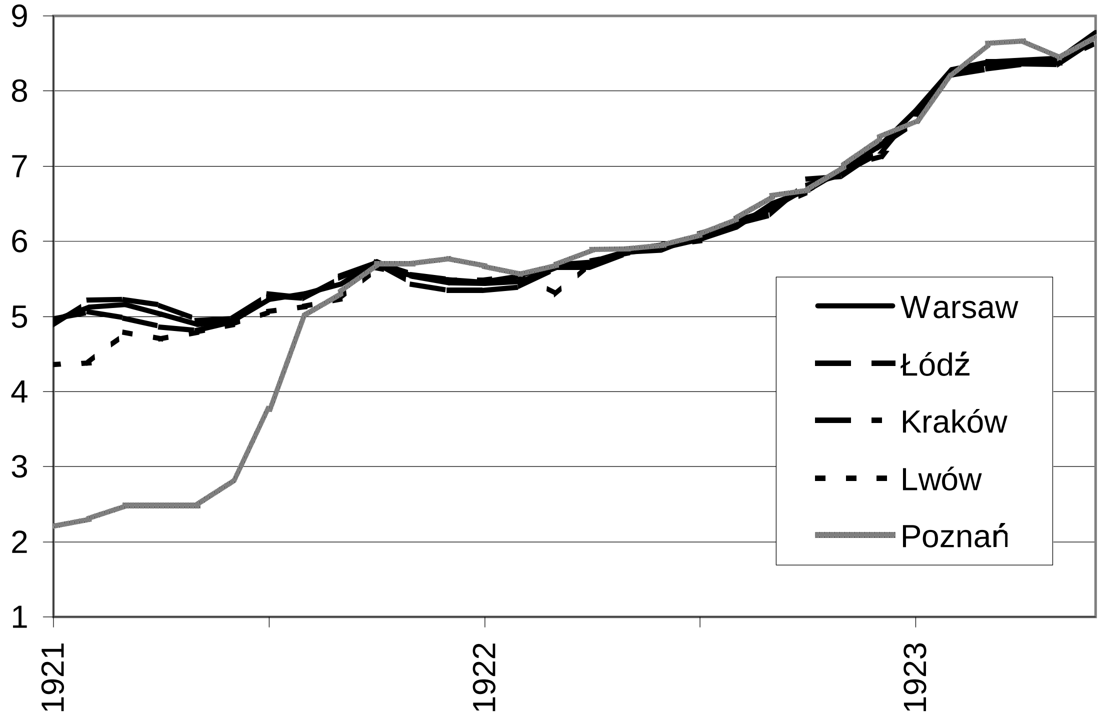

Figure 1. Logarithm of wheat flour prices 1921-1923

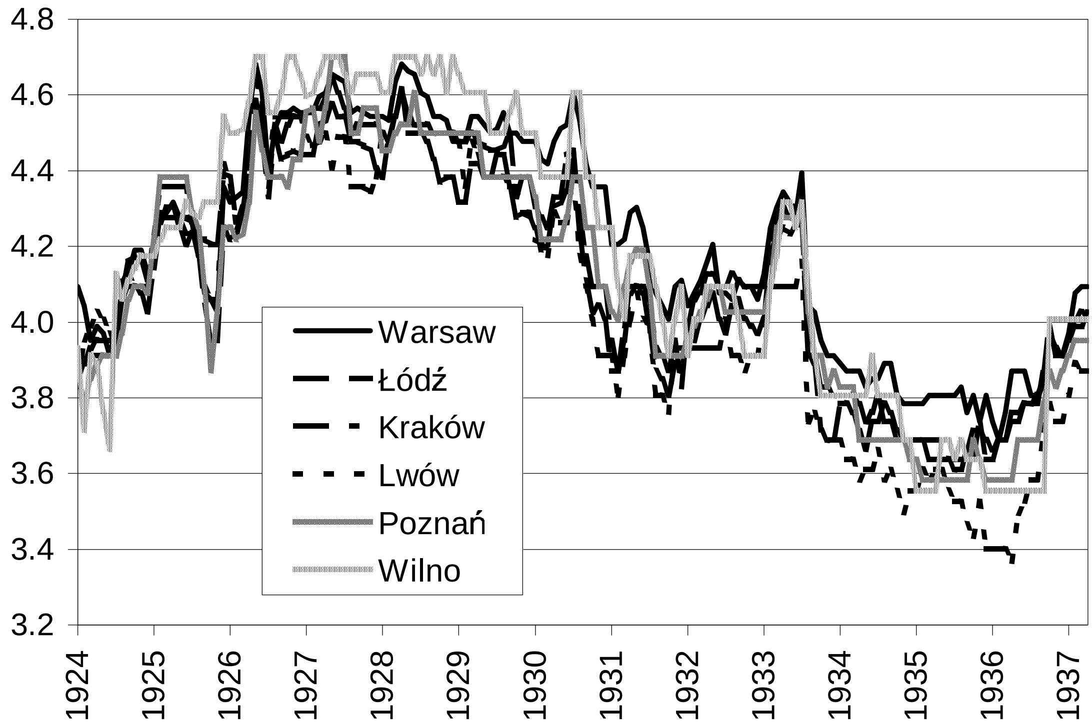

Figure 2. Logarithm of wheat flour prices 1924-1937

We have to address this issue within the threshold analysis. To determine a common break point we refer to the food price index (FPI) since break dates may be different for the price series. The FPI is the most complete price index available for interwar Poland on a monthly basis. It consists of 16 agricultural goods and wheat account to slightly less than 5% of the index. As break date we choose the observation May 1929 since from this month onward the FPI started to fall. A Chow breakpoint test confirmed this break date. Accordingly, we have 64 and 96 observations for the two subperiods respectively.

For the analysis of possible border effects according to the Engel-Rogers approach we can also take account of the data for January 1921-June 1923. As will be explained in the next section, our regression specification allows to handle periods of missing data and structural differences within and between the subperiods. Figure 1 shows the strong price increase due to the high inflation at that time. Moreover, the lower flour prices in Poznań in 1921 are apparent which reflect the regulated prices in the former Prussian area.

3 Econometric Model Frameworks and Methods

Both the framework of Engel & Rogers (1996) and the threshold cointegration analysis can be associated with the LOP. 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 transaction costs connected with arbitrage. Only if the price deviations between the two market places exceed the transaction costs, arbitrage is profitable and takes place. In fact, by referring to the LOP we measure market integration by price convergence. In contrast, McCallum (1995) links integration to trade and Moosa & Bhatti (1995) to investment flows. We prefer the LOP approach since lasting price differences can indicate integration failures even in the presence of trade or investments. In this sense, the occurrence of price convergence offers a kind of absolute measure for integration in contrast to relative measures related to flow data.

In the following we first comment on the approach by Engel & Rogers (1996) and afterwards, we describe the main steps of the threshold cointegration analysis and discuss the relationship between both model frameworks.

3.1 Pooled Regression Approach

Let P_{it} and P_{jt} be the prices of a good in the cities i and j respectively at time t and p_{it} and p_{jt} be the corresponding logs of the prices. Engel & Rogers (1996) suggest to compute the standard deviation of $\Delta z_t = \Delta(p_{it} - p_{jt})$ over the whole sample for each city pair. These standard deviations are considered to be a measure for the volatility of the price differences. Then, they regress the standard deviations on the log of the distance between the cities, city specific dummy variables and a border dummy variable indicating whether the border is crossed or not. Obviously, the border dummy variable summarizes all possible effects and costs associated with crossing a border. Thus, border effects are present if the border dummy coefficient is positively significant what means higher variation in the price differences.

The description of the development of the Polish interwar economy suggests that the importance of the former partition borders may have changed over time. Therefore, we pick up the idea of Engel & Rogers (1996) to split the sample in order to allow for time specific border effects. To be precise, we propose to compute the standard deviation of Δz_t for the first and second half-year of each year. Let $V(z_{t,s})^{(i,j)}$ denote the standard deviation of Δz_t for the city pair (i, j) in the half-year s $(s = 1, 2)$ of year t $(t = t_1, \ldots, T)$. Then, we regress $V(z_{t,s})^{(i,j)}$ on yearly border dummy variables $B_t^{(i,j)}$ $(t = t_1, \ldots, T)$ instead of a single border dummy and obtain the regression equation

$$V(z_{t,s})^{(i,j)} = \beta_1 d^{(i,j)} + \sum_{t=t_1}^{T} \beta_{2t} B_t^{(i,j)} + \sum_{m=1}^{6} \beta_{3m} C_m + u_{t,s}^{(i,j)} \quad \forall\, i, j \text{ with } i \neq j \text{ and } s = 1, 2, \quad (3.1)$$

where $d^{i,j}$ is the log of the distance between the cities i and j and the C_m's are city specific dummy variables which take on a value of one for the cities i and j if $m = i$ or $m = j$. The yearly border dummy variables $B_t^{(i,j)}$ are equal to one for the respective year t if the border is crossed and zero otherwise. We can trace possible changes in the importance of border effects through the corresponding yearly border coefficients β_{2t}. If $\beta_{2t} > 0$, crossing the border increases price variation. Regarding the error terms $u_{t,s}^{(i,j)}$ we allow for heteroscedasticity with respect to the different years t, hence $u_{t,s}^{(i,j)} \sim (0, \sigma_t^2)$. Accordingly, the unknown parameters β_1, β_{2t} $(t = t_1, \ldots, T), \beta_{3m}$ $(m = 1, \ldots, 6)$ are estimated by a feasible GLS procedure where we first estimate (3.1) by OLS in order to obtain estimates for the yearly error variances. A SUR estimation is not possible since we would have to determine too many unknown covariances. This problem occurs because the number of years is higher than the number of city pairs (compare e.g. Griffiths, Skeels & Chotikapanich 2002).

The specification (3.1) enables us to take also the period January 1921-June 1923 into account since the computation of standard deviations only for subperiods does not require a continuous data set as a time series approach does. Since the data for January to April 1937 do not cover a full half-year we exclude these observations and obtain 1921-1936 as our sample, i.e. $t_1 = 1921$ and $T = 1936$. We have chosen the half-year subperiod specification in order to use more observations for the estimation of the β_{2t} coefficients. Furthermore, the consideration of smaller subperiods makes it is more likely to have homogenous data for computing the price volatility. In addition, by allowing for time specific parameter coefficients and heteroscedastic error terms we can deal with structural heterogeneity over time. These adjustments are easily possible since we have cross sections over which we can calculate volatilities and estimate parameters at several points in time. Finally, the effects of linear trends are eliminated because the standard deviations are computed with respect to the first differences of $p_{it} - p_{jt}$. Even if a linear trend does not cancel out when subtracting the prices it vanishes if the first difference is taken. Similarly, the impact of breaks in the deterministic components diminishes owing to differencing.[2]

One may argue that the yearly border dummy variables in (3.1) measure general time specific effects. In the period 1921-1923 only two of the ten city-pairs are within-border pairs. For 1924-1936 we have four within-border pairs out of 15 pairs. Accordingly, the yearly border dummy variables contain a relatively low number of zeros which make them similar to general yearly time dummy variables. Therefore, we have added such yearly dummies to equation (3.1) in a second regression in order to distinguish yearly border and general time effects.

3.2 Threshold Cointegration Framework

The threshold cointegration analysis is directly motivated by the transaction cost view of the LOP. Therefore we formalize it now more precisely. Consider again two market places i and j and let N_{ji} denote some export level of a good from place j to i. Assume further that the transaction costs take the iceberg-form which is used in the recent literature of economic geography. Then, $e^{-\tau} P_{it}$ is the per-unit revenue when the good is sold in location i where

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

$\tau > 0$ is a cost parameter. Hence, $(1-e^{-\tau})P_{it}$ are the transaction costs which "melt away" a portion of the revenue. Intuitively, τ depends positively on the geographical distance between the locations i and j. Moreover, when border effects are present τ also differs depending on whether the locations lie in the same area or not. Finally, arbitrage from i to j is only profitable if $P_{it}N_{ji}e^{-\tau} > P_{jt}N_{ji}$. This results in the condition $\log(P_{it}/P_{jt}) = p_{it} - p_{jt} > \tau$. Hence, arbitrage from j to i takes place when the log-price difference $p_{it} - p_{jt}$ is larger than the cost parameter τ. Equivalently, one trades from location i to j only if $p_{it} - p_{jt} < -\tau$. Thus, we obtain $[-\tau; \tau]$ as a band of no arbitrage. Within this band no trade occurs 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.

This economic framework is econometrically translated into a threshold cointegration model (see Lo & Zivot 2001). Outside the band of no arbitrage adjustment processes ensure that the prices move back toward the price parity equilibrium. Since we use variables in logarithms this suggests that the log-price series are cointegrated with a cointegrating vector (1, -1). In other words, the log-price difference forms a stationary relationship. However, if the price difference is smaller than the transaction costs it behaves like a nonstationary time series because prices do not adjust. Accordingly, adjustment stops at the edges of the band of no arbitrage and does not continue until price parity. These considerations result in the following symmetric three-regime BAND-threshold autoregressive (BAND-TAR) model:

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

where $z_t = p_{it} - p_{jt}$ is the log-price difference at time t and $\eta_t \sim$ i.i.d. $(0, \sigma^2)$. The symmetric threshold band, i.e. the regime $[-\tau,\tau]$, relates to the band of no arbitrage in which z_t evolves like a random walk. Its limits are described by the so-called thresholds which coincide with the transaction cost parameter. Therefore they are also labelled as τ. By contrast, in the outer regimes, for which we have $|z_t| > \tau$, economic forces push the prices together implying $-1 < \alpha < 1$. The autoregressive coefficient α can be easily interpreted in terms of the speed of price adjustment by referring to the so-called half-life $h = \ln(0.5)/\ln(|\alpha|)$. The half-life states the number of periods required to reduce one-half of a deviation from the price-parity. Hence, a smaller value of $|\alpha|$ means that adjustment in the prices due to disequilibria is faster.

The BAND-TAR model assures that prices only adjust to the edges of the threshold band (3.2) by setting the means of the outer regimes to $-(\alpha-1)\tau$ and $(\alpha-1)\tau$ respectively. Moreover, the transaction cost view suggests symmetry regarding the adjustment coefficient α and the threshold τ since arbitrage should be induced in the same way independent of where the prices are higher.

It is possible to consider more general TAR models for z_t or multivariate threshold models with respect to the single time series $p_{i,t}$ and $p_{j,t}$.[3] Lo & Zivot (2001) evaluate the relative performance of multivariate and univariate procedures by means of an extensive Monte Carlo study. Their results, however, 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.

So far we have just considered the simple log-price difference $z_t = p_{i,t} - p_{j,t}$. The presentation of the data in the foregoing section has shown that the series may be characterized by a broken linear trend and different levels corresponding to the succession of inflationary and deflationary periods. The question is whether these or unbroken deterministic terms affect the log-price differences in the sense that they have to be included into the price relationship in order to obtain stationarity. As long as both price series contain the same deterministic terms with the same slope and level they cancel out when subtracting the series. But if the deterministic parts differ one may consider the extended relationship $z_t^* = p_{1,t} - p_{2,t} + \psi d_t$ instead of z_t where the relevant deterministic terms are collected in d_t. The inclusion of deterministic components in z_t^*, however, has important economic implications. If z_t^* contains a linear trend or a broken trend the prices in two market place diverge deterministically. Obviously, this contradicts economic (price) integration even if prices cointegrate, i.e. if z_t^* is a stationary relationship. The situation is different for a constant or broken constant entering z_t^*. In this case the prices still converge but not towards price parity. Instead, prices differ by some fixed (deterministic) amount in equilibrium. We may associate such kind of price convergence with a relative version of the LOP.

Deterministic price differences could be due to different local selling and buying costs

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

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

The threshold cointegration analysis is performed 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.

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 also enables us to test which deterministic components affect the price cointegration relationship in line with the foregoing discussion. Additionally, we can test whether the cointegrating vector can be restricted to $(1, -1)$ so that the log-price difference is in fact the relevant quantity for price adjustment.

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

$$\begin{aligned}&\Delta p_t = \alpha(\beta' p_{t-1} - \theta_1(t-1)D_{1,t} - \theta_2(t-1)D_{2,t}) + \nu_1 D_{1,t} + \nu_2 D_{2,t} + \gamma_2 d_{2,t} + \varepsilon_t,\\ &\qquad \varepsilon_t \sim N(0, \Omega) \text{ and } t = p+1, p+2, \ldots, T,\end{aligned} \tag{3.3}$$

where $p_t = (p_{i,t}, p_{j,t})'$, $D_{1,t}$ 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 and θ_1, θ_2, ν_1, and ν_2 are the corresponding $(n \times 1)$ parameter vectors for the linear trends and constants of the two regimes. In the empirical analysis we set T_1 to May 1929 as explained in Section 2. 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 prices in p_t. Hence, the rank r determines the number of cointegration relations. We consider the trace test version, i.e. the pair of hypotheses is $H_0(r_0): \mathrm{rk}(\Pi) = r_0$ *vs.* $H_1(r_0): \mathrm{rk}(\Pi) > r_0$. We expect a cointegrating rank of one since the LOP implies a cointegrating relationship between the log-prices. Critical or p-values of the test can be computed by using a response surface given in Johansen et al. (2000). Of course, one has to augment (3.3) by lags of Δp_t and $d_{2,t}$ if necessary.

Provided we have found a cointegrating rank of one, we test whether β can be restricted to $(1, -1)$. Similarly, we perform restriction tests on the parameters θ_1, θ_2, ν_1, and ν_2 to examine which deterministic terms are present and enter the cointegrating relationship. If θ_1 and θ_2 are nonzero, trend components enter the price relationship such that price convergence does not occur. All these restriction tests are asymptotically $\chi^2(s)$ distributed where s 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.

A different strategy of analysing deterministic terms with respect to economic integration would be to employ a cointegration test which excludes trend components from the cointegrating relationship. If cointegration is not found under that setup one would conclude that the LOP does not hold. This approach, however, has a practical problem. If we allow for breaks in the deterministic components within the short-run dynamics the limiting distribution of the corresponding cointegration test version depends on unknown quantities (see Johansen et al. 2000). Therefore, we prefer our approach where we first test for cointegration using a general deterministic modelling framework. Then, in a second step, we test whether economic integration is ruled out by deterministic price divergence due to trends in the cointegrating relationship.

To test for threshold nonlinearity several procedures are applied. 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 higher 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 and a three-regime TAR model respectively. The three regime model, however, is much more general than our BAND-TAR model (3.2). The procedures have a sup-F-type form comparing the sum of squared residuals under the null and the alternative hypotheses. Since the threshold parameter is not identified under the null hypothesis of linearity p-values have to be determined by bootstrap methods. Since the univariate procedures assume a known cointegrating relationship between the log-prices we have also applied a 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 in detail.

One may test for threshold cointegration directly instead of following the two-step approach which first applies the linear Johansen cointegration procedure and tests for threshold nonlinearity afterwards. We lack, however, suitable threshold cointegration tests. Available tests assume only a two-regime threshold model (e.g. Hansen & Seo 2001, Enders & Siklos 2001) and they often have rather low small sample power (compare Lo & Zivot 2001).[4] The former problem also applies to threshold unit root tests which could be used in principle if one is willing to assume a known cointegrating vector. Furthermore, threshold cointegration tests may also have power against linear cointegration. Finally, the treatment of deterministic terms, especially broken components, is much easier within a linear framework.

If the tests indicate threshold nonlinearity one proceeds to estimate the threshold models. A reasonable strategy would be to estimate first an unrestricted three-regime TAR model 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. Accordingly, we focus on (3.2) and comment only briefly on the results for the unrestricted models in the next section.

We estimate the three regime BAND-TAR model (3.2) via sequential conditional least

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

squares methods and apply a grid search to locate the threshold. The step length is set to 10^{-5} which is approximately 10^{-4} times the average log-price difference of all city pairs in the whole sample. Following the literature, we set the minimum number of observations per regime to 15% of the total number of observations. Hence, we obtain a minimum of 24 observations since our time series comprise 160 data points. Note, that the minimum number of observations applies to the sum of the observations in the outer regimes. The reader is referred to Hansen (1999) and Lo & Zivot (2001) for more details on the estimation of TAR models.

3.3 Relationship between the Econometric Frameworks

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

Within the threshold cointegration framework one can distinguish between two different degrees of border effects. First, a systematic border impact can prevent prices of across-border city pairs to adjust. This strong form of border effects results in a failure of cointegration with respect to the prices of across-border pairs. Second, if the prices still adjust, weak border effects can lead to a slower price adjustment and higher transaction costs, i.e. larger price differences are required to induce price adjustment. The former effect implies that the coefficient α in (3.2) is larger in absolute terms; the latter translates to larger threshold bands when corrected for the distance between the cities. As explained above, Engel & Rogers (1996) relate border effects to higher price variation. However, larger price deviations can be caused by the absence of cointegration but also by a slower speed of price-adjustment and higher transaction costs. Thus, they cannot distinguish between high price volatility and lasting price deviations which prevent the price series from cointegrating. Nevertheless, the Engel-Rogers approach can easily deal with large cross-sections and handle some type of structural instability as well as periods of missing data. In addition, nonsignificant border dummy variables are a strong result since it rules out both strong and weak forms of border effects. In this sense, one may regard the framework of Engel & Rogers (1996) as a first step to study the impact of borders which should be followed by a more detailed threshold analysis if the presence of border effects is indicated.

It should be clear from this section, that the question of price integration can only be addressed within the time series framework which analyses the dynamic properties of the data. This cannot be done using the approach of Engel-Rogers. The threshold cointegration framework also delivers information on the speed of price adjustment and the magnitude of transaction cost bands. Nevertheless, one has to pay a price for that by assuming structural stability over time and running threshold analyses for all city pairs. Hence, both econometric frameworks have advantages and disadvantages and tackle partly different problems. That is why we apply both approaches to exploit as much information out of our data as possible.

4 Empirical Results

4.1 Border Effects

In this subsection we present our findings regarding the importance of the former partition borders. Table 2 summarizes the estimation results for equation (3.1) according to the Engel-Rogers approach with respect to the period 1921:01-1936:12. The estimation is based on 440 observations because we do not have observations for the second half-year of 1923 and we lack data for Wilno until 1923:06.

We see that significant border effects increasing the price variation are only present in the years 1921-1924. Obviously, the magnitude of the effects decreases during this period and after 1924 the corresponding coefficients are not significant different from zero. The results do not change importantly if we sequentially delete the variables with insignificant coefficients. In the latter case the border effects are also significant in the years 1926, 1931, and 1933. However, they are rather small. The size of the dummy coefficients is only one third of the one in 1924.

In a second step we have added yearly dummy variables to equation (3.1) in order to distinguish yearly border and general time effects. Table 3 collects the results for the specification where all insignificant variables have been deleted in a sequential procedure. The log-distance as well as the mean term have always been considered.

Regarding the years 1921-1924 the findings indicate that the former partition borders only matter in 1923. In the other three years general time specific impacts increasing the volatility of the price differences are present. It is a somewhat surprising result that first time specific effects are important, then the border and again the time. Obviously, it is difficult to discriminate between both effects due to the low number of city pairs that make

Table 2. Results for regression equation (3.1) with time specific border effects: 1921:01-1936:12

Variable	Coefficient	Std. Error	p-value
log distance ($d^{(i,j)}$)	0.0029	0.0045	0.5264
City specific dummy variables (C_m)			
Warsaw	0.0117	0.0126	0.3539
Wilno	0.0349	0.0175	0.0466**
Poznań	0.0261	0.0145	0.0731*
Łódź	0.0146	0.0124	0.2428
Kraków	0.0137	0.0142	0.3376
Lwów	0.0236	0.0158	0.1349
Time specific border dummy variables ($B_t^{(i,j)}$)			
1921	0.1209	0.0235	0.0000***
1922	0.0748	0.0168	0.0000***
1923	0.0793	0.0181	0.0000***
1924	0.0221	0.0098	0.0248**
1925	−0.0005	0.0068	0.9465
1926	0.0036	0.0050	0.4701
1927	−0.0030	0.0066	0.6538
1928	−0.0082	0.0057	0.1550
1929	−0.0067	0.0048	0.1682
1930	0.0056	0.0065	0.3912
1931	0.0063	0.0046	0.1716
1932	−0.0048	0.0058	0.4120
1933	0.0052	0.0045	0.2411
1934	−0.0056	0.0053	0.2904
1935	−0.0095	0.0059	0.1050
1936	−0.0021	0.0078	0.7863

Note: We allow for heteroscedasticity in the error terms with respect to the different years. The adjusted R^2 and the standard error of regression are 0.404 and 0.025 respectively. Significance at the 1%, 5%, and 10% level is denoted by ***, **,* respectively. Computations have been done using EViews 4.1.

the dummy variables rather similar. For example, if time and border dummy variables are specified together both dummies are insignificant for the years 1923 and 1924. When deleting one of the dummy variables the other one becomes significant. We interpret these outcomes as having found evidence for the impact of the old borders in the first half of the 1920s. However, these border effects are likely to be part of general time impacts which died out after 1924. Nevertheless, we observe small, but significant, time effects in the years 1927-1929

Table 3. Results for regression equation (3.1) with additional time dummy variables: 1921:01-1936:12

Variable	Coefficient	Std. Error	p-value
log distance ($d^{(i,j)}$)	0.0052	0.0034	0.1239
City specific dummy variables (C_m)			
Wilno	0.0205	0.0036	0.0000***
Poznań	0.0126	0.0026	0.0000***
Lwów	0.0091	0.0030	0.0022***
Time specific border dummy variables ($B_t^{(i,j)}$)			
1923	0.0793	0.0181	0.0000***
1935	−0.0095	0.0059	0.0072***
Time specific dummy variables			
1921	0.1078	0.0180	0.0000***
1922	0.0763	0.0125	0.0000***
1924	0.0250	0.0074	0.0008***
1927	−0.0129	0.0046	0.0053***
1928	−0.0115	0.0040	0.0046***
1929	−0.0085	0.0034	0.0132**
1934	−0.0089	0.0033	0.0071***

Note: The results refer to a specification where all insignificant variables have been deleted in a sequential procedure. The log distance as well as a mean term (results are not reported) are always considered. We allow for heteroscedasticity in the error terms with respect to the different years. The adjusted R^2 and the standard error of regression are 0.521 and 0.023 respectively. Significance at the 1% and 5% level is denoted by *** and ** respectively. Computations have been done using EViews 4.1.

and 1934 and a border effect in 1935 decreasing price variation. Interestingly, the distance does not seem to be a relevant factor. In a standard regression with a time independent border dummy variable, however, the distance coefficient is significant.

The results on the importance of the former partition borders have also implications for the threshold cointegration analysis according to the discussion in Section 3. Since the impact of the borders vanishes in principle after 1924 we should not expect to find evidence for border effects, neither for strong nor weak forms, when examining the results of the threshold analysis covering the period 1924-1937. Indeed, we could no detect any important systematic differences between the within- and across-border city pairs with respect to the cointegration analysis and the estimated threshold models. Hence, in correspondence to the historical description in Section 2 we have found border and time specific effects increasing price variation in the Polish wheat market during the first half of the 1920s. We can conclude, however, that the integration policy in Poland was in fact successful in eliminating these negative effects. Whether the wheat flour prices between the cities really cointegrate, i.e. whether arbitrage processes lead to price adjustment, will be studied in the next subsections within the threshold cointegration analysis.

4.2 Threshold Cointegration Analysis

4.2.1 Preliminary Analysis

The foregoing subsection has shown that only some of the years after 1924 were affected by significant time or border effects. Therefore, we keep the assumption of parameter stability for our time series analysis regarding the period January 1924-April 1937. But we still have to consider a possible break in the deterministic terms in May 1929 owing to a change from inflation to deflation. In the following we do present results for all city pairs instead of focussing on results with respect to a certain benchmark city because some of our findings differ for the various pairs, although not systematically between within- and across-border pairs. This presentation approach is also justified by a comment in Harvey & Bates (2004) which says that (empirical) findings of convergence studies may depend on the specific benchmark chosen. Before presenting the outcomes of the threshold cointegration analysis we study the integration properties of the price series and test for the presence of seasonality.

In Table 4 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). Obviously, all log-series can be regarded as integrated of order one because the null hypothesis of a unit root is not rejected for the level series but rejected for the first differences.

The issue of seasonality is rather important since different seasonal patterns in the price series would raise doubts about market integration. A test by Canova & Hansen (1995)

Table 4. Unit Root Test statistics

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

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

clearly suggests to model possible seasonality in a deterministic way and not stochastically for all city pairs. Therefore we estimated in a second step univariate AR models for the first differences of the log-series including seasonal dummies in order to test for the significance of these dummy variables, i.e. to test for deterministic seasonality. The dummy variables are jointly significant only for Warsaw (10% level). Furthermore, the log-price differences of all 15 city pairs are not affected by deterministic seasonality. This 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.

4.2.2 Results of Cointegration Analysis

The outcome of the generalized Johansen procedure is given in Table 5. The misspecification test for vector autocorrelation described in Doornik & Hendry (1997) suggests no significant autocorrelation for the corresponding vector autocorrelation (VAR) models (compare column 2). We see that the price systems of all city pairs have a cointegrating rank of one at a 5% significance level. In a next step, we have tested whether the cointegrating vector β can be restricted to (1, -1). The results in the 5th column of Table 5 show that this assumption cannot be rejected at a 10% level regarding all city pairs. Hence, the price differences of the respective city pairs are part of a stationary relationship as the LOP requires.

As outlined in Section 3 the same stochastic dynamics for both prices series is not sufficient for price integration as long as trend components enter the cointegrating relationship. In order to study which deterministic terms are present we have applied a sequence of restriction tests. Note that all tests are conducted conditioned on the outcome of the previous ones in the sequence. First, we have tested whether the trend coefficients θ_1 and θ_2 in the cointegrating relationship (compare (3.3)) can be both set to zero. This restriction is rejected with respect to Warsaw-Łódź, Warsaw-Lwów, Wilno-Lwów, Łódź-Lwów, and Kraków-Lwów. For these pairs we have checked if the linear trend components are equal ($\theta_1 = \theta_2$).[5] The null hypothesis $\theta_1 = \theta_2$ is only rejected for Warsaw-Łódź.[6] Thus, we conclude that there occurs no price integration for this city pair owing to a broken linear trend. Regarding the other four pairs convergence is ruled out by a common deterministic trend in the cointegrating relationship.

Regarding the remaining 11 city pairs for which we could impose $\theta_1 = \theta_2 = 0$ we have tested whether linear trends can be completely excluded from the model. This restriction translates to the hypotheses $\nu_i = \Pi\mu_i$ $(i = 1, 2)$ in (3.3) with $\theta_1 = \theta_2 = 0$ which implies that the mean terms can be restricted to the cointegrating relationship.[7] We can maintain this restriction for all pairs. Finally, we have examined if the mean terms of the two subperiods are equal for these pairs, i.e. we have tested the null hypothesis $\mu_1 = \mu_2$.[8] This hypothesis is not rejected with the exceptions of Warsaw-Wilno, Warsaw-Poznań, and Warsaw-Kraków. The final results regarding the deterministic terms are summarized in the last column of Table 5. Detailed results of the various restriction tests are not presented here in order to save space.

Thus, we conclude to have evidence for the validity of a relative version of the LOP for 11 out of the 15 city pairs since the corresponding price differences adjust appropriately to disequilibria without deterministic price divergence. However, the price differences do not eliminate the mean terms in the cointegrating relations. Hence, prices do not adjust toward price parity but to some fixed difference. Interestingly, city specific effects seem to matter

[5] This test is asymptotically χ^2 distributed with one degree of freedom.

[6] We have checked that θ_2 cannot be set to zero for Warsaw-Łódź. The case $\theta_2 = 0$ would mean that the deterministic price divergence has stopped after May 1929.

[7] This test is asymptotically χ^2 distributed with two degree of freedom.

[8] This test is asymptotically χ^2 distributed with one degree of freedom.

Table 5. Results of Generalized Johansen Procedure

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	Deterministic Terms in Cointegrating Relation
Warsaw-Wilno (1)	0.225	$r_0 = 0$	48.72 (0.001)***	0.066*	0.921	broken constant
		$r_0 = 1$	7.13 (0.612)			
Warsaw-Poznań (2)	0.134	$r_0 = 0$	40.72 (0.011)**	0.941	0.718	broken constant
		$r_0 = 1$	11.41 (0.233)			
Warsaw-Łódź (1)	0.111	$r_0 = 0$	50.04 (0.001)***	0.186	0.009***	broken linear trend
		$r_0 = 1$	11.04 (0.260)			
Warsaw-Kraków (1)	0.501	$r_0 = 0$	49.27 (0.001)***	0.614	0.315	broken constant
		$r_0 = 1$	10.29 (0.311)			
Warsaw-Lwów (1)	0.562	$r_0 = 0$	49.61 (0.001)***	0.962	0.030**	linear trend
		$r_0 = 1$	9.84 (0.347)			
Wilno-Poznań (1)	0.106	$r_0 = 0$	39.35 (0.013)**	0.622	0.771	constant
		$r_0 = 1$	7.55 (0.567)			
Wilno-Łódź (1)	0.406	$r_0 = 0$	51.02 (0.000)***	0.509	0.170	constant
		$r_0 = 1$	9.79 (0.351)			
Wilno-Kraków (1)	0.717	$r_0 = 0$	65.43 (0.000)***	0.167	0.503	constant
		$r_0 = 1$	10.59 (0.288)			
Wilno-Lwów (1)	0.319	$r_0 = 0$	63.72 (0.000)***	0.111	0.041**	linear trend
		$r_0 = 1$	11.28 (0.241)			

Table 5. cont'd. Results of Generalized Johansen Procedure

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	Deterministic Terms in Cointegrating Relation
Poznań-Łódź (2)	0.264	$r_0 = 0$	46.55 (0.002)***	0.643	0.112	constant
		$r_0 = 1$	15.27 (0.074)*			
Poznań-Kraków (2)	0.581	$r_0 = 0$	45.60 (0.003)***	0.288	0.333	constant
		$r_0 = 1$	13.43 (0.131)			
Poznań-Lwów (2)	0.537	$r_0 = 0$	39.93 (0.013)**	0.710	0.460	constant
		$r_0 = 1$	14.90 (0.083)*			
Łódź-Kraków (1)	0.365	$r_0 = 0$	51.72 (0.000)***	0.477	0.136	constant
		$r_0 = 1$	9.06 (0.416)			
Łódź-Lwów (1)	0.348	$r_0 = 0$	48.16 (0.001)***	0.313	0.001***	linear trend
		$r_0 = 1$	9.55 (0.372)			
Kraków-Lwów(1)	0.841	$r_0 = 0$	45.10 (0.003)***	0.567	0.025**	linear trend
		$r_0 = 1$	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 computed 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 cointegration analysis has been done with PcFiml 9.10 (see Doornik & Hendry 1997).

Table 6. 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
Warsaw-Wilno	0.006 (1)***	0.718	0.112	0.224
Warsaw-Poznań	0.008 (2)***	0.335	0.706	0.882
Warsaw-Łódź	0.254 (1)	0.392	0.342	0.491
Warsaw-Kraków	0.758 (1)	0.648	0.425	0.248
Warsaw-Lwów	0.157 (1)	0.030**	0.037*	0.040**
Wilno-Poznań	0.000 (1)***	0.000***	0.075*	0.157
Wilno-Łódź	0.015 (1)**	0.000***	0.329	0.579
Wilno-Kraków	0.007 (1)***	0.004***	0.026**	0.035**
Wilno-Lwów	0.307 (1)	0.002***	0.081*	0.431
Poznań-Łódź	0.631 (2)	0.451	0.067*	0.243
Poznań-Kraków	0.787 (2)	0.736	0.654	0.323
Poznań-Lwów	0.262 (2)	0.247	0.118	0.327
Łódź-Kraków	0.122 (1)	0.295	0.774	0.275
Łódź-Lwów	0.269 (1)	0.185	0.586	0.852
Kraków-Lwów	0.351 (1)	0.674	0.790	0.543

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. The Tsay nonlinearity test statistics and their p-values are computed using own GAUSS programs. GAUSS programs from Bruce Hansen's web page (http://www.ssc.wisc.edu/~bhansen/progs/progs_threshold.html) are applied to compute the test statistics for the procedures by Hansen (1997, 1999) and their respective bootstrap p-values.

regarding the occurrence of deterministic terms in the cointegrating relationship: all pairs including Warsaw require a modelling of broken components and four of the five Lwów pairs have a linear trend specification.

4.2.3 Results of Threshold Nonlinearity Tests

Since the cointegration analysis clearly supports price integration for most of the city pairs we proceed to test for threshold effects by applying specific threshold nonlinearity tests. As described in Section 3 we apply the multivariate and univariate Tsay tests (Tsay-M, Tsay-U) and the univariate procedures by Hansen (1997, 1999) which test linearity against a two- and three-regime TAR model respectively (Hansen-U12, Hansen-U13).

The results in Table 6 indicate that threshold nonlinearity does not describe the dynamics of the price series in general. There only seems to be robust evidence for threshold effects for some city pairs including Wilno. One has to be careful, however, in interpreting the outcome of the tests. None of the procedures allows for broken deterministic components or a linear trend. This may be problematic for the pairs containing Warsaw or Lwów. The cointegration analysis has shown that broken deterministic terms or a trend have to be included into the price relationships for the corresponding city pairs. Furthermore, the univariate nonlinearity tests assume stationarity under the null hypothesis. If the price differences are near unit root processes the nonlinearity tests may be size distorted (compare Lo & Zivot 2001). Nevertheless, our results are confirmed by similar findings of a test suggested by Hansen & Seo (2001) which takes account of a linear trend.

Although the evidence for threshold nonlinearity is weak we continue with estimating the TAR models. The estimation results may give us more insights on the relevance of threshold effects because the nonrejection of linearity could be due to small threshold bands.

4.2.4 Estimation of Threshold Cointegration Models

As already explained in Section 3 we do not present estimation results for unrestricted TAR models since no reliable restriction tests are available. Nevertheless, we want to state here that the estimated unrestricted TAR models seem to be far away from what is economically implied by the transaction cost view of the LOP. With this outcome in mind we have estimated the BAND-TAR model (3.2). The estimation results are summarized in Table 7. When estimating (3.2) we consider the deterministic terms according to the findings of the cointegration analysis. Hence, the relationship $z_t^* = p_{1,t} - p_{2,t} + \psi d_t$ is used in line with the discussion in Section 3. Here, d_t contains the deterministic terms given in Table 5 for the respective city pair. In case of a (broken) trend a (broken) constant is considered in addition.

In general, the estimated threshold bands are of importance only for the pairings including Wilno. For most of the other pairs the number of observations in the thrshold regime is very low, in some cases even equal to the minimum number of 24 observations. In line with this low number of observations the width of the threshold bands is rather small, e.g. in case of Warsaw-kraków it amounts to only 1.1% percent of the average wheat flour log-price

Table 7. Estimation of BAND-Threshold Model

City pair	Linear model	BAND-Threshold model				
	$\hat{\alpha}_l$	$\hat{\tau}$	$\hat{\alpha}$	Observations per Regime		
	$(s.e.)$		$(s.e.)$	Lower	Thresh.	Upper
Warsaw-Wilno	0.774	0.032	0.543	58	53	48
	(0.065)		(0.127)			
Warsaw-Poznań	0.719	0.015	0.572	65	30	64
	(0.056)		(0.074)			
Warsaw-Łódź	0.816	0.010	0.636	69	24	66
	(0.047)		(0.073)			
Warsaw-Kraków	0.760	0.023	0.595	54	55	50
	(0.066)		(0.093)			
Warsaw-Lwów	0.876	0.015	0.613	77	24	58
	(0.042)		(0.072)			
Wilno-Poznań	0.717	0.078	0.418	43	70	46
	(0.080)		(0.164)			
Wilno-Łódź	0.803	0.207	−0.078	14	135	10
	(0.064)		(0.248)			
Wilno-Kraków	0.720	0.087	0.393	33	88	38
	(0.082)		(0.178)			
Wilno-Lwów	0.680	0.182	−0.425	15	135	9
	(0.096)		(0.461)			
Poznań-Łódź	0.706	0.068	0.256	42	81	36
	(0.058)		(0.126)			
Poznań-Kraków	0.627	0.020	0.522	66	32	61
	(0.065)		(0.078)			
Poznań-Lwów	0.755	0.066	0.526	40	73	46
	(0.046)		(0.073)			
Łódź-Kraków	0.644	0.009	0.604	64	24	71
	(0.073)		(0.081)			
Łódź-Lwów	0.843	0.012	0.556	68	24	67
	(0.045)		(0.085)			
Kraków-Lwów	0.805	0.013	0.539	74	24	61
	(0.049)		(0.085)			

Note: The given standard errors are computed according to White (1980) to be robust of unknown heteroscedasticity.

of all cities over the whole sample. Obviously, the transaction costs are quite low. This may explain why the estimated unrestricted model deviate so clearly from the BAND-TAR structure and why the threshold nonlinearity tests do not detect threshold dynamics in general: It is difficult to discriminate between linearity and nonlinearity.

As mentioned, the situation is different for the Wilno pairs. The threshold bands are much larger and the width amounts up to 10% of the average price in case of Wilno-Łódź. The larger bands for the Wilno pairs should be no surprise since Wilno is the most remote city. Accordingly, the threshold bands are supposed to be wider owing to higher transportation costs. Interestingly, the threshold nonlinearity tests suggest threshold effects rather robustly with respect to the Wilno pairings. We have to note regarding Wilno-Łódź and Wilno-Lwów, however, that the outer regimes contain only the minimum number of observations. Thus, adjustment is merely observed for a low number of time points. Accordingly, the adjustment coefficient is estimated quite imprecisely and is even negative. The latter means that adjustment occurs in an oscillating pattern. However, we may also interpret these findings as evidence against threshold effects since the outer regimes seem to be unimportant.

In line with the foregoing discussion we observe the strongest reduction in the coefficient α for the Wilno-pairings. The half-life h for Wilno-Poznan e.g. reduces from 2.08 months in the linear model to 0.79 months in the threshold model. Hence, the price adjustment increases clearly if we account for threshold effects, i.e. adjustment is faster if we are further away from the equilibrium. [9] In general, the reduction in α for the non-Wilno pairs is less pronounced lower although still present.

Although there occur differences in terms of the estimated threshold models they are not systematic with respect to the within- and across-border city pairs. When relating the estimated threshold parameter $\hat{\tau}$ and the adjustment coefficient $\hat{\alpha}$ to the log-distance and a border dummy variable only the distance has a significant impact.

Therefore, we can conclude in the following way. With the exception of the Wilno pairings, the threshold bands are not very large compared to the absolute wheat flour prices. This implies relatively low transaction costs. Taking also the outcome of the nonlinearity tests into account we see that threshold nonlinearity seems to be present but its degree is small. We interpret this as a further sign for functioning price adjustment and economic integration in line with the main results of the cointegration analysis.

[9] 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).

5 Summary and Concluding Remarks

In this paper we have studied the topic of integration across borders for the Polish interwar economy by analysing the wheat flour market with respect to six cities between 1921 and 1937. To be precise, we have asked whether borders hinder economic integration where we consider an economy as integrated if the LOP holds taking transaction costs into account.

Let us briefly summarize our findings. First, using the approach of Engel & Rogers (1996) we find evidence that the former partition borders matter until 1924, but that their effect vanishes after that date. In fact, we have evidence that a relative version of the LOP holds for 11 out of our 15 city pairs during the period 1924-1937. That means that the prices of wheat flour adjust to disequilibria but a (broken) constant enters the price relationship. Hence, some price differences remain between the market places so that integration is not perfect. It seems that city specific effects matter in this respect but there is no clear pattern that suggests a persistent effect of the former partition borders. Among the 15 city-pairs we have four parings that were not formerly separated by a partition border. In 11 cases we find a cointegration relationship without trending components, including three of the four within-border pairs. Moreover, our findings on threshold bands of no-arbitrage suggested again rather city-specific factors than border effects. E.g. the multivariate Tsay-Test gives evidence that there was a significant threshold-nonlinearity for city-pairs including Wilno, the most remote city in our sample, but this applied to both within-border parings (Wilno-Warsaw, Wilno-Łódź) and across-border pairings (Wilno-Poznań, Wilno-Kraków). Hence, nonlinearities and the implied transaction costs seem to exist but they are generally small and not related to the former borders. Thus we regard the Polish interwar economy as integrated with some restrictions and interpret this as a success of the Polish integration policy after the reunification in 1919.

Our results for Interwar Poland markedly differ from the findings of Moodley et al. (2000) and Engel & Rogers (1996) who analyse economic integration between U.S. and Canada in the wake of the Free Trade Arrangement (FTA) of 1990. Interestingly, both studies show no important effect of the FTA on the degree of integration in contrast to our results for changing borders in Poland. Most probably, this is due to the fact that integration in the Polish case involved not only the removal of tariff barriers, but improved regional mobility and communication along virtually all aspects.

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