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Abstract

In Malthusian economies, crop shortages could be a matter of life and death. The development of regional and national markets for grain held the potential to provide insurance against the demographic consequences of local crop failure. Weather shocks that are reflected in price data, however, entail a measurement problem for market integration studies, which we solve against the background of the end of the Little Ice Age. We exemplify our method to measure price convergence and the link from grain prices to mortality for Germany based on a new data set of rye prices for 15 cities in 1650–1790. We find price convergence in North-Western Germany as well as along major rivers. In addition, a substantial moderation of aggregate rye price volatility occurred, which we link theoretically to increased arbitrage. Mortality was positively related to the aggregate rye price and thus, the decline of rye price volatility decreased the risk to die of hunger in pre-industrial Germany.

JEL Classification: N13, N53, O13.

Keywords: Market integration, price volatility, agriculture, weather, climate, unified growth theory.

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1 Introduction

Grain prices rank among the longest available economic time series and are one important source of information about the evolution of pre-industrial economies. In Europe, bread from grain constituted the major spending share in consumer baskets for the pre-industrial period (ca. 30%, Allen, 2001). Economic historians study market integration, because it improves allocative efficiency via market access (Federico, 2012; Kelly, 1997). In addition to increased allocative efficiency of production factors, better integration of *agricultural* markets in pre-industrial economies is also likely to have improved the allocation of grain for consumption. Integrated markets are expected to have a lower price volatility (Chilosi et al., 2013; Jacks et al., 2011). In a Malthusian regime, which characterized industrialized economies for most of history (Galor and Weil, 2000), more stable food prices could lead to higher food security and thus attenuate the 'positive check', that is, the negative relationship between income and the death rate, which is often assumed to work through hunger crises.

One potential alternative explanation for price convergence and/or declining price volatility are weather shocks.[1] A large part of the market integration literature analyzes grain prices that are highly influenced by weather. Trend estimates of price convergence might be influenced by weather shocks, a problem the literature has not addressed thus far. A further issue is climate change; here in particular the Little Ice Age (LIA) that began to end in the early eighteenth century. The LIA was a period of cooler climate in pre-industrial Europe and included the Maunder Minimum (1645–1715), a period of lower solar irradiance (Masson-Delmotte et al., 2013, 389). One could argue that the increase of mean temperature raised the level of grain output and thus impacted on price levels in a direct way. Furthermore, a decline of weather shocks may have reduced local crop failures that would have been reflected in lower price volatility independent of the state of market integration (cf. Federico, 2012, 484; Chilosi et al., 2013, 48).

We exemplify this methodological problem and the link from grain prices to mortality by analyzing pre-industrial Germany, one of the most populous economies of 18th century Europe (Malanima, 2010). We compile a stable sample of rye prices with less than 5% missing observations from 15 German cities in 1651 until 1790. The latter period defines the 140 years between two major historical events: the end of the Thirty Years' War and the

[1] Like most of the economics literature, we refer to weather as inter-annual variation of meteorological variables (Dell et al., 2014). Climate is the long-run temporal mean and variability of weather (usually 30 years); "Climate change refers to a statistically significant variation in either the mean state of the climate or in its variability, persisting for an extended period (typically decades or longer) (WMO, 2017b)." E.g., in our framework, inter-annual variations of spring temperature are weather; the mean and the standard deviation of spring temperature across 30 years are climate; a significant persistent change in that 30-year-mean and/or the 30-year-standard deviation is climate change.

War of the First Coalition following the French Revolution. Rye was the most important cereal used for bread, the main source of calories in Germany (Pfister, 2017), which is also evident in cropping shares: 38% rye, 23% oats, 17% barley and 7% wheat (van Zanden, 1999, 368, table 16.6). Additionally, data from the urban grain market in Cologne show that ca. 71.8 % of the traded quantity of cereals was made up of rye, 10.8% of wheat, 10.9% of barley and 6.5% of oats, respectively (average of monthly data in 1716–25 from Ebeling and Irsigler, 1976). To ease comparison with earlier work that focused mostly on the cash crop wheat, we extended the data set to the latter. To gain a broader picture, we also included barley and oats. Altogether, our database comprises more than 100 time series.

An important method to measure market integration as price convergence is the cross-sectional coefficient of variation (CV) (Chilosi et al., 2013; Federico, 2011). To address the problem of measuring price convergence robust to weather shocks and climate change, we formally analyze how shocks impact on the CV. The main insight is that the CV is neither robust to weather shocks nor to climate change. We argue that the measurement of market integration based on grain prices can be improved in two ways. First, one component of the CV, the standard deviation (SD), is affected by shocks to a lesser extent, although deflation of data becomes necessary when using this indicator. Second, by analyzing five-year average prices rather than annual prices a substantial part of the potentially distorting short-term variation caused by weather shocks can be removed from grain prices.

Our empirical results on Germany bear out two stylized facts: first, cross-sectional price convergence at the national level and, second, a decline of the temporal volatility of the aggregate rye price in Germany. At the national level, prices converged by 0.3% per year (1651–1790) from 31% cross-sectional dispersion relative to the mean price (1651–75) to 17% (1766–90; based on the standard deviation of real five-year average prices). Temporal volatility declined from 26% (1651–75) to 14% (1766–90). We consider the magnitude of the decline of price volatility across time so relevant that we term it the *Great Moderation of Grain Price Volatility*. Our results show much stronger market integration than previous studies (e.g., Bateman, 2011). Furthermore, our methodological approach ensures that price convergence is measured robust to weather shocks and climate change. To relate the moderation of grain price volatility to increasing price convergence, we conduct a formal analysis of temporal price volatility and show that the decline in aggregate price volatility must be at least partly a result of spatial arbitrage, which is reflected in price convergence.

The main consequence of market integration is Smithian growth (Kelly, 1997). In addition, decreasing grain price volatility might have had further effects. Empirically, growth rates of the aggregate rye price and mortality are significantly positively related; a 1%

increase of the lagged rye price is associated with a 0.49% increase of the death rate (1730–1790), which points at a strong positive check in Germany. But the price volatility moderation coincided with a decrease of the variability of the German crude death rate (Pfister and Fertig, 2010). By way of this channel, lower grain price volatility likely increased the survival probability of children (and adults) which points at two additional aggregate effects of grain market integration beyond Smithian growth. First, the positive check might have become weaker. Second, higher survival probability might have reduced pre-cautionary demand for children. In the unified growth model developed by Kögel and Prskawetz (2001), increasing survival probability is an important exogenous trigger for the transition to sustained growth.

The study is organized as follows. We present our data in Section 2 and describe the aggregate rye price against the background of evidence on subsistence crises in Section 3. We then explore how the CV is affected by weather shocks and climate change in Section 4 and develop a method to identify price convergence that is not contaminated by weather shocks or climate change. Section 5 presents our empirical results as two stylized facts. We document the evolution of price convergence and the decline of grain price volatility. In Section 6, we show that market integration appeared mainly along major rivers and in all segments of the grain market. Section 7 analyzes the relationship between the aggregate rye price and mortality. Section 8 concludes.

2 A new data set of pre-industrial grain prices in Germany

This study develops a new data set of grain prices from the 15th to the 19th century. In what follows, we focus on a stable sample of calendar year rye prices for 15 cities in 1651–1790 (Figure 1, black) measured in grams of silver per liter (g Ag / l). The stable sample includes price series with $\leq$5% missing observations. Next to the stable sample, Figure 1 shows 14 additional cities which we include in the unbalanced sample (grey). To account for inflation in our analysis of price convergence, we deflate all price series using the consumer price index (CPI) developed by Pfister (2017).

Cities of the stable sample are either part of North-Western Germany (black filled square) or continental Germany (white filled square). This sample split is applied to account for the possibility of regionally specific climate change, which we discuss below in Section 4.3. The criterion for the sample split is the continentality of the climate measured by the within-year temperature difference between the hottest and the coldest month. If the latter is ≤ 17.5°C, a city is allocated to the North-Western sub-sample; otherwise to continental Germany (see supplementary appendix SA7 for details).

For extensions of the main results we constructed further data sets for other cereals,

namely wheat, barley and oats. We compiled these grain price data from edited and few selected archival sources, partly unused thus far. To improve the data quality, we address how grain price data from sources containing crop year prices can be converted to a common time base (calendar year) by building on the procedure proposed by Bateman (2011, 451)

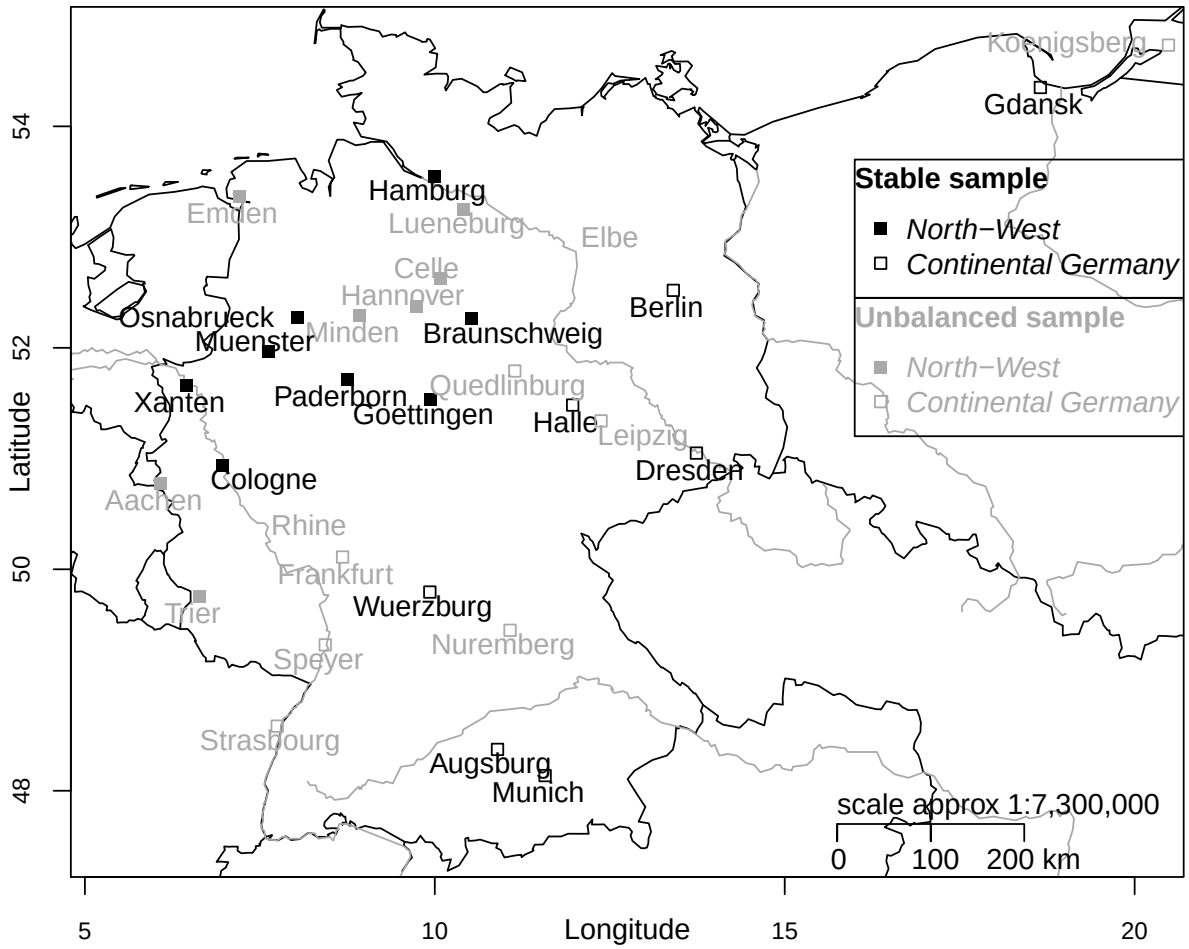

Figure 1: Cities in the sample. Stable sample 1651–1790: $\leq$5% missing observations per individual series. Cities are allocated to North-West if continentality$\leq$ 17.5°C. Data on continentality from Müller-Westermeier et al. (2001, map 7) and WMO (2017a).

who approximates calendar year prices with averages of neighboring crop year prices. We apply this approach to a further type of prices found in sources, *Martini* prices.[2] The transformation of Martini prices increases the correlation of prices in our stable sample.

[2]Customarily, many peasant obligations were due on Martinmas, November 11, so market liquidity was particularly good around this time. Since *Martini* prices sometimes served as reference to monetize arrears of peasant dues, they had a good chance of being recorded and preserved.

Hence, the standardization of the time reference of grain prices is highly relevant for market integration studies (Table S12 in SA3.1.2). We also develop regional series of silver contents for currency conversion, which is crucial for analyses of price dispersion accross space. All information on data preparation and sources for each individual city and cereal are relegated to SA1–SA4. While Federico et al. (2018) present a very broad European picture for wheat, our study solves several data problems and provides data sets for four cereals with the focus on rye, the most important food grain in pre-industrial Germany.

3 Evolution of the aggregate real rye price

For the analysis of aggregate volatility, we calculate the aggregate real rye price as the arithmetic mean of the city prices in our sample. We deflate the nominal aggregate price with the national CPI by Pfister (2017; see SA6 for details).

Panel (a) of Figure 2 shows the aggregate real rye price for Germany together with a smoothed trend for the stable sample series, and panel (b) the detrended price or, effectively, the cyclical component.[3] The stable sample for 1651–1790 contains rye prices for 15 cities for 140 years and contains 2093 observations. The trajectory of the aggregate real price of rye based on the stable sample (black solid line) is consistent with the one based on a larger unbalanced sample (black dashed line; 2983 observations from 29 cities).

The peaks of the cyclical component (marked with vertical bars and corresponding years in Figure 2) indicate serious food or subsistence crises, because they can be linked to increasing mortality and decreasing fertility rates. On average, notably more peaks occurred during the first half of the sample period, that is, until 1720 (nine vs. five peaks). Furthermore, on average, most price peaks are larger before 1720, particularly in the 17th century. Until 1700 all but the crises of 1684/85 and 1698/99 show up in the series of mortality crises documented for France by Dupâquier (1989, 191–2); they obviously correspond to continental food crises. In addition, from 1675/76 all crises show up in regional series of vital events developed by Pfister and Fertig (2010). The crisis of the early 1690s is reputed as one of the worst food shortages of the post-Thirty Years' War Ancien Régime (Ó Gráda, 2005; Ó Gráda and Chevet, 2002). The crisis of 1698/99 seems largely unknown but appears to have been particularly severe in Germany.

The years 1700–1725 saw four price peaks in 1709, 1713/14, 1719/20 and 1724/25. All but the one in 1713/14 show up in demographic data for France, and all crises can be identified in the regional series of vital events for Germany. During the period 1726–1790 only four

[3]Plots of all individual nominal rye price series and the aggregate nominal rye price are available in the SA (Figures S2–S4 and Figure S6).

price peaks occurred. The crises of 1740/41 and the early 1770s are well-known European crises (Post, 1985; 1990). The price peaks 1756/57 and 1761/62 are less known but may be confounded with inflationary pressures connected with the Seven Years' War (1756–1763). Nevertheless, all crises can be clearly identified in national series of vital events (Pfister and Fertig, 2010, 31).

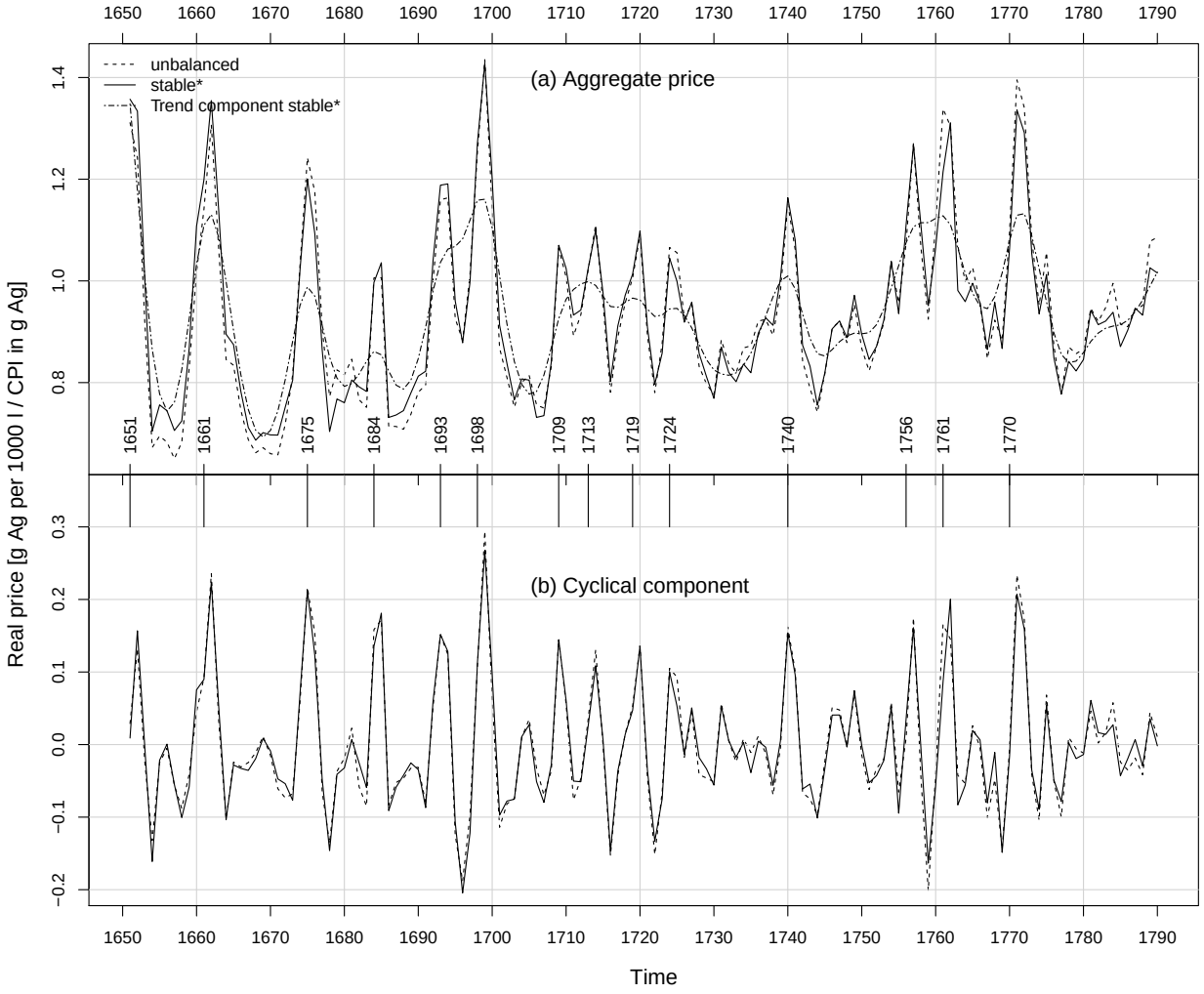

Figure 2: Aggregate real rye price Germany. Trend (shown for stable sample series only) and cyclical component from Hodrick-Prescott-Filter, $\lambda = 6.25$ (Ravn and Uhlig, 2002). Vertical lines and given years on upper horizontal axis in panel (b) mark major price peaks associated with subsistence crises. *: $\leq$5% missing observations per individual series. The aggregate nominal rye price is deflated with the CPI from Pfister (2017). Data sources: see SA2.

Several of these subsistence crises have been associated with weather shocks (e.g., 1741; Post, 1985). In fact, the aggregate rye price is statistically significantly related to aggregate

temperature and precipitation reconstructions from Luterbacher et al. (2004) and Pauling et al. (2006) (data from NOAA, 2017; 2018; regression results reported and discussed in detail in SA8.2). The signs of the relationships we find are in line with evidence from long-term agronomic experiments for winter rye yields (Chmielewski, 1992; Chmielewski and Köhn, 2000) (all details in SA8.2).

The existence of short-term relationships between weather variables and the aggregate real rye price creates the possibility that grain market conditions were influenced by climate change. Indeed, the behavior of weather variables underwent significant change during the period under study. First, the negative peaks in the winter temperature series from Luterbacher et al. (2004) (defined as mean minus twice the standard deviation), which is a significant predictor of rye prices during the period 1651–1720, disappear almost entirely during 1721–1790 compared to 1651–1720 (Figure S7). Second, annual spring temperature from Xoplaki et al. (2005) was higher in 1721–90 than in 1651–1720. The difference of +0.41°C is significant at the 1% level (Newey-West standard errors; data are *not* smoothed; see critique by Kelly and Ó Gráda, 2014; further details in SA8). Additionally, the Kwiatkowski-Phillips-Schmidt-Shin (KPSS) test rejects level and trend stationarity for both weather variables for the period 1651–1790. Warming in parts of the northern hemisphere around 1700 has been associated with the waning of the Little Ice Age (Masson-Delmotte et al., 2013, 389, 409; Kelly and Ó Gráda, 2014, 1374; Xoplaki et al., 2005, 2).

The aggregate real rye price is stationary or trend stationary during the period 1651–1720 (Augmented Dickey Fuller (ADF) and KPSS-tests; Tables S25 and S26). However, these time series properties do not carry over to the second half of the sample period 1721–1790: For these 70 years the aggregate real rye price cannot be regarded as stationary anymore as evidence from ADF and and KPSS tests shows (Tables S25 and S26).[4] Next to climate change, market integration constitutes a potential explanation for the changing time series properties of the aggregate price. Even without inducing technological progress, trade between cities leads to regional specialization according to comparative advantage and increases total factor productivity at the aggregate level (for the theoretical background consider a Ricardian trade model, e.g., van Marrewijk, 2012, chapter 3).

For the 15 individual rye price series of the stable sample the evidence from ADF-tests is less ambiguous than at the aggregate level. For all prices series, ADF-tests reject the null hypothesis of a unit root regardless of whether the series is nominal, real or in logs (see Table S19 in SA5.1). Similarly, the majority of price series is classified as stationary or trend stationary based on KPSS-tests (Table S22). If we split the sample period, a unit root

[4]The ADF and KPSS tests using data for the entire period 1651–1790 point at level or trend stationarity of the aggregate real price (Tables S25 and S26).

can be rejected at the 10% level in all except one city for the years 1721–1790 (Paderborn; Tables S20 and S21). The KPSS test results suggest that more series contain a unit root compared to the ADF test (four of fifteen price series for the years 1721–1790), however, for both sub-periods (1651–1720 and 1721–1790), the majority of real rye price series can be regarded as level or trend stationary based on KPSS tests (Tables S23 and S24).

According to Dobado-Gonzàlez et al. (2012, 682–3), market-clearing prices should be non-stationary. This results from the fact that *persistent* shocks to prices translate into the behavior of relative prices of two integrated markets. E.g., the level of technology follows a non-stationary process in most economies (for a somewhat different argument based on the serial correlation of harvest damages, see Shiue and Keller, 2007, 1198).[5] If one follows this line of reasoning, stationary prices mean that the German economy would have been characterized by non-clearing grain markets during the period under study. However, non-stationary technology is an attribute of post-Malthusian and modern economies; a Malthusian economy with largely static technology is characterized by stationary real wages and vital rates (Møller and Sharp, 2014, 107, 115–7). In pre-modern Germany, real wages and vital rates were stationary or trend-stationary over longer periods of time (Pfister and Fertig, 2010, 43; Pfister, 2017, 713, 717). Thus, the stationarity of most grain prices reflects the predominantly Malthusian character of the German economy during the period under study and the absence of persistent shocks and does not necessarily constitute a sign of non-clearing markets.

4 Anatomy of the coefficient of variation

Stationarity of of most grain prices in our sample for most of the time precludes the application of cointegration analysis to study market integration in a Malthusian setting (see, e.g., Dobado-Gonzàlez et al., 2012, Jacks, 2004, Kopsidis, 2002 or Shiue and Keller, 2007 for applications of this methodology). Hence, we focus on a simple but straightforward measure to test the law of one price (LOP), the coefficient of variation (CV; Federico, 2012). As will be shown below, the CV of grain prices does not only reflect the extent to which the LOP holds but is also sensitive to weather shocks to agricultural output. Furthermore, by altering agricultural supply conditions, climate change that took place during the period under study may potentially invalidate conclusions concerning market integration derived from the CV. Therefore, based on an analysis of the CV and of its main component, the standard deviation (SD), this section develops a framework for analyzing market integration

[5]Cereals are annual crops. Thus, a harvest shock cannot be persistent for a cereal plant like rye as such. Still, if seed density and/or labor productivity are lower in years following major crop failure, persistent effects of harvest shocks might prevail.

that is robust to changes in agricultural supply conditions emanating from weather shocks and climate change. Section 4.1 analyzes shocks in the CV. Sections 4.2 and 4.3 discuss how to measure price convergence robust to weather shocks and climate change, respectively.

4.1 Shocks in the cross-sectional coefficient of variation

We analyze several cases of shocks depending on how these affect cross-sectional units (all: symmetric or only one: asymmetric), on the presence of arbitrage (perfect or none) and on the type of shock (absolute or proportional to price level).

In what follows we assume that a shock alters the price level in an absolute way. Later we relax this assumption and discuss several cases of proportional shocks. This is necessary, because we cannot rule out either type of shock on empirical or theoretical grounds. While illustrating the key ideas, the representation for the absolute shock is simpler and hence, we start with this variant.

4.1.1 Symmetric absolute shock

The CV is calculated as follows:

$$CV_t = \frac{\sqrt{\frac{1}{N-1}\sum_{i=1}^{N}(p_{it} - \bar{p}_t)^2}}{\bar{p}_t} \tag{1}$$

with city $i = 1, \ldots, N$, year t, and $\bar{p}_t = \frac{1}{N}\sum_{i=1}^{N} p_{it}$.

The effect of a symmetric absolute shock s_t to the prices in all cities on the mean price is:

$$\begin{aligned} \bar{p}_t^z &= \frac{(p_{1t} + s_t) + ... + (p_{Nt} + s_t)}{N} \\ \bar{p}_t^z &= \frac{p_{1t} + ... + p_{Nt}}{N} + s_t = \bar{p}_t + s_t, \end{aligned} \tag{2}$$

where superscript z indicates the cross-sectional mean price $\bar{p}_t$ including the shock. The shock s_t cancels from the sum of squared deviations:

$$\sum_{i=1}^{N}(p_{it} - \bar{p}_t^z)^2 = (p_{1t} + s_t - [\bar{p}_t + s_t])^2 + ... + (p_{Nt} + s_t - [\bar{p}_t + s_t])^2. \tag{3}$$

Consequently, the symmetric shock remains only in the denominator of the CV.

$$CV_t = \frac{\sqrt{\frac{1}{N-1}\sum_{i=1}^{N}(p_{it} - \bar{p}_t)^2}}{\bar{p}_t + s_t} \tag{4}$$

A symmetric positive price shock $s_t > 0$ affecting all markets equally decreases the CV. Severe weather shocks leading to price increases in all markets might thus be misunderstood as price convergence signaling market integration.

4.1.2 Asymmetric absolute shock

The effect of an asymmetric absolute shock s_{1t} to the price in one city depends on whether arbitrage between cities takes place or not.

Perfect arbitrage

Arbitrage leads to an equal[6] distribution of the shock across all cities. That is, each of the N cities experiences a share s_{1t}/N of the local shock s_{1t}.

$$\begin{aligned} \bar{p}_t^z &= \frac{(p_{1t} + s_{1t}) + ... + p_{Nt}}{N} \\ \bar{p}_t^z &= \frac{p_{1t} + ... + p_{Nt}}{N} + \frac{s_{1t}}{N} = \bar{p}_t + \frac{s_{1t}}{N}. \end{aligned} \tag{5}$$

This shock affects the sum of squared deviations as follows:

$$\sum\nolimits_{i=1}^{N} (p_{it} - \bar{p}_t^z)^2 = (p_{1t} + \frac{s_{1t}}{N} - [\bar{p}_t + \frac{s_{1t}}{N}])^2 + ... + (p_{Nt} + \frac{s_{1t}}{N} - [\bar{p}_t + \frac{s_{1t}}{N}])^2. \tag{6}$$

The local shock cancels. However, the CV is still affected and decreases as well (if $s_{1t} > 0$), but to a lesser extent ($\frac{s_{1t}}{N}$ in the denominator, not s_t). This result confirms Rönnbäck (2009, 101), who states that a statistical drawback of the CV is its dependence on the equilibrium price level.

No arbitrage

In this case, the shock s_{1t} to city 1 does not spread to any other city and thus affects the squared deviations as follows:

$$\begin{aligned} \sum\nolimits_{i=1}^{N} (p_{it} - \bar{p}_t^z)^2 &= (p_{1t} + s_{1t} - [\bar{p}_t + \frac{s_{1t}}{N}])^2 + ... + (p_{Nt} - [\bar{p}_t + \frac{s_{1t}}{N}])^2 \\ &= (p_{1t} + (1 - \frac{1}{N})s_{1t} - \bar{p}_t)^2 + ... + (p_{Nt} - [\bar{p}_t + \frac{s_{1t}}{N}])^2. \end{aligned} \tag{7}$$

The shock does not cancel. In addition, the effect is ambiguous in sign. The sum of squared deviations *decreases*, if the shock s_{1t} moves the price of the city experiencing the shock, $p_{1t}^z = p_{1t} + s_{1t}$, closer to the mean price with shock ($\bar{p}_t^z$) relative to the situation without shock. For example, the prices in three cities are $2, 4, 6$. Assume an adverse weather shock that increases the price of the first city $p_{1t} = 2$ to $p_{1t}^z = 3$. Now the squared difference of the price with shock and the mean price with shock is *smaller* compared to the situation before the shock: $(p_{1t}^z - \bar{p}_t^z)^2 \approx (3 - 4.33)^2 < (2 - 4)^2 = (p_{1t} - \bar{p}_t)^2$. The sum of squared deviations *decreases* and thus, the SD decreases from 2 to ca. 1.53. The CV would decrease from 0.5 to about 0.35.

[6] This is a simplifying assumption. The share of the shock which each city absorbs depends on trade costs that are not equal across cities in reality.

The denominator of the CV (the mean price including the shock, that is, $\bar{p}_t^z$) will unambiguously increase in case of a positive price shock. Hence, in our example, the denominator amplifies the decrease of the numerator, because now a smaller numerator is divided by a larger denominator. Specifically, the SD decreases to 76.5% of the pre-shock value but the CV to 70%. By contrast, if the shock moves the local price further away from the mean, the increase of the denominator will dampen the increase of the SD.

4.1.3 Proportional shock

We summarize the main results of this sensitivity check and relegate all details to SA15, where we show that the results are qualitatively equivalent to those for absolute shocks with two important exceptions. First, compared to absolute symmetric shocks, there is a difference between the cases with *perfect arbitrage* and *no arbitrage* under proportional symmetric shocks. Second, with a positive *proportional symmetric shock rate and no arbitrage*, the SD increases while the CV remains unchanged. This result is intuitive because it illustrates why the CV corrects for symmetric monetary inflation while the SD does not (Rönnbäck, 2009, 102). As we rely on the SD, we thus have to deflate all prices to avoid increases of the SD due to inflation.

4.2 Measuring price convergence robust to weather shocks

Cereals are annual crops and thus, annual local prices are strongly influenced by weather shocks as long as markets are not characterized by perfect arbitrage. The insight of the formal analysis of the CV is that it is neither robust to spatially symmetric nor asymmetric shocks except for one case, namely, a *symmetric proportional shock with no arbitrage*.

In the presence of weather shocks, application of the CV might lead to a misinterpretation of price convergence as market integration. This results from the way how shocks affect SD and mean price level. Shocks that lead to large jumps in the cross-sectional CV might affect trend estimates, which are usually performed to measure price convergence over time. We use the cross-sectional SD of prices instead of the CV, because the SD is less affected by shocks as shown in the previous Section 4.1. While the CV reacts to weather shocks in all but one case out of seven, measurement using the SD reacts in only three out of seven cases (Table 1). We now discuss how we deal with these cases.

To attenuate the influence of (absolute or proportional) asymmetric weather shocks on the SD, we average the price for each city over five years. Weather can be regarded as random across years (Schlenker and Roberts, 2009, 15596). Thus, the choice of a five-year-period is based on the idea of approximating a zero mean shock for each city, i.e., to nullify the effect

Table 1: Effect of weather shocks on cross-sectional CV and SD

Case	CV annual data	SD annual data	SD 5-year-averages
Absolute price changes due to weather shock			
- Symmetric	↓↓	(.)	(.)
- Asymmetric, perfect arbitrage	↓	(.)	(.)
- Asymmetric, no arbitrage	↓↑	↓↑	(.) assumption: local shock nullified by average
Proportional price changes due to weather shock			
- Symmetric, perfect arbitrage	↓↓	(.)	(.)
- Symmetric, no arbitrage	(.)	↑↑	(.) assumption: local shock nullified by average
- Asymmetric, perfect arbitrage	↓↓	(.)	(.)
- Asymmetric, no arbitrage	↓↑	↓↑	(.) assumption: local shock nullified by average

Note: Arrows show sign (number of same arrows the intensity) of change of the considered measure in reaction to a price increase resulting from a decrease in agricultural output. (.): denotes that the considered measure is not affected. Source: own representation.

of annual asymmetric shocks to supply and the resulting fluctuations of local grain prices. Admittedly, the choice of five years is to some degree arbitrary and constitutes a compromise between approximating long-run average conditions (to get closer to climate) and loosing too many observations for making statistically valid inferences.[7] In addition, from an economic perspective, using five-year-averages brings prices closer to equilibrium than relying on annual averages.

The remaining problem with regard to weather shocks is that the SD increases in the case of a positive *proportional symmetric shock without arbitrage*. This problem can also be attenuated by using five-year-average prices under the plausible assumption that we nullify shocks by averaging. The difference with respect to the cases of absolute or proportional asymmetric weather shocks is that the full sample is affected instead of only one or a few cities.

A potential alternative approach consists in applying a moving average (MA) filter to each cross-sectional time series (to keep more observations) and then use the filtered time series where high frequencies from weather shocks are attenuated or even eliminated. But the MA filter can create irregular cycles which blur the measurement rather than improving it (Kelly and Ó Gráda, 2014, 1387–8).

To summarize, measuring price convergence with the CV (as is current practice in the literature) makes the tacit assumption of one particular type of shock, namely common shocks which are proportional to the price level under the additional assumption of *no arbitrage*. The measurement of price convergence is unaffected by weather shocks or climate

[7]The main results are very similar, if we use seven- or eleven-year-average prices.

change only in this particular case. However, demonstrating the existence and extent of spatial arbitrage is the purpose of any (grain) market integration study and hence, the assumption of *no* arbitrage cannot be credibly defended.

We showed formally that one component of the CV, the SD, is affected to a lesser extent by weather shocks and develop our empirical strategy to measure market integration based on this insight. However, the SD does not constitute a perfect measure of market integration, because in the most interesting case also the SD is susceptible to shocks (spatially asymmetric shock, no arbitrage). To minimize the effect of local shocks, we use five-year-average prices. Furthermore, deflation of prices using a CPI is necessary to account for monetary inflation.

4.3 Measuring price convergence robust to climate change

Climate change can be analyzed in two categories: changes in the mean state and/or the variability around that mean state. If the mean state (e.g., average temperature) or the variability around it (i.e. the frequency/severity of weather shocks, e.g., very low temperatures) *changes* equally across all cities, this could affect our empirical strategy to use cross-sectional SD of five-year-mean prices as follows.

By construction, the SD is affected by symmetric shocks in only one case: *proportional symmetric shock with no arbitrage*. Assume that the end of the LIA led to an increase in temperature (change in mean state) which benefitted yields at low temperature levels. Output would increase and the real price decrease, *ceteris paribus*. If this effect is strong enough, it could lead to a downward trend in the real price of rye, which would decrease the standard deviation and bias results towards price convergence and thus, market integration.

Furthermore, the variability around the mean state could have changed. Specifically, against the background of the end of the LIA, we expect that shocks to output became less frequent and less severe over time and thus, fewer and less pronounced price spikes occurred in the real price of rye. The latter phenomenon could also lead to a downward trend of the real rye price, because the changing variance pattern would be persistent and create less price spikes after the beginning end of the LIA in ca. 1700. Note that this downward trend would appear even if we use five-year-average-prices. In other words, we cannot 'average away' the changing variance pattern as we do with weather shocks. A downward trend in the real price would decrease the standard deviation and bias our results towards market integration (Table 2).

We can approach both dimensions of the issues created by climate change (change in mean state of climate; change in the variability) by evaluating whether we observe any downward (upward) trend for the real price for the full sample (Table 2).

Table 2: Effect of climate change on cross-sectional CV and SD

Case	CV annual data	SD annual data	SD 5-year-averages
Absolute price changes due to climate change			
- Symmetric	as absolute price changes due to symmetric weather shock		
- Asymmetric, perfect arbitrage	as absolute price changes due to asymmetric weather shock		
- Asymmetric, no arbitrage	↓↑	↓↑	↓↑**
Proportional price changes due to climate change			
- Symmetric, perfect arbitrage	as proportional price changes due to symmetric weather shock		
- Symmetric, no arbitrage	(.)	↑↑	↑↑*
- Asymmetric, perfect arbitrage	as proportional price changes due to asymmetric weather shock		
- Asymmetric, no arbitrage	↓↑	↓↑	↓↑**

Note: Arrows show sign (number of same arrows the intensity) of change of the considered measure in reaction to a price increase resulting from a decrease in agricultural output. (.): denotes that the considered measure is not affected. *: problem approached by evaluating trend of average real price; **: problem approached by sample split and evaluation of trends of average real prices in subregions. Climate change refers to (1) a change of the "mean state" and/or (2) a change of the variability around the "mean state" (= persistent change of weather shocks). Source: own representation.

However, persistent changes can also appear only in a sub-sample of cities characterized by a specific climate that differs from the one experienced by the rest of the sample. Consequently, a sub-region could exhibit a different trending behavior compared to the full sample's average real rye price, which biases the SD at the national level. To deal with this problem, we split the sample according to the climatological criterion introduced in the data section (continentality).

When focusing on sub-regions, we analyze their within variances. We thus follow Federico (2011, 97, 125) to track down the source of price convergence using a variance decomposition. We then evaluate the trends of the two sub-regions' average real prices. If we observe a downward trend of the real price in a sub-region, results using the SD of five-year-average-prices are biased towards price convergence in this sub-region.

Price convergence in the full sample might also be the result of decreasing between variance. If a sub-region shows a downward trending real price, national level price convergence could be the result of this downward trend. Hence, we check whether the trend of the average real price in a sub-region drives inter-regional price convergence.

In short, we henceforth argue that the SD of real five-year prices allows to eliminate the remaining effects of weather shocks and monetary inflation on the measurement of market integration. The assumption we rely on is that five-year mean prices approximate local equilibrium prices, not distorted by weather shocks. With regard to climate change, we evaluate trends in real prices to understand whether negative trends could bias results towards price convergence. Furthermore, we assume that splitting the data set as described allows to net out possible substantial differences between climate changes of the two sub-regions. Second, we assume that within either sub-region no further spatially different (asymmetric)

climate change existed that could lead to a misinterpretation of measured price convergence as market integration.

5 Price convergence and the Great Moderation of Grain Price Volatility

We present our main results as two stylized facts. First, price convergence between 15 German cities emerged during the period 1651–1790. Second, a substantial reduction of grain price volatility occurred. We discuss robustness checks and extensions in Section 5.3.

5.1 Fact 1: Price convergence before the French Revolution

The cross-sectional SD based on five-year-averages of rye prices exhibits a clear downward trend (Figure 3, panel a). At the national level, prices converged by ca. 0.3% per year in 1651–1790 (0.3% $\approx$ parameter/average SD 1651–75 (= $-0.0009/0.28$). In 1651–75, the SD was roughly 0.28 (CV$\approx$31%); this value declined to 0.16 (CV$\approx$17%) for the years 1766–90.[8] Bateman (2011, 459) instead finds no strong trend of price convergence for Germany. Federico (2011, 102, 114) reports negative but not significant changes for 1750–88 which is consistent with our analysis because most of the convergence appeared from 1651 to 1750. Our results confirm studies pointing towards market integration as a gradual process in Europe starting before the 19th century (e.g., Chilosi et al., 2013).

Our result is not confounded by asymmetric/symmetric weather shocks or spatially symmetric climate change, which leads to absolute price changes. Furthermore, because there is no negative trend in the average real price, which could bias the SD towards price convergence, we can further exclude that spatially *symmetric proportional* price changes driven by climate change affect the result of national price convergence. Note that if we used the CV on annual data, the trend estimate would be substantially more negative (Figure S13 in SA10).

But without further investigation, we cannot exclude effects of *asymmetric absolute/proportional* price changes under *no arbitrage*, induced by climate change on the variation of the SD between the five-year-sub-periods in the complete sample (Figure 3 panel a). A sub-sample with regionally specific climate might have experienced on average better growing conditions for grain or less (severe) shocks and thus, fewer crop failures than the

[8] Results for data at annual frequency and additional discussion are available in the SA. National price convergence is apparent irrespective of whether we use the CV, the SD or the median absolute deviation, a robust measure of dispersion (Figures S12–S15 in SA10).

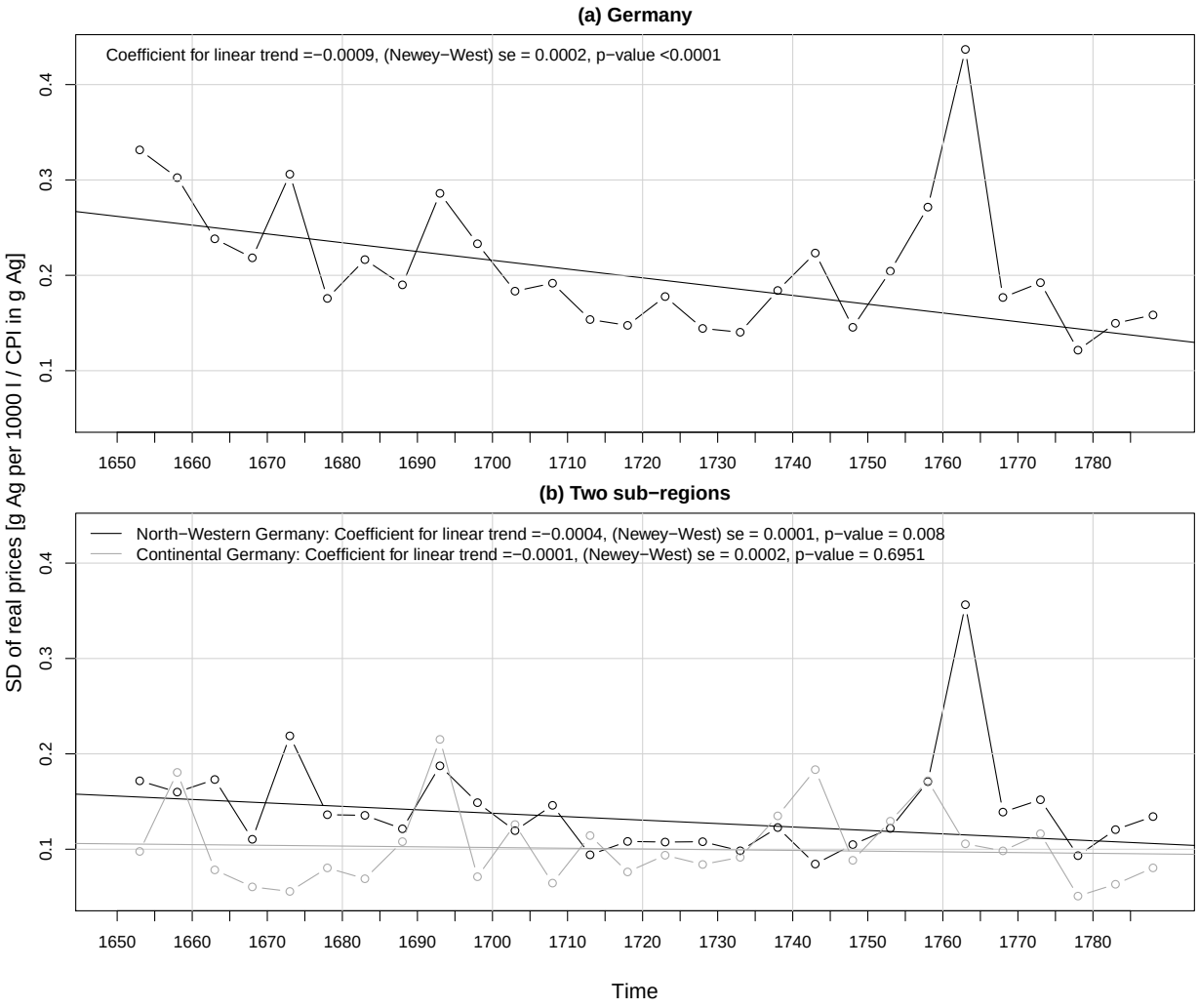

Figure 3: Inter-urban price dispersion 1651/5–1786/90. Cross-sectional standard deviation of real 5-year-mean-prices, rye (stable sample). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

rest of the sample (cases: asymmetric absolute/proportional changes; cf. Section 4.3). Although both regions might not have been integrated in one market, we would measure changes in the inter-urban price dispersion at the national level, because one sub-region experienced a trending real price, affecting the within-region variance. Additionally, price convergence at the national level can be the result of a decrease of the between variance of the sub-regions if no arbitrage takes place between them.

To approach the problem of asymmetric, that is, regional climate change, we produced a set of results for two sub-regions: North-Western and continental Germany. The idea is that an asymmetric change for the complete stable sample, becomes analytically a symmetric

change for either of the sub-regions. The result of decreasing inter-urban price dispersion measured as the within SD[9] remains robust within North-Western Germany but the trend weakens; convergence is not apparent in continental Germany (Figure 3 b).

Furthermore, the regional results reveal that during the crisis of 1741 the within SD increased only in continental Germany (comprising Southern and Eastern Germany), indicating regionally different price effects within this region. This phenomenon might be explained by *Fruchtsperren*, trade restrictions preventing arbitrage (cf. Göttmann, 1991, 93–4). The period of the Seven Years' War showed a higher level of price dispersion in North-Western Germany. However, the data during the Seven Years' War are of doubtful quality because of war-related inflation.

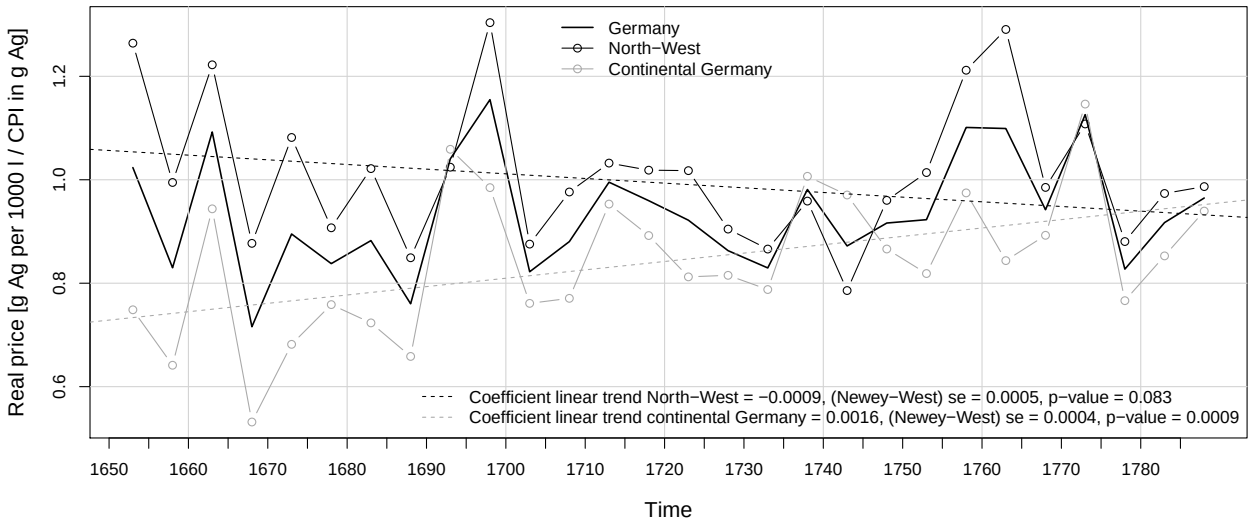

Figure 4: Real 5-year-mean-prices for Germany, North-Western and continental Germany, rye (stable sample) 1651/5–1786/90. Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

A downward trend of the real rye price in North-Western Germany could bias our result towards price convergence. Once we include a dummy variable for the Seven Year's War, we observe a statistically significant downward trend for the real rye price in North-Western Germany; for continental Germany we observe a significant upward trend (Figure 4). The downward trend for North-Western Germany could indicate increasing output and thus decreasing real prices due to regional climate change. On the other hand, the upward trend in

[9]The within SD is defined as the square root of the within variance of each region. This is based on the decomposition proposed by Federico (2011, 125) but the variances are not normalized at each point in time using the average price of the complete market. In this way, we analyze the absolute variance (of real prices) and all results of our formal analysis regarding shocks hold.

continental Germany is inconsistent with more favorable growing conditions due to the end of the LIA (or these are overcompensated by an increase in demand).

Figure 5 plots the between-SD, that is, the square root of the between variance of the two sub-regions. The figure shows that a major fraction of the convergence we observe at the national level occurred between the two regions. The negative trend is much stronger than the price convergence observed within North-Western Germany (Figure 3 b).

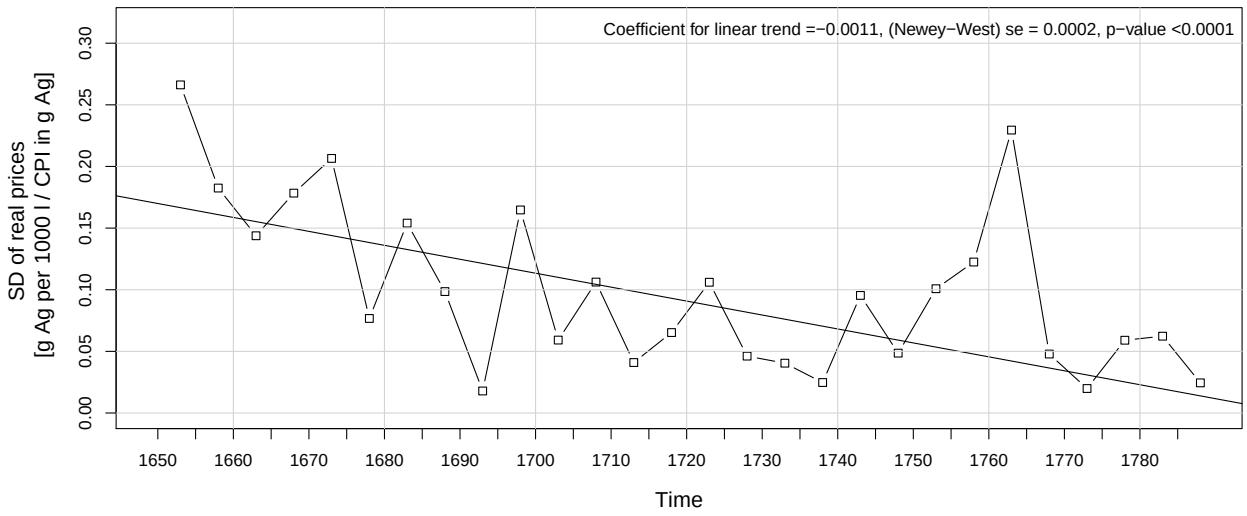

Figure 5: Inter-regional price dispersion 1651/5–1786/90. Between-region standard deviation (square root of variance between North-Western and continental Germany); real 5-year-mean-prices, rye (stable sample). Regression for linear trend includes a dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

As discussed in Section 4.3, a negative trend of the average price in North-Western Germany (potentially caused by warming, the end of the LIA) might lead to convergence of North-Western and South-Eastern prices even when using the SD of five-year mean prices. We test whether the downward trend of the inter-regional price dispersion remains significantly negative after subtracting the negative trend of the North-Western real rye price. For this purpose, we use trends from regressions where the dependent variables are in logs so that the coefficients for linear trends are semi-elasticities and thus comparable. Particularly, we test whether -0.0119 [trend parameter inter-regional price dispersion] $-(-0.0008)$ [trend parameter real rye price Northern Germany] $= 0$ ($p < 0.001$; Newey and West standard error).[10] Thus, the trend of the inter-regional dispersion remains substantially negative in magnitude, -0.0111, and is significantly different from zero. That is, inter-regional dispersion decreases by 1.11% per year net of price convergence due to the decreasing real price

[10]The trend counts years not five-year-periods where the parameter would be five times the magnitude.

in North-Western Germany. This test is directly related to our formal analysis, because we show that, *without arbitrage*, a *symmetric proportional shock* with a negative shock rate remains as a factor that decreases the SD (see SA15, eq. S12). A similar test reveals that the negative trend estimate for price dispersion within North-Western Germany remains significant with $p = 0.07$.

The logic of our methodology is to rule out that climate change can affect our measurement of price convergence by exclusion. The results obtained with the sample split warn against the interpretation that national price convergence signals national market integration—this is a notable result given current practice in the literature. Price convergence between regions could be a result of asymmetric climate change. In our particular case, we cannot rule out that a small part of the inter-regional price convergence stems from regional climate change in North-Western Germany. But the size of the downward trend of the North-Western real rye price cannot explain the magnitude of the observed inter-regional price convergence. We cannot exclude that climate change affected agriculture in continental Germany (a growing population might have overcompensated increased output due to the End of the LIA) but we can exclude that measurement is biased towards price convergence, that is, towards finding market integration.

On the one hand these results confirm the approach by Chilosi et al. (2013) who identify sub-regions where prices converged. On the other hand, our results call into question approaches using prices of cities from all over Europe. Europe's climate is more diverse than the one in Germany.

In short, our results show a process of market integration in seventeenth- and eighteenth-century Germany. At the national level, price convergence is robust to weather shocks and climate change resulting in symmetric absolute price changes. Once we allow for two sub-regions to control for spatially different effects of climate change, the results show that no price convergence is observed within continental Germany. The decrease of price dispersion within North-Western Germany is weakened by the downward trend of this sub-region's average price. However, the latter is not large enough to explain all of the observed price convergence within the North-West. Convergence between the two regions accounts for a major fraction of national convergence. After netting out the effect of the downward trend of the average price in North-Western Germany, still a substantial decline of the cross-sectional standard deviation of five-year-average-prices by 1.11% per year remains, which suggests that some market integration took place beyond North-Western Germany. We explore the nature of this phenomenon below in Section 6.

5.2 Fact 2: The Great Moderation of Grain Price Volatility

The leveling-out of price peaks during the first part of the eighteenth century discussed in the data section implies a reduction of price volatility. Figure 6 shows what we call, in analogy to 'The Great Moderation' of volatility in output and inflation during the 1980s to early 2000s (e.g., Summers, 2005), *The Great Moderation of Grain Price Volatility*: Aggregate volatility (calculated as the CV over time)[11] decreased substantially over time until the French Revolution as illustrated by the significant negative linear trend. The coefficient of the trend counts years and indicates that in ten years volatility is reduced by roughly one percentage point.

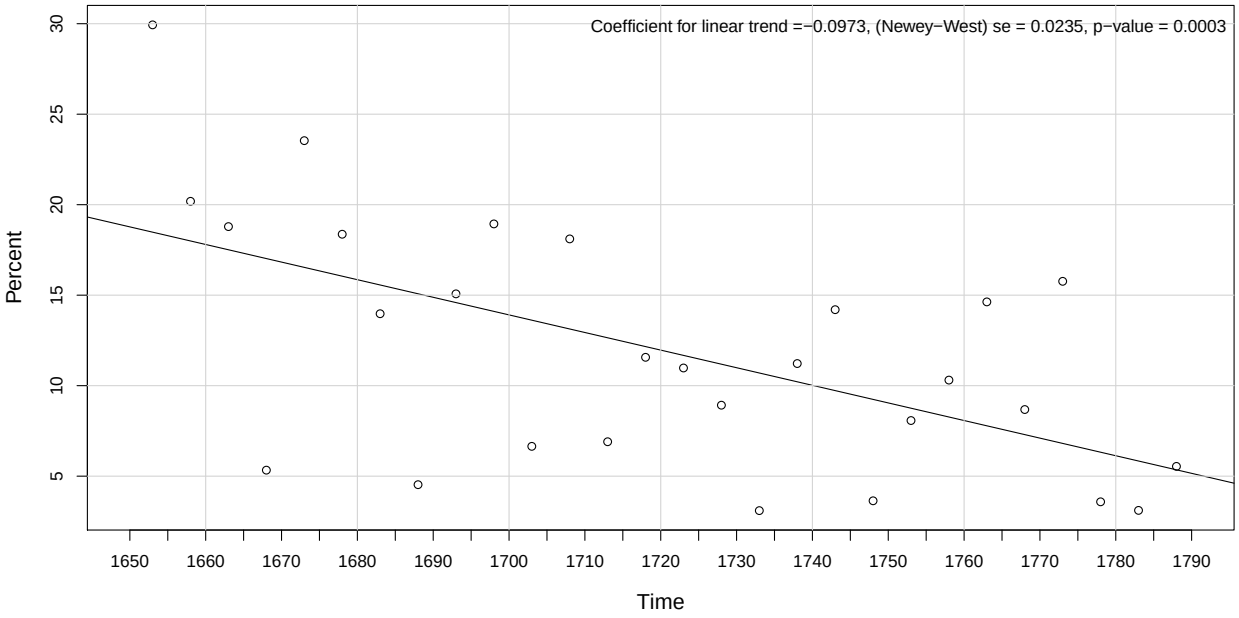

Figure 6: The Great Moderation of Grain Price Volatility 1651/5–1786/90. Volatility of the aggregate real rye price in Germany (stable sample). Each circle represents a five-year-period centered at the given year (e.g., the circle for year 1653 represents the period 1651–55). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

The pattern visible in Figure 6 is confirmed by a panel regression of the five-year volatilities of the individual city price series on a linear trend (with city fixed effects; the parameter is slightly smaller: 0.07; Table S30).[12] The decline of volatility prior to the Napoleonic Wars

[11] Volatility is a measure of relative variability; the CV is one standard method to quantify volatility, the other one is the standard deviation of log differences (Albers et al., 2017, 55; Bateman, 2011, 455; Jacks et al., 2011, 805).

[12] Note that the moderation of volatility is restricted to the 140 years until the Napoleonic Wars—the first half of the 19th century witnessed a new surge in price volatility. The Napoleonic Wars (1792–1815) led to war finance driven inflation that might not be correctly accounted for in currency conversions and real war related shocks (see also Jacks et al., 2011, 807).

is not confined to Germany. Related work on wheat prices in Europe by Chilosi et al. (2013, 51, figure 4) shows a reduction of wheat price volatility in the 18th century (cf. Bateman, 2011, 455–6, 459 who finds only weak evidence of decreasing volatility in Europe and none for Germany).

An important question is whether the decline of aggregate price volatility can be linked to fact 1, price convergence. Integrated markets are expected to have a lower price volatility (Chilosi et al., 2013, 48). But a reduction of the frequency and severity of asymmetric or symmetric weather shocks due to the end of the LIA, particularly following the end of the Maunder Minimum at the beginning of the eighteenth century, could also have reduced price volatility via a reduction of the fluctuation of cereal yields.[13] In Appendix A (with further details in SA16) we show that more spatial arbitrage (and a smaller magnitude and a lower degree of symmetry of shocks) must reduce *aggregate* price volatility. Hence, while we cannot exclude that climate change plays a role, at least a part of the volatility moderation is a result of spatial arbitrage, that is, market integration.

Of course, the quantitative importance of the two factors, changing shock patterns vs. market integration is very relevant. We cannot provide a definite answer but merely a back-of-the-envelope calculation, which makes use of the results of a long-run sample 1576–1790 and our theoretical analysis (all details in SA14). Our tentative estimate is that roughly half of the reduction in price volatility can be attributed to market integration.

5.3 Robustness

We performed two robustness checks using sample variation. First, we dropped cities from our sample. Given the modest size of our sample one might argue that our results are sensitive to including particular cities or combinations of them. In other words, geographical proximity of particular cities forming a common market may affect the result. To investigate the importance of this argument, we systematically tested all $\binom{n}{k}$ possible combinations of dropping $k = 2$ out of the $n = 15$ cities from the stable sample 1651–1790. The trend estimates for the SD of five-year-prices at the national level is between -0.0013 and -0.0006 (stable sample: -0.0009) and always significant at $p < 0.05$. Similarly, the trend of the inter-regional SD also remains negative between -0.0014 and -0.0009 (stable sample: -0.0013) and statistically significant at $p < 0.01$. The trend estimate for the within SD of North-Western Germany is not significantly different from zero in 13 out of 105 cases.

For aggregate volatility, the trend estimates range between -0.1150 and -0.0827 (stable sample:-0.0973). The estimates are significant at least at the 1% level.

[13] Even sole changes of the mean state of climate can lead to different fluctuations of grain output, because yields are non-linearly related to temperature with changing signs.

In our second robustness check, we carried out the tests of our analysis for an unbalanced sample of 29 cities for the period 1651–1790 (see SA9). The results remain very similar. In addition, continental Germany shows price convergence of similar size as North-Western Germany but the parameter is not significant at conventional levels ($p = 0.15$).

Overall, robustness checks using sample variation fully confirm the results of sections 5.1 and 5.2.

6 Market integration along major rivers and across all cereals

In this section we extend the main results in two ways. First, we explore the causes of the decrease of the variance between North-Western and continental Germany in the baseline result. Second, we replicate the main steps of the above analysis for the three other cereals barley, oats and wheat.

While North-Western Germany is a relatively homogenous area with respect to climate, geography and distance between cities, particularly our second region, continental Germany, is more diverse. Furthermore, the large decrease in the variance between North-Western and continental Germany observed in our baseline result above awaits further explanation, particularly because overland transport costs were high even for distances of roughly 20 kilometers (Jacks, 2004, 302). Thus, in what follows we test whether price gaps fell stronger between towns located in the same river system than among other city pairs. If this holds, the reduction of the cost of river trade may have contributed to the integration of grain markets in large areas (cf. Chilosi et al., 2013, 47).

The map in the data section (Figure 1) shows two major rivers in Germany, the Elbe and the Rhine, flowing roughly from South-East to North-West. To test the hypothesis that falling trade costs on navigable rivers contributed to price convergence, we conduct a more detailed variance decomposition where we introduce four sub-regions in total, two of which again form an aggregate region each. In this way, we add a further layer to the variance decomposition as proposed by Federico (2011, 125). Additionally, we exploit as much information as possible by using all available 29 price series.

The first sub-region consists of cities connected by the Elbe and its tributaries, namely, Hamburg, Lüneburg, Dresden, Halle and Berlin. The remaining cities that are located in North-Eastern Germany but have no navigable access to the Elbe river system form what we call the 'North-East'.[14] Cologne, Xanten, Speyer, Strasbourg, Trier, Frankfurt and Würzburg form the Rhine region. These cities are located along the rivers Rhine, Main and Moselle.

[14] Braunschweig, Celle, Emden, Göttingen, Hannover, Minden, Münster, Osnabrück, Paderborn, Leipzig, Gdansk, Königsberg, Quedlinburg.

The fourth sub-region is called 'South-West' and includes cities in South-Western Germany that are not part of the Rhine region.[15]

The decomposition carried out in this way shows that the variance within both the Rhine region and the Elbe region is decreasing until about 1715/20, very much in line with the pattern of national convergence (see Figures S16 and S17). By contrast, the variances within North-East and within South-West show no statistically significant downward trend. This is consistent with decreasing trade costs along the main river systems. Admittedly, what exactly led to a decrease of trade costs must be left for future empirical research. Potential explanatory factors include shipping technology (Chilosi et al., 2013) but also institutional changes impacting on tariffs or increasing competition in the transport sector.

The variance between the first aggregate region (Elbe region and the remaining North-Eastern cites) and the second aggregate region (Rhine region and remaining South-Western cities) still shows a significant downward trend (dummy variables for the crisis of the 1690s and the Seven Years' War included). But the magnitude of the trend of this between variance is much smaller than in the case of the baseline specification with regions defined on the basis of their climate using the entire unbalanced sample (-0.0008 vs. -0.0004, compare Figures S9 and S17). Moreover, an additional variance decomposition demonstrates that the convergence between the large aggregates (Elbe plus North-East vs. Rhine plus South-West) is by and large due to price conversion within North-Western Germany, which reflects the baseline result (SA13). The other components of the overall between variance (between Rhine region and South-West; between Elbe region and North-East) both show a significant downward trend, which indicates that the river regions integrated with geographically close cities.

The second extension analyzes in how far the baseline results can be replicated for other cereals. The potential for market integration is higher for goods exhibiting high value-to-bulk ratios, and grains differed with respect to the value-to-bulk ratio: In 1716–25, the price in grams of silver per litre was 0.43 for wheat, 0.32 for rye, 0.25 for barley and 0.16 for oats (mean of aggregate price, unbalanced sample). Furthermore, whereas rye, together with wheat, was cultivated as a winter cereal, oats constituted a spring cereal; for barley both spring and winter types were possible but spring barley was quantitatively more important (Göttmann, 2006). Spring cereals were sown only after winter and their output fluctuations may thus have reacted to weather shocks and/or climate change in other ways than rye. Results for the main indicators, that is, the cross-sectional standard deviation and volatility of prices are shown in Table 3.

We find price convergence also for all cereals other than rye at the national level, within

[15] Aachen, Augsburg, Munich, Nuremberg.

north-western Germany, between this region and the continental parts of Germany, and to a weaker extent within continental Germany. This fully replicates the findings reported above; market integration was not restricted to rye, but involved all cereals. The pace of market integration was correlated with the value-to-bulk ratio, with wheat experiencing a faster decline than other cereals fetching lower prices. However, this tendency is weak, and oats does not conform to the expected pattern (see column 'in percent' displaying the annual rate of change of the cross-sectional SD in Table 3).

Table 3: Trend estimates for cross-sectional standard deviation of five-year-average-prices and volatility of aggregate price, four cereals 1651–1790

Cereal, decreasing value-to-bulk ratio	Annual trend of standard deviation		within North-West	within continental	between North-West and continental	Annual trend of volatility of aggregate price
	National					
	absolute	in percent				
Wheat						
St., $N = 11$	-0.0015^{***}	$\widehat{=} -0.40\%$	-0.0008^{***}	-0.0003^{***}	-0.0016^{***}	-0.0534^{***}
Unb., $N = 27$	-0.0015^{***}	$\widehat{=} -0.40\%$	-0.0009^{***}	-0.0004^{*}	-0.0012^{***}	-0.0513^{***}
Rye						
St., $N = 15$	-0.0009^{***}	$\widehat{=} -0.33\%$	-0.0004^{***}	-0.0001	-0.0011^{***}	-0.0973^{***}
Unb., $N = 29$	-0.0010^{***}	$\widehat{=} -0.38\%$	-0.0004^{***}	-0.0003	-0.0008^{***}	-0.0921^{***}
Barley						
St., $N = 10$	-0.0008^{***}	$\widehat{=} -0.35\%$	-0.0003^{***}	-0.0001	-0.0009^{***}	-0.0464
Unb., $N = 26$	-0.0007^{***}	$\widehat{=} -0.35\%$	-0.0004^{***}	-0.0003^{**}	-0.0005^{***}	-0.0511^{*}
Oats						
St., $N = 9$	-0.0007^{***}	$\widehat{=} -0.46\%$	-0.0003^{***}	-0.0001	-0.0009^{***}	-0.0134
Unb., $N = 26$	-0.0006^{***}	$\widehat{=} -0.46\%$	-0.0003^{***}	-0.0003^{***}	-0.0003^{***}	-0.0082

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$; based on Newey-West standard errors. 'St.' refers to stable sample ($\leq$5% missing observations per individual series); 'unb.' refers to unbalanced sample using all available cities. N is the number of included cities. Own calculations based on data described in SA2. Annual rate of convergence at the national level in percent is calculated relative to the average of the five cross-sectional standard deviations of five-year-average-prices for 1651–1675. Volatility is measured as the coefficient of variation over time. Results for rye as discussed in main results and included for comparison.

Price volatility also declined among all three grains other than rye, but the trend is not significant at conventional levels for barley in the case of the stable sample ($p = 0.12$) and insignificant for both specifications in the case of oats. This may be due to the fact that the initial level of volatility and, hence, the absolute magnitude of the subsequent decline, was much smaller in the case of oats compared to rye (15% vs. 26% in 1651–1675, and 9% vs. 14% in 1766–90). These differences in volatility levels should not be taken as an indicator of a higher integration of markets for oats compared to rye. Rather, they may reflect differences with respect to the reaction and/or exposure to weather shocks. As mentioned above, rye is a winter cereal, oats a spring cereal and thus, the latter is not exposed to shocks in late autumn or winter.

Overall, the results in Table 3 corroborate our earlier results. All cereals exhibit price

convergence and all show some reduction of volatility until 1790. To be sure, results vary among the four grains considered here, but the differences do not suggest a regular pattern. At the bottom line we can say that market integration was not restricted to rye but involved all segments of grain markets between the mid-seventeenth and the late eighteenth century.

7 Demographic implications of lower food price risk

The 'Great Moderation of Grain Price Volatility' matters economically via the impact of grain prices on mortality. Food prices are indirectly related to empirical specifications of Malthusian growth models, because it is usually assumed that the positive check operates through nutrition (among other channels; Møller and Sharp, 2014, 110). Due to a high share of bread from grain in the typical consumer basket in pre-industrial Germany and rare opportunities for substitution with other sources of calories such as potatoes[16], grain price spikes had the potential to push a household's budget below the poverty line. Thus, the correlation of the aggregate rye price with aggregate mortality can inform us about the role of food prices for the positive check in Germany.

In our descriptive analysis of the aggregate price we noted coincident spikes in demographic variables. In fact, the aggregate real rye price and the German crude death rate (CDR) (see also Figure S26) exhibit a strong short-term relationship during the period 1730–1790, where overlapping and reliable mortality data are available from Pfister and Fertig (2010). We considered a dynamic specification in an exploratory bivariate vector autoregressive (VAR) model, because one would typically assume that there is a considerable lag between a hunger crisis and actual death.[17] Furthermore, increasing mortality might lead to lower demand and thus lower prices. The VAR model of the log-differences of the crude death rate and the real rye price shows that both variables Granger cause each other ($p < 0.01$, based on Newey-West standard errors; model with 2 lags of the price and 4 lags of the CDR; includes dummy variable for Seven Years' War and winter precipitation, a significant predictor of rye prices). An increase of the first lag of the rye price is related to a 0.71% increase of the CDR (0.49% with additional dummy variable for 1772; $p < 0.01$).[18] Vice versa, the second and fourth lag of the CDR are negatively related to the rye price with -0.27% and -0.26%, respectively (-0.29% and -0.27%; $p < 0.05$), which is in line with the

[16]Potatoes were of minor importance in Germany prior to 1800. The cropping share of potatoes was only roughly 2% around 1800 (van Zanden, 1999, 368, table 16.6); at best about 8% of consumed calories were provided by potatoes (Pfister, 2017, S2, p. 3).

[17]The contemporaneous growth rates (log-differences) of the death rate and the aggregate price are significantly positively related: a 1% increase of the rye price is associated with a 0.34% increase of the death rate (1730–1790; $p < 0.05$, $R^2 = 0.19$; dummy variable for Seven Years' War included).

[18]In 1772, the highest CDR of 48 deaths per 1000 and a very high real price are observed.

intuition that a reduction of population size should decrease demand for grain (the third lag is positive but smaller: 0.18%). While such a model does not constitute an accurate quantification of Malthusian mechanisms, the positive short-term relationship between rye price and mortality can be regarded as indirect empirical evidence for the severe consequences of grain price variations on household budgets in a Malthusian growth regime. The strength of the relationship is notable in comparison to England, where Møller and Sharp (2014, 126, 129) find no significant support for the positive check.

Survival probability increases *ceteris paribus*, if price peaks wane and food prices become more stable, as it was the case during the Great Moderation of Grain Price Volatility. Intuitively, the real budget of a consumer falls short of the required minimum for survival less often and less deeply. One might conjecture that market integration dampened the impact food price spikes on mortality. Although we can appeal to correlations for a short period only, the death rate also becomes less volatile over time once the mortality peak in 1756–1763 is attributed to the Seven Years' War (Figure S26, right axis). The main exception to the pattern of decreasing mortality is the outstanding subsistence crisis around 1770: either the crop shortage was so highly correlated across space that markets became illiquid and/or disintegrated, e.g., as a result of trade bans. Thus, the moderation of grain price volatility, at least partly a result of spatial arbitrage, correlated with a decline in mortality shocks. But this should not lead to the interpretation that the emergence of grain markets already provided perfect insurance of households against the consequences of price shocks. The crisis of 1771/72 shows that still a fair degree of demographic vulnerability remained.

While we have focused mostly on the short-term relationship with mortality so far, long-term consequences of increased survival probability are possible as well. Via increasing survival probability, the decline of grain price volatility can be linked to the model by Kögel and Prskawetz (2001) within the framework of unified growth theory (Galor, 2011). This model has a crucial exogenous element, the increase in survival probability of children, which is relevant for the transition to the modern growth regime. Higher survival probability reduces precautionary demand for children and simultaneously increases human capital investments (Kögel and Prskawetz, 2001, 348–9).

In sum, the joint reduction of the volatility of grain prices and mortality indicates that the consequences of agricultural market integration were potentially not restricted to Smithian growth. First, the positive check might have become less severe and second, higher survival probability might have reduced precautionary demand of children. In the absence of a quantitatively important demand for human capital in the eighteenth century, resources might have been redirected towards child-bearing in a low(er) mortality regime; after falling for a considerable time, the fertility rate stabilizes after 1750. An exact quantification of the

demographic consequences is beyond the scope of this paper; first, more data work is needed to derive reliable aggregate vital rates prior 1730.

8 Conclusion

This study addresses the issue of measuring changes in grain market integration using prices from a sample of markets in a Malthusian situation where important economic variables are stationary and cointegration techniques cannot be applied. Furthermore, asymmetric weather shocks and potentially climate change in the form of the end of the Little Ice Age impact on the behavior of prices and give rise to a measurement problem: Is price convergence a result of market integration or of changing shock patterns? We demonstrate that the cross-sectional standard deviation of five-year averages of real grain prices is robust to spatially symmetric and asymmetric weather shocks. To exclude further potential bias in the measurement of market integration that results from climate change, we test whether the observed negative trends of within-region and inter-regional price dispersion are sensitive to downward trends of regional mean prices.

We apply this methodology to a new data set of German grain prices in 1651–1790 for a stable sample of 15 towns and an unbalanced sample of additional 14 towns. In contrast to earlier historical work on market integration with regard to wheat we focus on rye, which dominated both grain farming and food trade in Germany prior to c. 1800. During this period Germany was an inland region that benefited little from the development of Atlantic trade (e.g., Acemoglu et al., 2005). Four findings emerge. First, we find unequivocal support for market integration within North-Western Germany during the period 1651–1790. This finding is also robust to climate change associated with the Little Ice Age, e.g., the observed contemporaneous increase in average spring temperature and the potential change of short-term supply shock patterns it might have engendered. Even if warming led to higher agricultural productivity and thus a lower average price in North-Western Germany, the observed downward trend of the North-Western average price does not explain entirely the decrease of cross-sectional price dispersion in the North-West.

Second, the average price gap between the North-Western markets and those located in continental Germany fell faster than the real price of rye in the North-West. This indicates that although both sub-regions belong to different climatic regions measured by their continentality, any potential warming in the North-West cannot explain the size of the observed inter-regional price convergence. Further analysis demonstrates that this process was concentrated on cities located in two important river systems, Elbe and Rhine. These rivers provided navigable trade connections between the continental parts of Germany and the

North-West. Thus, Elbe and Rhine provided a basis for linking inland regions with the North-Western core even with respect to trade with goods characterized by a low value-to-bulk ratio such as rye.

Third, market integration was not restricted to rye but was present in all segments of the grain market, that is, also in the cases of barley, oats and wheat. There is a weak tendency of a relatively faster integration for crops with a higher value-to-bulk ratio such as wheat during the period under study.

Finally, we observe a drastic moderation of grain price volatility. Whereas this may have been partly caused by climate change—particularly the end of the Maunder Minimum—and its effect on the behavior of supply shocks, we demonstrate analytically that spatial arbitrage reduces *aggregate* grain price volatility. Thus, what we term the 'Great Moderation of Grain Price Volatility' is at least partly a result of market integration. Quantitatively, both factors climate change and market integration seem to have been of similar importance.

In sum, these results contribute to the literature on the history of grain market integration and long-run economic development. Methodologically, we clarify in how far the widely used coefficient of variation is affected by shocks and produces misleading conclusions about market integration. Furthermore, we prove that lower aggregate volatility can be linked to price convergence (but also to weaker shocks), something the literature has either tacitly taken for granted or criticized. Empirically, previous research has found little evidence so far for an advance of market integration in Central Europe between 1650 and 1800, a view that our study allows to overcome. The extent of markets improved significantly before the railway age. Market integration had at least one important consequence, which we think is difficult to challenge, namely, Smithian growth. The timing of price convergence coincides with more regional specialization found by earlier research and a non-stationary German real wage since the late 17th century (Pfister, 2017, 723). Møller and Sharp (2014, 118) argue that Smithian (or related Boserupian) growth is one plausible mechanism that can ease or ultimately overcome the constraints of a Malthusian setting.

In addition, our evidence on the sizable relationship between aggregate rye price and death rate points at a strong positive check in Germany compared to England so that Smithian growth was possibly more limited in the German case. At the same time, the positive relationship between the rye price and the death rate indicates an important role of food price fluctuations in the positive check. But Malthusian pressure was not necessarily constant; the decline of price volatility is also mirrored in a decreasing volatility of mortality. Due to a lack of reliable data prior to 1730, this result must remain tentative but suggests that agricultural market integration could have had two further consequences beyond Smithian growth working via food prices. First, the positive check might have become

weaker in Germany over time. Second, survival probability likely increased during the Great Moderation of Grain Price Volatility, which points at an argument from the unified growth literature (Kögel and Prskawetz, 2001): the reduction of precautionary demand for children due to higher survival probability.

A Appendix: spatial arbitrage reduces aggregate price volatility

In what follows, we analyze how volatility reacts to shocks depending on the type of shock. We focus on absolute shocks. All details and an extension to the case of proportional shocks is relegated to SA16.

Volatility V is defined as the coefficient of variation (CV) calculated for a time series: $V = \frac{1}{\bar{p}}\sqrt{\frac{1}{T-1}\sum_{t=1}^{T}(\bar{p}_t - \bar{p})^2}$, where index $t = 1, ..., T$ counts years. The mean over time is denoted with $\bar{p} = \frac{1}{T}\sum_{t=1}^{T}\bar{p}_t$. The single $\bar{p}_t$ within the summation operator correspond to the cross-sectional average prices in year t in the analysis of the cross-sectional CV (and to the aggregate price in Figure 2).

We now introduce a positive asymmetric absolute shock s_{it} in city 1 in year 2: $s_{12} > 0$. We introduce the factor $a = \frac{1}{N}$, $0 < a \leq 1$, where N is the number of cities in the cross-section as before. Thus, a measures the potential for arbitrage. A large a indicates that less cities comprise one common market and participate in arbitrage. In other words, the larger a, the less cities can dampen the local shock. Note that this precludes prohibitively high trade costs. We assume that an additional city is indeed available for cross-sectional arbitrage. The case $a = 1$ means that there is no cross-sectional arbitrage.

Effect on mean

The average price over time including the shock is $\bar{p}^z$:

$$\bar{p}^z = \frac{\bar{p}_1 + [\bar{p}_2 + as_{12}] + \bar{p}_3 + ... + \bar{p}_T}{T} = \frac{\bar{p}_1 + \bar{p}_2 + \bar{p}_3 + ... + \bar{p}_T}{T} + \frac{as_{12}}{T} = \bar{p} + \frac{as_{12}}{T}. \quad \text{(A1)}$$

Effect on sum of squared deviations

We focus on the sum of squared deviations (SSD), because the standard deviation (SD) (the numerator of the CV) is a monotonic transformation (Simon and Blume, 1994, 497–8) of the SSD (and the variance). This allows using the sign of the derivatives of the SSD instead of the SD, which simplifies the calculation. The SSD including the shock are defined as the function $u(.)$:

$$
\begin{aligned}
u(.) &= \sum_{t=1}^{T} (\bar{p}_t - \bar{p}^z)^2 \\
&= (\bar{p}_1 - \bar{p} - \frac{as_{12}}{T})^2 \quad + (\bar{p}_2 + as_{12} - \bar{p} - \frac{as_{12}}{T})^2 + (\bar{p}_3 - \bar{p} - \frac{as_{12}}{T})^2 + ... \\
&+ (\bar{p}_T - \bar{p} - \frac{as_{12}}{T})^2 \\
&= ... \qquad\qquad\qquad + (\bar{p}_2 - \bar{p} + \frac{(Ta - a)s_{12}}{T})^2 + ...
\end{aligned}
\tag{A2}
$$

To evaluate the effect of the shock we now assume $\bar{p}_1 = \bar{p}_2 = \bar{p}_3 = ... = \bar{p}_T$. This assumption means that the time series of cross-sectional average prices is a 'flat line' except for the shock we introduce in $t = 2$. In SA16 we illustrate the 'flat line' assumption (Figure S25) and provide further discussion. If all cross-sectional average prices without the shock are equivalent, this implies that their average over time (also without any shock) is: $\bar{p} = \bar{p}_1 = \bar{p}_2 = \bar{p}_3 = ... = \bar{p}_T$. The squared deviations for periods where no shock occurs (here $t = 1$ and $t = 3$) are only influenced through the altered mean price, that is, the mean price $\bar{p}^z$ including the shock s_{12}.[19]

Due to the 'flat line' assumption, we find the same expression for the squared deviations $T - 1$ times (only the expression for the period with shock is different). Eq. A2 simplifies to (details in eq. S17):

$$
u(.) = (T - 1)\frac{a^2 s_{12}^2}{T}. \tag{A3}
$$

The expression for $u(.)$ in eq. A3 is clearly positive, because $T > 1$. Based on this result and further analysis, SA16.1.3 shows that a shock increases volatility.

The role of spatial arbitrage can be analyzed by taking the first derivative of $\frac{u(.)}{\bar{p}^z}$ with respect to a:

$$
\frac{\partial u/\bar{p}^z}{\partial a} = \frac{2(T-1)\frac{as_{12}^2}{T} \cdot [\bar{p} + \frac{as_{12}}{T}] - [(T-1)\frac{a^2 s_{12}^2}{T}] \cdot \frac{s_{12}}{T}}{[\bar{p} + \frac{as_{12}}{T}]^2} > 0, \tag{A4}
$$

because the numerator can be simplified to: $2(T-1)\frac{as_{12}^2}{T} \cdot \bar{p} + (T-1)\frac{a^2 s_{12}^3}{T^2} > 0$, because $T > 1$ (see SA16.1.4). Thus, a larger number of cities N, which are available for spatial arbitrage, means that $a = \frac{1}{N}$ decreases and reduces volatility. In SA16.2, we show that this result holds also for proportional shocks.

The factor a has a second intuitive interpretation. Let s_{12} denote a shock that is not only limited to city 1: $s_{12} = s_2$. The larger a, the 'more symmetric' is the shock s_2. We could define $a = \frac{n}{N}$, where $n \leq N$ is the number of cities experiencing the shock. As before,

[19] To simplify this analysis, we assume no intertemporal arbitrage, e.g., through storage (which would also reduce volatility).

a larger a means a larger volatility; $a = 1$ (implying $n = N$) corresponds to the case of a perfectly symmetric shock.

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Supplementary Appendix

SA1 Overview of data preparation

The data series underlying this study are annual grain prices in grams of silver per litre. Original data are in different and highly variable units, however, which requires several steps of preparing the data for analysis. Specifically, we use published compilations that report grain prices in historical volume units and currencies that vary between markets and sometimes over time as well. Original data also differ with respect to time reference and frequency. Wherever possible we employ retail prices because we expect that these reflect local market conditions. Where we draw on a mixture between retail and wholesale prices we adjust the latter to the level of retail prices (e.g., Hamburg).

Section SA1.1 explains the principles followed in converting original prices into grams of silver per litre. Section SA1.2 outlines the standardization of the time reference to the calendar year. The final part of this section describes the procedures applied in interpolating and extrapolating missing data (Section SA1.3).

SA1.1 Conversion to grams of silver per litre

Prices are reported in silver based currencies for the majority of German speaking cities. Hence, we converted all prices to grams of silver per litre (g Ag per l) via the fine metal content. For this purpose we apply the following basic relationships: 1 *Mark* of Cologne = 233.8555 g Ag (Rittmann, 1975, 535–7).

Following the Vienna treaty of 1857 most German states shifted to a new system in 1858 where 1 Mark equaled 500 grams of fine silver (Zich, 2009, 126). The resulting difference in the silver content of regional currencies was small, however (Statistisches Reichsamt, 1935, 310). In a few cases, prices are reported in money of account (which has no metal content; Gerhard and Engel, 2006, 40–6, 59; Metz, 1990). We converted money of account to silver using exchange rates with gold currencies (*Rheinischer Gulden*) and gold-silver ratios.

One limitation of using silver contents remains. Our information on metallic content usually refers to a time point previous to an inflationary episode. Thus, our series overestimate silver price inflation during a period of currency debasement. To the extent that the intensity of currency debasement differed across towns and territories, inter-urban price dispersion is overestimated. Three periods of intensive currency debasement, during which the quality of our silver price information is highly doubtful, stand out: the *Kipper and Wipper* era at the

beginning of the Thirty Years War (1620–1623; Kindleberger, 1991), the Seven Years' War (1756–1763; Denzel and Gerhard, 2005, 169–76) and the Napoleonic Wars (c. 1799–1815). Following Denzel and Gerhard, during the Seven Years' War, prices in Hamburg were not affected by currency debasement whereas exactly this was happening in Lower Saxony, leading to a divergence of prices in Northern Germany. Price dispersion would also be overestimated in this case.

SA1.2 Calendar year average prices

The predominant part of the data for the stable sample 1651–1790 is obtained as annual calendar year average from the literature (Table S1). However, two forms of data require a transformation of the original series: crop year and *Martini* prices.

Table S1: Overview of time reference of raw data

Stable sample 1651–1790				Additional cities in unbalanced sample 1651–1790			
City	calendar	crop	Martini	City	calendar	crop	Martini
Berlin	x			Aachen	x		
Braunschweig	x			Celle	x		
Cologne	x			Emden	x		
Dresden	x			Königsberg	x		
Gdansk	x			Nuremberg	x		
Hamburg	x			Quedlinburg	x		
Xanten	x			Strasbourg	x		
				Trier	x		
Augsburg	x	x		Leipzig		x	
Munich	x	x		Frankfurt a. M.		x	
Würzburg	x	x		Speyer		x	
Göttingen	x		x	Hanover	x		x
Halle	x		x	Lüneburg	x		x
Münster			x	Minden	x		x
Osnabrück			x				
Paderborn			x				

Note: Symbol x denotes that time base in column is used in underlying data. E.g., the time base for Berlin is the calendar year. The final series for Augsburg is based on both calendar year and crop year prices which were converted to calendar year prices. a. M.: am Main.

Martini prices refer to average prices during four to twelve weeks around Martinmas (November 11). They were recorded because they frequently served as basis for the monetization of peasant dues (Gerhard and Kaufhold, 1990, 396; see also Elsas, 1933, 228). We refer to these prices as average prices for November and December. *Martini* prices cannot be compared with calendar year averages. Hence, we developed a method to extrapolate calendar year prices from *Martini* prices; details are relegated to Section SA3. We dealt with crop year prices in a similar way. Crop year prices refer to a shifted period such as August 1501 to July 1502 (details in Section SA3).

There are three reasons for using all data in calendar years rather than crop years or *Martini* prices. (i) Harvest dates vary by region and agricultural commodity. Defining the precise dates for the crop years is rather difficult, as already Elsas admitted (Elsas, 1933, 224–5; 1936, 92–3). Thus, defining a period always constitutes an average over the regional and commodity specific differences. (ii) It is reasonable to choose a time base that renders possible international comparisons and comparisons with other variables, e.g., nominal wage data. (iii) To keep the data preparation as lean as possible it is reasonable to normalize data to calendar years since these constitute the reference for the bulk of the available data.

SA1.3 Extrapolation and Interpolation

The following two data preparation rules summarize how we deal with extrapolation and interpolation.

1. To preserve local information, we preferred the local relationship for extrapolation from *Martini* or crop year prices. Results are shown in Section SA3 in Tables S6–S11 (*Martini* prices) and S16–S17 (crop year prices). If estimation of a local relationship was no possible, we extrapolated calendar year prices on the basis of panel data regressions for other cities (*Martini* prices: Celle, Osnabrück; crop year prices: Augsburg, Frankfurt, Leipzig, Speyer, Würzburg). The rule for *Martini* prices is given in eq. S4; the one for crop year prices in eq. S8. The commodity-specific parameters for crop year conversions are in Table S18.

2. Single missing data points in series relating to *Martini* and crop year prices were interpolated with the mean of adjacent years for the stable sample 1651–1790 and the additional period 1791–1850 (extension of stable sample to 19th century) before the conversion to calendar years. This procedure avoids the loss of information due to the fact that each single calendar year price is obtained using up to three neighboring crop year (two *Martini*) prices. Interpolation was not applied to any data before 1649 (the last lag needed to extrapolate a calendar year price for 1651). Table S2 shows interpolated years for each series. Inclusion of a series into the stable sample requires that the sum of interpolated values and missing observations is less or equal 5% of the number of observations per interpolated series during the period under study. Multiple missing data points were not interpolated. In addition to the series of the stable sample, the unbalanced sample contains all remaining series which have more missing observations.

Table S2: Interpolated years in data set

Augsburg			Frankfurt				Leipzig	
barely	oats	rye	barley	oats	rye	wheat	barley	oats
1653	1759	1650	1722	1729	1747	1651	1660	1767
		1759	1735	1745	1755	1660	1670	
			1740	1767	1780	1706	1689	
			1743	1781	1785		1767	
			1772	1786				
			1781					

Leipzig		Munich			Wuerzburg			Speyer
rye	wheat	barley	rye	wheat	oats	rye	wheat	rye
1655	1664	1684	1689	1662	1704	1685	1654	1677
1660	1666	1687		1666	1712	1687	1690	1686
1681	1675			1682	1724	1779	1695	1698
1687	1697				1740	1791	1697	1799
1767	1701				1750		1792	
	1758							

Note: Given years refer to interpolated years in crop year and *Martini* price series which were used for the conversion to calendar year prices. Periods which contain interpolation: 1651–1790 (1651–1850). The shares of interpolated values per existing observations is always $\leq 5\%$ for each period. Interpolation method: mean of neighboring years.

SA2 Description of individual grain price series for pre-industrial Germany

What follows provides the information for each price series from the 15th to the 19th century by city in alphabetical order. Years in brackets indicate the period covered by information on prices. The first paragraph addresses conversion of currencies and volumes. Second, we list the different grain types for the city and their sources. We specify the stage of marketing and quality; the absence of further information indicates that prices are retail prices and refer to the same quality. Third, we describe the rules followed in data preparation wherever such steps were necessary. Therefore, no description of data preparation indicates that there is none. Finally, we list extant data that we did not include and give the reason why we did not consider them. Figures S2–S4 in Section SA4 provide plots of each nominal rye price series which are part of an unbalanced sample for the period 1601–1850.

Linear regressions (ordinary least squares estimator) for the purpose of extrapolation were performed on silver prices for metric units unless indicated otherwise. We report heteroscedasticity and autocorrelation consistent standard errors according to Newey and West

(1987; 1994); the used codes of statistical significance are presented as follows if not indicated otherwise: ***: 0.001, **: 0.01 and *: 0.05. Additionally, we provide information on model fit (R^2). All estimates are rounded to four digits (two digits for R^2).

Aachen

Currency and volume conversion

Conversion to grams silver follows Gerhard and Kaufhold (1990, 57, 416); volume conversion as for Berlin.

Rye, barley, oats, wheat (1784–1871)

Calendar year prices 1784–1819 in *Reichstaler* per *Berliner Scheffel* from Kopsidis (1994, Appendix Table V.a/8), 1820–59 from Geheimes Staatsarchiv Berlin (b), 1824 and 1860-4 from Königliches Statistisches Bureau (1867, 117, 121, 126, 130), 1865–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72).

Crop year prices (July – June) in silver equivalents and metric litres for 1532–1783 are available from Rahlf (1996, 150–6). However, these data are not used, because it is impossible to establish a stable relationship between crop year and calendar year prices which prevented the extrapolation of calendar year prices for this market (cf. section SA3.2.2).

Augsburg

Currency and volume conversion

Basic currency system in Augsburg (e.g., Kruse, 1782, 59): 1 *Gulden* = 60 *Kreuzer* = 240 *Denar*. Currency conversion until 1760 follows Pfister (2017). Augsburg joined the *Konventionstaler* regime in 1760 (Kruse, 1782, 60). The new rate is used from 1761. 1 *Kölner Mark* = 10 *Konventionstaler* (Gerhard, 2002, 213). 1 *Konventionstaler* = 144 *Kreuzer* (Kruse, 1782, 60). This yields an intrinsic value of 0.406 g Ag per *Denar*. This value is used until 1837.

From 1838 we asses the silver content of *Kreuzer* based on the values implied by the Munich currency treaty of 1837: 1 *Gulden* = 1/24.5 *Mark* of Cologne; the latter = 233.8555 g Ag (Rittmann, 1975, 535–7). 1 *Gulden* = 60 *Kreuzer*; 1 *Kreuzer* = 4 *Pfennig*.

For all prices from the ledger of the urban hospital (*Hospitalrechnungen*) we follow the volume conversion rules applied by Allen (2001), based on Elsas (1936, 153–4) (see also Pfister, 2017); see Verdenhalven (1993, 49) for a very similar volume conversion rate. This holds for barley and wheat until 1744, for oats and rye until 1799, and spelt until 1807 (Elsas, 1936, 361–9, 382–5).

We apply the conversion rate given by Fassl (1988, 104, note 26) to all prices of the urban grain market (*Schrannenpreise*), which are reported in the *Intelligenzblätter* (Fassl, 1988, 104, note 26; Elsas, 1936, 361–9, 382–5). This holds for barley, husk spelt, and wheat for the period 1745–1813, and for oats and rye 1800–13 (Elsas, 1936, 361–9, 382–5). In 1811 the official volume measures changed because of the integration of Augsburg into Bavaria; the grain market stuck to the old volume measures until 1813, however (Fassl, 1988, 104, note 26; cf. Verdenhalven, 1993, 49 for a very similar rate). To acknowledge these volume changes, we construct factors based on the ratio of volumes. Additionally, a different volume for oats applies from 1814 (Witthöft, 1993, 76).

Barley (1461–1820), oats (1459–1820), husk spelt (1455–1820)
Crop year prices for spelt in local units from Elsas (1936, 593–600); other crop year prices in g Ag per litre from Pfister (2017); cf. Allen (2001, 437). Original source is Elsas (1936). Prices of barley and oats are converted to calendar years according to the commodity specific relationships (Table S18).

Rye (1459–1855)
Until 1749 prices for crop years from Pfister (2017). Crop year prices 1457–1749 are converted to calendar years using the coefficients from the local time series relationship in table S16. From 1750 to 1850 we use calendar year prices in local currency per *Scheffel* based on quotes from the urban grain market (*Schrannenpreise*) from Fassl (1988, 421); 1851–55 from Seuffert (1857, 283).

Wheat (1670–1855)
Until 1814 crop year prices in metric units from Pfister (2017). Conversion to calendar year prices applies the commodity specific relationship (Table S18). Missing values are extrapolated by relating the weighted inter-annual growth rates of the husk spelt series and the last know observation for wheat. Growth rates (g) of crop year prices of wheat and husk spelt 1675–1820 are related as follows: $g_{\text{wheat}} = 0.9047^{***}\ g_{\text{husk spelt}}$, $R^2 = 0.83$, constant is zero. Extrapolation using husk spelt rests on the idea that spelt is a type of wheat and thus, both grains react in a similar way to weather, a major determinant of the inter-annual variation (correlation of both series in levels is $r = 0.96$; 1674–1820). From 1815–55 calendar year prices from Seuffert (1857, 282).

Berlin

Currency and volume conversion

Information on the number of *Silbergroschen* per *Taler* and on the latter's silver content back to 1623 is drawn from Statistisches Reichsamt (1935, 309–10, 314–5). We follow the calculation in Statistisches Reichsamt (1935, 309, 314–5) and apply the Prussian *Mark* of 1816 which is equivalent to *Mark* Cologne (233.85550 $\approx$ 233.856 g). This is because a third source lists *Mark* Cologne (233.856 g) as reference for coined *Reichstaler* in the early 18th century, i.e., 1 *Reichstaler* of *Graumann'scher Fuß* = 16.704 g Ag (Schrötter, 1903, 568; Schrötter, 1908, 85). We apply the new silver content of the *Taler* based on the new Prussian *Mark* = 500 g Ag from 1858 (and not from 1857). This is because the law is from May 1857 and the practical difference in silver contents small (Statistisches Reichsamt, 1935, 310); *Verordnung* issued in 1858 (Zich, 2009, 126).

Debasement of Prussian small coins 1808–21, the Prussian vellon inflation (*Preußische Scheidemünzinflation*), is acknowledged by using the Berlin specific adjustment factors from Statistisches Reichsamt (1935, 310) which are based on an *Agio*. The latter was due if payment was in *Scheidemünze*. Derivation of such factors in a related publication (Jacobs and Richter, 1935) for Königsberg was replicated and proved valid. Factors for Berlin in Jacobs and Richter are equal to those in Statistisches Reichsamt; for further details see Königsberg.

Scheffel are converted to litre following Witthöft (1993, 26), because Statistisches Reichsamt (1935, 314–5) does not provide evidence for frequent changes in volumes as indicated by Verdenhalven (1993, 49). No extra oats measure (available in Verdenhalven) used, because this is not indicated in the source.

Barley, oats, rye, wheat (1652–1871)

Calendar year and crop year prices in *Silbergroschen* per *Scheffel* are from Statistisches Reichsamt (1935, 317–8). Data for rye and wheat 1624–51 are not included because the annual averages are based on five to six observations at best, and several values rest on only one observation (Statistisches Reichsamt, 1935, 308, 321). Without checking the original source it is not possible to distinguish between averages based on five observations and annual values resting on only one observation.

Braunschweig

Currency and volume conversion

Conversion of *Gulden* and from 1661 *Reichstaler* based on Gerhard and Kaufhold (1990,

28–9, 413–4). The first information on silver contents is for 1534. *Scheffel* are converted into litre following Gerhard and Kaufhold (1990, 398).

Barley, oats, rye, wheat (1572–1850)
Calendar year averages of the sales of the collegiate church *St. Blasius* until 1744, then retail market prices (Gerhard and Kaufhold, 1990, 29, note 2 and 8). Data for 1513–71 omitted, because of unknown silver contents. The same applies to some scattered *Martini* prices 1330–1512.

Celle

Currency and volume conversion
Following Gerhard and Kaufhold (1990, 414) we applied the conversion rates for Hanover (see below). Additionally, we assumed that the first known silver content for *Denar* (*Pfennig*, respectively) for 1740 can be applied to the period 1727–39.

Barley, rye, wheat (1727–1871), oats (1727–1846)
Prices 1727–1846 in *Mariengroschen* and *Denar* per *Hannoversche Himten* from Gerhard and Kaufhold (1990, 31–3, 92–4, 148–50, 207–9); 1818 change in currency to *Gute Groschen* and *Pfennig*, 1854 to *Neue Groschen*, respectively. Prices from 1847–71 are extrapolated from *Martini* prices (except oats). We applied the commodity specific parameters from the panel data regressions (eq. S4), because overlapping data are available for only 1835–1846.

Cologne

Currency and volume conversion
Prices from 1532–1796 in *Albus (Rechengeld)* per *Malter*. Currency converted as in Pfister (2017). The relevant series of *Rechenalbus Mittelkurse* from Metz (1990, 366—95) runs until 1790. We assume that the ratio of 1790 can be applied until 1796. *Malter* are converted according to Verdenhalven (1993, 34) (cf. Schimmelfennig, 1820, 63; Ebeling and Irsigler, 1976, XI, note 3 give a different value). Data from Kopsidis (1994) and Prussian sources converted as for Berlin.

Barley, rye, wheat (1532–1871)
Data for 1532–1786 consist of arithmetic means of monthly data from Ebeling and Irsigler (1976, 536–663), accessed through GESIS Köln (2005b). For some years monthly observations are missing, and the mean consists of less than 12 prices. Because of too little observations (less than six months) or a weak coverage of the calendar year we omit the following

years: 1531, 1541, 1787–1791. 1787–1819 calendar year prices in *Reichstaler* per *Scheffel* from Kopsidis (1994) (cf. GESIS Köln, 2015). 1820–59 from Geheimes Staatsarchiv Berlin (b). Prices are given in *Gute Groschen/Silbergroschen* (from 1822). 1824 and 1860–4 from Königliches Statistisches Bureau (1867); 1865–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72).

Martini prices are calculated as November-December averages from the monthly data for 1531–1796 (purpose: comparison to calendar year prices).

Rahlf (1996) provides a rye calendar year price series for the years 1531–1797 based on Ebeling and Irsigler (1976) (accessed through GESIS Köln, 2005a). We were able to reproduce both the calendar year series from Rahlf and the crop year series by Ebeling and Irsigler (1976) (except minor differences) using monthly data.

We omitted a mixed rye/wheat series 1443–1530 from Irsigler (1975, 519–21) *Kölnische Mark (Rechengeld)* per *Malter*, accessed through GESIS Köln (2005b). For 1445–1530 data refer to the mean price of rye *and* wheat in *october* and november. Thus, too many assumptions would be necessary for extrapolating calendar year prices and attempts of extrapolation.

Oats (1532–1871)
Cf. description of the rye series. However, for oats we omitted data in 1541, 1678–9, 1683, 1787–91.

Dresden

Currency and volume conversion
Currency conversion as for Leipzig. Volumes are converted following Witthöft (1993, 141).

Barley, oats (1602–1782), rye, wheat (1602–1869)
Calendar year prices 1602–1782 in *Taler Kurant* per *Scheffel* from Kraus (1808). Calendar year average prices 1783–1869 were calculated based on monthly data from Uebele et al. (2013); original source is a weekly newspaper (Uebele et al., 2013, 3). Data for 1825–34 are not available.

Emden

Currency and volume conversion
Conversion of currencies and volumes follows Gerhard and Kaufhold (1990, 402, 414–5). We assumed that the conversion rates for *Gemeiner Taler* and *Stüber* for the year 1788 can be

applied to the period 1780–87, and those for *Reichstaler* and *Gute Groschen* for the year 1820 to 1796–1819.

Barley, oats, wheat (1780–1850)
Prices 1780–1813 in *Gemeinen Talern* and *Stüber* per *Emder Last* from Gerhard and Kaufhold (1990, 40–2, 100–2, 215–7); 1814–37 in *Dalern*, and 1838–50 in *Reichstalern* and *Guten Groschen*. We omit data 1746–79 due to unclear currency conversion. The price for the year 1838 is the mean of both values given in different currencies by Gerhard and Kaufhold (1990). Calendar year prices for the period 1807–13 are extrapolated from *Martini* prices (in *Reichstalern* and *Guten Groschen*) according to the results in table S6.

Further omitted price series include a series for *Sommergerste* (spring barley) for the same period. These prices are not taken into account, because *Sommergerste* should be considered as a different commodity (Gerhard and Kaufhold, 1990, 43–4). This also holds for *Brau-Hafer* (oats, brewing quality) (1772–1850), which is likely to be of particularly good quality (Gerhard and Kaufhold, 1990, 102–3). There are price series for *Ostsee-Weizen* (wheat, imported from the Baltic Sea) (1748–1832) and *Ostsee-Roggen* (rye, imported from the Baltic Sea)(1748–1842) which are both shorter and do not offer opportunities for extrapolation. Furthermore, there are more *Martini* prices for rye, barley, and oats in Oberschelp (1986, 86–97) for which, however, the currency conversion remains unclear to us, because they are listed in the currency that usually applies to Hanover.

Rye (1780–1850)
Cf. barley, oats, wheat. Prices are from Gerhard and Kaufhold (1990, 156–8). Additionally, prices for 1780–83 are extrapolated from the series *getrockneter Roggen* (dried rye) by adjusting these prices for the mean level difference for the years 1775 and 1784 (-16.68%).

Frankfurt (am Main)

Currency and volume conversion
Conversion follows Pfister (2017).

Barley (1604–1797), oats (1372–1799), rye (1352–1799), wheat (1421–1799)
Crop year prices in *Pfennig* per *Achtel* from Pfister (2017); original data from Elsas (1940). Calendar year prices were obtained by extrapolation following eq. S8 using the commodity-specific parameters in Table S18.

Gdansk

Currency and volume conversion

Currency and volume conversion follows Pfister (2017) with two exceptions. (1) In addition, we account for the Prussian vellon inflation. (2) Like Königsberg, data from Prussian sources in 1814–1871 are converted as for Berlin. To account for the vellon inflation in 1808–22, we apply the specific adjustment factors for Königsberg from Jacobs and Richter also to Gdansk (reason: geographic proximity; see Königsberg for details on adjustment factors).

Barley, rye (1501–1871), oats (1501-1700; 1814–71), wheat (1814–71)

Calendar year prices for barley and rye until 1815 (oats until 1700) are in *zloty* per *last* from Pfister (2017). Original sources are Furtak (1935) and Pelc (1937). The observations in 1814/15 are averages of data from Furtac and Prussian sources starting in 1814. The relative movement for the two overlapping years (1814/15) for rye and wheat is similar but large relative mean differences for the latter exist (e.g., mean of Furtak/Pelc rye series is ca. 36% *smaller*). A systematic shift of either series is unlikely because the 10-year-mean around the overlapping years show that Furtak/Pelc data (1806–1815) are ca. 17% *higher* than Prussian data (1814–1823). Thus, we simply average the observations of both sources for 1814/15 and use Prussian data from 1816.

Barley, oats, rye and wheat for 1814–65 from Königliches Statistisches Bureau (1867, 125–6, 129–30, 120–1, 116–7) and for 1866–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72); all data in *Silbergroschen* per *Scheffel*.

Goch

Currency and volume conversion

See Xanten.

Barley, wheat (1800–1860), oats, rye, wheat (1800-1882)

From 1800 prices local/Prussian currency per *Berliner Malter* from Beissel (1889) accessed through Jacks' Database (1803–17 checked with Beissel; observations for 1805 and 1807 refer to Xanten and are omitted; Beissel, 1889, 116).

Göttingen

Currency and volume conversion

Cf. Hanover for conversion of currencies and volumes Gerhard and Kaufhold (1990, 403, 405, 415). For the period 1736–1857 this is identical to Göttingen in Pfister (2017).

Barley/oats (1632/4–1850), rye, wheat (1632–1867)

Prices from Gerhard and Kaufhold (1990, 45–9, 104–7, 161–5, 219–22) in *Reichstalern* and *Mariengroschen* per *Hannoversche Malter*, since 1832 (rye, wheat 1834) in *Reichstalern* and *Guten Groschen*. Rye and wheat since 1859 in *Reichstalern* and *Neuen Groschen*. Calendar year price for barley, oats, rye, and wheat 1632–1715 are extrapolated from *Martini* prices according to the regression results in Tables S6 and S7. Regressions are based on *Martini* prices for 1764–1863 in *Groschen* and *Denar* per *Hannoversche Himten* from Oberschelp (1986, 82–97). For 1716–66 there are two prices for each year which cover spring (*Frühjahrspreis*) and autumn (*Herbstpreis*). For this period we chose the mean of both prices as calendar year price. Since 1767 calendar year prices as such are provided (exception: barley, oats 1812–3; for these years we calculated the calendar year price as for the period 1716–66).

Halle

Currency and volume conversion

Intrinsic value of *Taler* follows Berlin (including currency debasement 1808–21). Change from *Gute Groschen* to *Silbergroschen* is applied in 1822/3 (in accordance with source of prices). 1 *Hallesche Scheffel* = 1.4 *Berliner Scheffel*; 1 *Wispel* = 24 *Berliner Scheffel* (Naudé and Schmoller, 1901, 516, 541); the latter converted as for Berlin.

Barley, oats, rye, wheat (1601–1871)

Calendar year prices 1601–1691 and 1757–1815, 1818, 1824 (1824 not for rye, see below) converted from *Martini* prices applying the local time series relationship (Table S7). Prices 1692–1739 are averages of annual minimum and maximum prices from Naudé and Schmoller (1901, 541—3) and Naudé et al. (1910, 623); original list by Löwe published in 1789 (Naudé and Schmoller, 1901, 514). Units are *Reichstaler* and *Groschen* per *Wispel*. Prices 1740–56 in *Reichstaler* and *Groschen* per *Berliner Scheffel* from Naudé et al. (1910, 615–622). Prices 1816–59 in *Gute Groschen/Silbergroschen* per *Berliner Scheffel* from Geheimes Staatsarchiv Berlin (b). Rye price (annual average of monthly values) in 1824 and 1860–4 from Geheimes Staatsarchiv Berlin (a). Prices 1865–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72).

Martini prices 1600–1749 in *Guten Groschen* per *Hallesche Scheffel* (until 1713) or *Berliner Scheffel* (from 1714). *Martini* prices 1600–1749 from Jacks' Database (2016), who took them from the Beveridge papers. Original list was published in 1750 by Dreyhaupt (see discussion in Naudé and Schmoller, 1901, 512–6), an accessible later publication is

Königliches Statistisches Bureau (1867, 108–9). *Martini* prices 1749–1834 in *Groschen* (until 1822) and in *Silbergroschen* (from 1823) per *Berliner Scheffel* kindly provided by Katrin Moeller from Historisches Datenarchiv Sachsen-Anhalt; originally from Runde (1933).

Hamburg

Currency and volume conversion

Currency conversion of data from Gerhard and Engel (2006) and Gerhard and Kaufhold (1990) as in Pfister (2017). Volumes are converted according to Gerhard and Kaufhold (1990, 404) but (as Pfister, 2017) we use the old *Fass* rate valid until 1829 from Gerhard and Engel (2006, 317).

Currencies and volumes of prices from Statistisches Reichsamt (1935) (identical to series in Jacobs and Richter, 1935) for the period 1792–1850 were first reconverted from *Mark* (= *Goldmark* of 1873) per 1000 kg to original local currency and units. Second, local currencies and units were converted to g Ag per litre. Thus, we apply 1 *Mark banco* of Hamburg = 1.517 *Mark* (Jacobs and Richter, 1935, 17) to prices from Statistisches Reichsamt to obtain prices in *Mark banco*. We then apply the exchange rate of *Mark courant* per *Mark banco* used by Jacobs and Richter (1935, 18). This yields prices in *Mark courant* which were converted to g Ag using the g Ag per *Schilling* series (1 *Mark courant* = 16 *Schilling*) from Pfister (2017).

Volumes are converted by deriving the kg per litre ratio from the kg per *Last* values (different for each grain type) from Jacobs and Richter (1935, 16) and the litre content of *Last* from Gerhard and Kaufhold (1990, 404) taking into account changes in litre volumes of *Last*.

Barley (1736–1850), oats (1736–1850), wheat (1736–1850)

Until 1850 (barley: 1849) wholesale bid prices in *Reichstaler* and *Schilling* per *Hamburger Last* from Gerhard and Kaufhold (1990, 50–2, 108–9, 223–4); 1823–42 (wheat 1822–42) in *Mark* and *Schilling* (price 1). We reduced these prices by 4.5% in order to approximate actual market prices (Gerhard and Kaufhold, 1990, 50, note 1, 395). The barley and oats series are extrapolated with stock market prices in *Mark* per 1000 kg from GESIS Köln (2008) (original: Statistisches Reichsamt, 1935, 300–3, 304–7) (price 2).

Barley: Value for the year 1850 extrapolated from price 2. The estimated relationship 1799–1849 is: $\alpha = 0.0789^{**}$, $\beta_1 = 0.6891^{***}$, $R^2 = 0.82$, dependent variable: price 1; explanatory variable: price 2.

Oats: Values for the years 1801, 1804, 1846–9 are extrapolated from price 2. The estimated relationship 1792–1850 is: $\alpha = -0.0079$, $\beta_1 = 0.8086^{***}$, $R^2 = 0.95$, dependent variable: price 1; explanatory variable: price 2.

Rye (1540–1850)

The core of the series are prices of the *St. Hiob* hospital (1540–1821) in *Mark* and *Schilling* per *Wispel* from Gerhard and Engel (2006, 108–13) (price 1). Missing values and data after 1821 are extrapolated as follows.

(1) Values for the years 1546, 1551, 1557, 1580, 1595, 1600, 1604 are extrapolated from prices of *St. Georg* hospital (price 2) (Gerhard and Engel, 2006, 108–9; volumes refer originally to *Scheffel*). The estimated relationship 1540–1612 is: $\alpha = 0.0159$, $\beta_1 = 0.8484^{***}$, $R^2 = 0.68$, dependent variable: price 1; explanatory variable: price 2. We avoided backwards extrapolation of the values for rye 1443–75, 1500, 1510–39 with prices from *St. Georg* hospital due to the relatively weak model fit. (2) Values for the years 1795, 1798–1800, 1822–50 are extrapolated from stock market prices (price 3) (GESIS Köln, 2008, original Statistisches Reichsamt, 1935, 292–4; in *Mark* per 1000 kg). The relationship 1792–1821 is: $\alpha = 0.0505$, $\beta_1 = 0.9768^{***}$, $R^2 = 0.87$, dependent variable: price 1; explanatory variable: price 3. (3) Value for 1656 is extrapolated from an unspecified rye series from Gerhard and Kaufhold (1990, 166–9) (price 4). The relationship 1638–1790 is: $\alpha = 0.0606^{***}$, $\beta_1 = 0.9043^{***}$, $R^2 = 0.74$, dependent variable: price 1; explanatory variable: price 4. (4) Value for 1806 is extrapolated from the rye (*Mecklenburg*) series from Gerhard and Kaufhold (1990, 170–1) (price 5). The relationship 1736–1821 is: $\alpha = 0.0024$, $\beta_1 = 1.0688^{***}$, $R^2 = 0.88$, dependent variable: price 1; explanatory variable: price 4.

Hanover

Currency and volume conversion

Conversion to g silver follows Gerhard and Kaufhold (1990, 415–6) and Oberschelp (1986, 101). Modifications 1788 for *Reichstaler* and *Gute Groschen* 1790–1833 as in (Pfister, 2017, Online appendix S1, 5; refers to Göttingen). We decided to apply the rate for the volumes prior 1714 according to Verdenhalven (1993, 23, 49, 72), because the value from Gerhard and Kaufhold (1990, 405) is not plausible and because the date of introduction of the new volume is inconsistent with other literature (cf. Oberschelp, 1986, 27, 47). After 1714 we follow Gerhard and Kaufhold (1990, 405); these rates are very similar to those given by Verdenhalven.

Barley, oats, rye, wheat (1590–1863)

Prices 1590–1691 in *Mariengroschen* per *Scheffel* from Oberschelp (1986, 13–5) (for oats we calculated the mean of the two qualities *rauh* and *weiß*); 1700–49 calendar year averages based on monthly data in *Thaler* and *Mariengroschen* per *Himten* from Oberschelp (1986, 20–45) (original Unger, 1752, 237–62); 1750–1817 in *Mariengroschen* and *Denar* per *Himten*, since 1818 in *Guten Groschen* and *Pfennig* from Gerhard and Kaufhold (1990, 53–5, 111–2, 172–3, 225–6). The observation in 1847 is calendar year average from a different primary source. In 1817 and 1834 there are two calendar year averages provided for each year; hence, we used the mean.

Martini prices are calculated as November-December averages from the monthly data for 1700–49; for 1764–1863 they are taken from Oberschelp (1986, 82–97).

Missing values in the calendar price series for the years 1796–1808 (oats, rye until 1810) and 1851–63 are extrapolated from *Martini* prices according to the results from table S8.

Königsberg

Currency and volume conversion

1688–1756: Prussian *Gulden* (*Groschen*) converted to *Reichstaler* applying the relationships: 1 *Reichstaler* = 3 Prussian *Gulden*; 1 *Reichstaler* = 90 Prussian *Groschen* (Naudé et al., 1910, 662; Schrötter, 1903, 247, footnote 1). Silver content of *Reichstaler* derived as follows.

Before 1740 currency conversion follows Berlin but applies adjustment factors to reflect the slightly higher silver contents of minted coins in Eastern Prussia compared to Berlin (which holds for many but not all coins, Schrötter, 1903, 568). E.g., 1724 the mint master of Königsberg had to apply a silver content of 18.9 g per *Taler* while minting coins of a nominal value of 2 *Groschen* (Schrötter, 1903, 400). The Berlin series gives a value of 17.9 g silver per *Taler* for 1724. To reflect the higher silver content of minted coins in Königsberg, we base the adjustment factor (1.056) on the ratio of the values from Königsberg and Berlin. For 1700 a similar factor (1.062) based on a coin with a nominal value of 18 *Groschen* obtains (Schrötter, 1903, 568). We use the average of both factors (1.059) to scale the silver content for Königsberg upward (factor used for period 1688–1724). To reflect the ongoing export of coins with high silver content to Poland reported by Schrötter, the adjustment factor decreases exponentially from 1725 until it reaches 1 in 1740.

From 1740 the silver content is based on the unadjusted Berlin series. This is justified by the scarcity of money since 1743 (Schrötter, 1908, 249) and the King's order to apply the *Graumann'schen Fuß*, i.e. 14 *Taler* per *Mark* Cologne or 16.704 g Ag per *Taler*, in

1754 (Schrötter, 1908, 252–4; Berlin series: *Graumann Taler* 1750–1857 taking account of debasement in 1759–64).

1797–1810: Reconversion of data from Jacobs and Richter (1935, 18–21) to original currency using the factor 1 *Reichstaler (courant)* = 3 *Mark* of 1873. The silver content of the *Reichstaler* follows Berlin. Payment was likely done in smaller coins such as *Gute Groschen* (*Scheidemünze*), which were subject to debasement in the period 1808–1823 according to Jacobs and Richter. In their conversion, these authors already accounted for the *Aufgeld* (*Agio*; an additional payment) which was demanded if payment was not in *Reichstaler*.

Jacobs and Richter also discuss whether payment was actually done in *courant* or *Scheidemünze*; reasons for payment in *courant*: (i) source contains notes stating that payment during bad years must be in *courant*, (ii) payment to 'foreign' trade partners who would accept only *courant* (*Preis-Kurant* is the source). We follow Jacobs and Richter and assume that the *Aufgeld* was necessary. Hence, we do *not* reverse this part of their calculation.

Our replication of their data treatment shows only minor differences for 1808–12 and thus, we keep their calculations. We apply these factors also to data from Prussian sources in 1811–22.

1811–1871: Conversion of currency as for Berlin, but applies Königsberg specific adjustment factors from Jacobs and Richter for 1811–22.

Volume conversion applies the following relationships to *Scheffel* of Berlin: 1 *Last* (of Königsberg) = 56.5 *Berliner Scheffel*; 1 *Last* of oats = 51 *Berliner Scheffel* (Naudé et al., 1910, 620).

Conversion of prices per kg from Statistisches Reichsamt (1935) back to volumes as follows: kg per *Scheffel* of Königsberg and kg per *Preußische Scheffel* from Jacobs and Richter (1935, 16) (cf. Statistisches Reichsamt, 1935, 281); litre per *Scheffel* of Königsberg from Witthöft (1993, 267); litre per *Preußische Scheffel* (as for Berlin) from Witthöft (1993, 26).

Barley, oats, rye, wheat (1688–1871)

1688–1730: Average of minimum and maximum prices in Prussian *Gulden* per *Last* from Naudé and Schmoller (1901, 618–9). The type of prices is unspecified. They might be wholesale prices, because the volume is in *Last*), however, this is not certain (hence, no adjustment of price levels to retail prices is made). 1731–39: gap.

1740–1756: Annual averages of weekly prices in Prussian *Groschen* per *Berliner Scheffel* calculated by Naudé et al. (1910, 664—71) based on the *Wöchentlichen Königsbergischen Frag- und Anzeigungsnachrichten*. To fill gaps in 1748, 1749 and 1751 they used information from the *Preiskurant* compiled by the merchant guild of Königsberg. These are wholesale prices. We did not adjust these prices to approximate retail prices due to a lack of support for

the choice of the adjustment factor. Assuming a mark-up without empirical backup would be arbitrary. Keeping wholesale prices works against our hypothesis because it increases the spread with other markets and signals disintegrating markets.

1797–1810: prices in *Mark* per 1000 kg from Jacobs and Richter (1935, 52–3) (accessed through GESIS Köln, 2008). Prices 1792–6 are not used because Jacobs and Richter extrapolated these data from prices in Berlin (see their note 7).

1811–1871: Data in *Gute Groschen* (until 1821)/*Silbergroschen* (from 1822) per *Berliner Scheffel* are from the following sources: 1816–59 from Geheimes Staatsarchiv Berlin (b); 1818 (in *Silbergroschen*), 1824 and 1860–4 from Königliches Statistisches Bureau (1867); 1865–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72). Visual inspection of overlapping data for 1811–26 from Jacobs and Richter and Prussian sources shows that the observations for 1819 are not well in line but the correlation of series from both sources is still satisfactory (ranges from 0.86 for wheat to 0.94 for rye). Removing 1819 increases these correlations for all cereals (e.g., for wheat to ca. 0.91).

Leipzig

Currency and volume conversion

Follows Pfister (2017). The *Zollpfund* (500 g silver) is applied from 1858 (see Berlin). 30 Taler were minted from the *Zollpfund*, which implies a slightly reduced intrinsic value for *Taler* and *Pfennig*.

Barley/oats (1574/87–1820), rye, wheat (1577/74–1860)

Crop year prices for barley, oats, rye, wheat (1564–1820) in *Denar* per *Scheffel* from Allen's Database (2001), based on Elsas (1940). We applied the conversion rule developed above to extrapolate calendar year prices (Table S18; rye, wheat: until 1818). Rye and wheat: 1819–1860 calendar year prices in *Groschen* per *Scheffel* from Pfister (2017); original Koehler (1967, 366–78).

Lüneburg

Currency and volume conversion

Cf. Hanover for conversion of currencies. We assume that the silver content for *Guter Groschen* (1817) is valid for the period 1790–1817 (Cf. Gerhard and Kaufhold, 1990, 62, note 14). Volumes are converted according to Gerhard and Kaufhold (1990, 407).

Barley, oats, rye, wheat (1764–1863)

Prices in *Mariengroschen* and *Pfennig* per *Lüneburger Himten* from Gerhard and Kaufhold (1990, 58–63, 114–7, 176–8, 229–31); from 1790 in *Guten Groschen*. *Martini* prices in *Groschen* and *Denar* from Oberschelp (1986, 90–3, 94–7, 86–9, 82–5). Missing calendar year prices for barley, oats, rye, wheat for the years 1765, 1820–31 (rye 1816–31) and 1851–63 are extrapolated from *Martini* prices according to the results in Tables S8 and S9.

Prices 1550–1763 are omitted due to missing information on the conversion to grams silver. We omitted implausibly low calendar year prices for rye 1816–9 given by Gerhard and Kaufhold. According to these values, rye would be cheaper than barley which is usually not the case. Thus, we replaced these prices with calendar year values estimated from *Martini* prices, which did not show this anomaly.

Minden

Currency and volume conversion

Currency conversion of data from Naudé and Schmoller (1901) and Naudé et al. (1910) for 1638–1747 applies the intrinsic value of the *Reichstaler* developed for Westphalia (see Currency Westphalia). Currency conversion for data from 1775 from Gerhard and Kaufhold applies intrinsic values for Herford (Gerhard and Kaufhold, 1990, 416–7). From 1834 these silver contents are equivalent to those in Berlin (Prussian currency).

Volume conversion for data 1638–47 applies the same rate as for Berlin. Instead of the rates provided for volume conversion in Gerhard and Kaufhold (1990, 408), we also applied the same rates as for Berlin to data from Gerhard and Kaufhold, because all prices are per *Berliner Scheffel*.

Barley (1651–1850), oats (1640–1850), rye (1641–1850), wheat (1724–1850)

All November/December prices (treated as *Martini* prices) 1638–1747 from Naudé and Schmoller (1901, 534–8) and Naudé et al. (1910, 602) in *Taler*, *Mariengroschen* and *Pfennig* per *Berliner Scheffel*. All prices from 1775 in *Reichstalern* and *Guten Groschen* per *Berliner Scheffel* from Gerhard and Kaufhold (1990, 64–7, 118–20, 180–2, 232–4); since 1822 (*Martini* prices since 1817) in *Reichstalern* and *Silbergroschen*. Calendar year prices 1640–1747, 1805–7 and 1811–4 are extrapolated from *Martini* prices according to the results in table S9.

Munich

Currency and volume conversion

The currency system is 1 *Gulden* = 60 *Kreuzer* = 240 *Denar* (Elsas, 1936, 116). Currency conversion until 1740 and 1800–1820 as in Pfister (2017). 1741–1800: The *Konventionstaler* regime that was agreed with Austria in 1754 was revoked in 1755. In 1759 the Konventionstaler was set to 2.5 *Gulden* = 150 *Kreuzer* (Kruse, 1771, 270). 1 *Kölner Mark* = 10 *Konventionstaler* (Gerhard, 2002, 213). This implies an intrinsic value of the *Denar* of 0.390 g Ag. For the period 1741–1758, the values are determined by exponential interpolation. This reflects the failure of monetary cooperation at the Imperial level in 1740 (Gerhard, 2002, 262–5) and the beginning of devaluation around that time in Cologne and Hamburg. In 1766, the *Kreuzer* was stabilized at 144 *Kreuzer* per *Konventionstaler* (Kruse, 1771, 270). From 1800 the rates for g Ag per *Denar* are from Gerhard (1984, 623–4).

Volume conversion until 1820 as in Pfister (2017); a separate measure applies for oats (Witthöft, 1993, 329).

For conversion of currencies and volumes of prices in 1811–2, 1819, and the period 1821–1885 from Statistisches Reichsamt (1935) (identical to series in Jacobs and Richter, 1935), the following procedure applies (cf. Hamburg). First prices in *Mark* (= *Goldmark* of 1873) per 1000 kg are reconverted to original local currency and units. Second, local currencies and units are converted to g Ag per litre. Thus, we apply 1 *Gulden* = 1.75 *Mark* (1792–1810) and 1 *Gulden* = 1.7143 *Mark* (1811–1875) (Jacobs and Richter, 1935, 21) to prices from Statistisches Reichsamt to obtain prices in *Gulden*. We assume that prices are in *Rechnungsgulden* (like prices from Elsas, 1936 prices in Statistisches Reichsamt are from the urban grain market *Schranne*). We apply the relationship between *Gulden*, *Kreuzer* and *Denar* as given by Elsas (1936, 116) and multiply with the series of g Ag per *Denar* by Pfister (2017).

Volumes are converted by deriving kg per litre ratio from the kg per *Scheffel* values (different for each grain type) from Jacobs and Richter (1935, 16) and the litre content of *Scheffel* from Witthöft (1993, 329).

Barley (1514–1874), oats (1452–1874), rye (1452–1874), wheat (1512–1874)

Crop year prices for oats (1400–1690) and rye (1404–1690) in *Denar* per *Scheffel* from Allen's Database (2001, Munich file; checked with original source) (Allen, 2001; original in Elsas, 1936, 539–45: *Kammerrechnungen*). Calendar year prices are obtained following the results for oats and rye from *Kammerrechnungen* in Table S16. Prices for barley (1514–1690), wheat (1512–1690), and if possible missing values in the series for oats and rye are extrapolated using the calendar year conversion of the crop year series referring to the *Heilig-Geist* hospital

(Allen's Database, 2001, Munich file; checked with original source; original Elsas (1936, 560–5). Cf. results for barley, oats, rye and wheat from *Heilig-Geist* in Tables S16 and S17. All estimations for the conversion to calendar year prices rely on the overlap with the following series.

1691–1820: Calendar year prices calculated as arithmetic averages of monthly data in *Denar* per *Scheffel* from Elsas (1936, 677–9, 680–2, 674–7, 671–4). We obtained prices for 1780–7, which are not provided by Elsas, in *Gulden* and *Kreuzer* per *Scheffel* for each month from the *Intelligenzblätter* (Churfürstlich Pfalzbaierrisches Intelligenz- und Addreß-Comtoir, 1780–3; 1784–7). We applied Elsas' method by using the price for average quality (*Vom Mittern*) of the first *Schranne* per month (Elsas, 1936, 350–1). A comparison with the mean of all monthly observations from the *Intelligenzblätter* for the same period shows that Elsas' method is valid: the correlation ranges between 0.9135 (oats) and 0.9846 (rye). The relatively poor fit for oats is due to an outlier for the observation November 1787 which may originate from a printing error.

Three missing observations for 1811–2, 1819, and prices for the period 1821–55 in *Mark* per 1000 kg are from GESIS Köln (2008), original in Statistisches Reichsamt (1935, 300–3, 304–7, 293–5, 296–9). Prices from Elsas for 1691–1820, and from Statistisches Reichsamt for 1811–2, 1819, and 1821–1855 refer to the urban grain market (*Schranne*). The calendar year prices from Statistisches Reichsamt are higher than those based on monthly data from Elsas. (A possible reason could be that the kg per *Scheffel* values Jacobs and Richter applied are different from those used to recalculate per *Scheffel* prices prior to conversion to litres.) Both series correlate almost perfectly for the overlapping period 1791–1820. To avoid the shift in the mean price level, we adjusted each series from Statistisches Reichsamt with a commoditiy specific factor downwards when splicing it with data from Elsas (downwards adjusment ranges from 10 to 13%). Calendar year prices Prices 1856–85 omitted (focus of analysis until 1850).

Crop year prices for both periods July-June and August-July 1690–1817 and *Martini* prices (as November-December averages) 1690–1818 for barley, oats, rye and wheat are calculated on the basis of monthly data if possible.

Münster

Currency and volume conversion

Currency conversion of *Martini* prices from Gerhard and Kaufhold (1990) until 1763 follows detailed description in *Currency Westphalia* (see below). From 1763–1826 we apply the *Konventionstaler*; from then on as for Berlin: 1827–57 Graumannscher Fuß (16.704 g); from

1858 we follow *Wiener Münzvertrag* (16.667 g). *Malter* converted following Gerhard and Kaufhold (1990, 409); *Scheffel* as for Berlin.

Currency Westphalia

Traditional currency systems: bishopric of Münster (Gerhard and Kaufhold, 1990, 71; Schwede, 2004, 436) 1 *Reichstaler* = 28 *Schillinge*; 1 *Schilling* = 12 *Pfennige*.

Traditional currency systems: bishoprics of Osnabrück and Paderborn (Kennepohl, 1938, 25–6; Schwede, 2004, 43) 1 *Reichstaler* = 21 *Schillinge*; 1 *Schilling* = 12 *Pfennige*.

Widely used divisions of the Reichstaler in North-Western Germany from 1623: 1 *Reichstaler* = 36 *Mariengroschen*; 1 *Reichstaler* = 24 *Gute Groschen*.

Prussian currency reform of 1822/3: 1 *Taler* = 30 *Silbergroschen*; 1 *Silbergroschen* = 12 *Pfennige*.

Silver content of Reichstaler: explanations. The Imperial mint ordinance of 1566 was implemented in Westphalia in 1569 (Kennepohl, 1938, 167); before this year, no certain information for the silver content of the Taler exists. Until 1655 we assume silver content according to the Imperial mint ordinance (25.984 g).

From 1691 to 1740 silver content is set according to the Leipzig convention of 1690 (19.488 g). Exponential interpolation is used to calculate values in 1655–90. The starting point of debasement is set to 1656 because contemporaries deplored a flooding of Westphalia with debased *Mariengroschen* following the war between Brandenburg and Poland in 1656/7. Extrapolated values are consistent with the results of examinations of Westphalian coin by the regional mint authorities in 1675 and 1680 (Kennepohl, 1938, 202—4; Schwede, 2004, 177—86, 258—60, 267).

In 1763 most Westphalian territories accepted the *Konventionstaler* regime (17.539 g; Schwede, 2004, 28). We let debasement start in 1741 and again define values in 1741—57 by the exponential trend between the silver content of 1690—1740 and from 1763 onwards. 1741 conforms to the initial phase of the Austrian War of Succession (1740—8) and the beginning of a longer phase of debasement in Cologne (silver content of *Albus* according to Metz, 1990, 366–95).

Massive debasement during the Seven Years' War (1756–63) is captured by a short series of exchange rates of local currency against the *Louis d'or* in Paderborn (Schwede, 2004, 442; value in June for 1757, values for January for subsequent years until 1763; we compute pairwise averages of the latter figures to obtain mid-year values). We use changes of the exchange

rate against the value in June 1757 and our interpolated value for 1757 to extrapolate the intrinsic content of the *Reichstaler* in 1758–62.

Barley, oats, rye, wheat (1570–1868)
Calendar year prices 1570–1815 extrapolated from *Martini* prices applying the local times series relationship (Table S10). 1816–1871 based on Prussian data in *Gute Groschen* (since 1822 in *Silbergroschen*) per *Scheffel*: 1816–59 calendar year prices from Geheimes Staatsarchiv Berlin (b); 1818, 1824, 1860–64 from Königliches Statistisches Bureau (1867); 1865–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72).

Martini prices 1569–1863 (*Kappensaat*; see notes by Gerhard and Kaufhold, 1990, 71—2) in *Reichstaler* and *Schilling* (since 1827 in *Silbergroschen*) per *Malter* of Münster from Gerhard and Kaufhold (1990, 68—72, 121—4, 183—6, 235—8).

Nuremberg

Currency and volume conversion
Original prices until 1671 are in a system of *Rechengeld* (money of account). The silver content was assessed in three steps. We replicate Bauernfeind's conversion to gold and then convert to silver via a gold silver ratio. Details: First, we converted to *Rheinischer Gulden* (a gold currency) using the rates of *Denar* or *Kreuzer* per *Rheinischer Goldgulden* according to Bauernfeind (1993, 390–9). Missing observations in these rates are interpolated with the last known value. For the following years (sub-periods), no value is given at all: 1430, 1446–7, 1449, 1453, 1461, 1474–1515, 1517–8, 1520, 1526, 1529, 1531, 1534.

Second, we converted *Rheinische Gulden* to g gold (Au) according to Bauernfeind (1993, 60; based on Metz, 1990, 345–63).

Third, we converted gold prices to g Ag with the help of the Ag / Au ratio for Cologne from Metz (1990) as in Pfister (2017).

Currency conversion of prices 1761–1811 from Bauernfeind et al. (2001) and prices from Seuffert (1857) 1812–55 following Pfister (2017).

Volume conversion for prices until 1671 and for the period 1761–1811 applies the ratio given by Bauernfeind (1993, 72, 511); litre rate for *Scheffel* for data in 1812–55 from Witthöft (1993, 76).

Rye, wheat (1490/1498–1671; 1761–1855)
1490–1671: Calendar year prices are calculated as arithmetic mean of monthly prices. Raw data are kindly provided by Walter Bauernfeind and are from Bauernfeind et al. (2001, 286–7) (personal communication with Walter Bauernfeind). Monthly grain prices are derived

from official grain price estimates from the retail market, which were used to fix the bread weight (a process called *Raitung*) and therefore represent retail prices (Bauernfeind et al., 2001, 285–6). Until 1671 these data rest on Bauernfeind (1993). Additionally, the data are corrected for a change from the Julian to the Gregorian calendar which appeared in 1700 (Bauernfeind et al., 2001, 287, note 15). Raw data are in *Denar* per *Nürnberger Sümmer*. As in Bauernfeind (1993), prices from September 1504 until June 1514 are reduced by the tax of 32 *Denar* per *Nürnberger Sümmer*; the same holds for February 1576 until December 1579 with 126 *Denar* (Bauernfeind, 1993, 78, 220 note 376, 241).

1672–1760: Gap; raw grain price data would be available from Bauernfeind et al. (2001). However, there is no reliable currency information to obtain silver prices.

1761–1811: Mean of monthly data from Bauernfeind et al. (2001).

1812–1855: Calendar year prices are calculated as arithmetic mean of monthly prices. Rye: data from Pfister (2017), original in *Gulden* and *Kreuzer* per *Scheffel* from Seuffert (1857, 200–9). Wheat from Bauernfeind et al. (2001) (who also rely on Seuffert).

Osnabrück

Currency and volume conversion

Currency conversion follows *Currency Westphalia* until 1762; from 1763–1833 *Konventionstaler*; from 1834 the same as Berlin. *Osnabrücker Malter* converted following Witthöft (1993, 372).

Barley, oats, rye (1615–1861), wheat (1625–1861)

Calendar year prices for entire period extrapolated from *Martini* prices applying the general extrapolation rule (eq. S4). Raw data are in *Reichstaler* and *Schilling* (*Groschen*) per *Osnabrücker Malter* from Gerhard and Kaufhold (1990, 73–7, 125–8, 187–90, 239–41). Data are *Korntaxe* and treated as *Martini* prices (cf. Gerhard and Kaufhold, 1990, 76). Data for barley, oats, rye 1601–14 omitted because the absence of interannual variation for various years casts doubt on prices being market prices.

Paderborn

Currency and volume conversion

Currency conversion follows *Currency Westphalia* until 1762; from 1763–1813 *Konventionstaler*; from 1814 *Graumannscher Fuß* (16.704 g) (equivalent to Berlin). *Paderborner Scheffel* converted following Witthöft (1993, 375); *Berliner Scheffel* as for Berlin.

Barley, oats (1677–1871), rye, wheat (1641–1871)
Calendar year prices 1677–1810 (barley), 1677–1808 (oats), 1641–1808 (rye), and 1641–1802 (wheat) extrapolated from *Martini* prices. Extrapolation applies local time series relationships (Tables S10 and S11). Data from 1811 (barley), 1809 (oats, rye) and 1803 (wheat) until 1871 are calendar year prices from different sources: Until 1850 from Gerhard and Kaufhold (1990, 79–82, 130–2, 192–4, 244–5) in *Reichstalern* and *Mariengroschen* or *Gute Groschen* (since 1814; Gerhard and Kaufhold, 1990, 81) or *Silbergroschen* (since 1822) per *Paderborner Scheffel*; exception wheat: 1803–10 and 1814–21 in *Gute Groschen* but 1811–13 in *Mariengroschen*.

We omit earlier data for calendar years starting from 1780 for barley, oats, rye. The main reason is that the calendar year prices are almost identical to *Martini* prices in these earlier years but both series show the usual disagreement in later years. Furthermore, a visual comparison of rye to a neighboring market (Münster) shows that the Paderborn *Martini* price moves together with the Münster rye Martini price (which again show the usual difference to calendar year prices). These issues casts doubt on the early observations of annual averages for Paderborn.

We extend calendar year averages of later years from Gerhard and Kaufhold with further data in *Silbergroschen* per *Berliner Scheffel* from different sources: 1851–59 from Geheimes Staatsarchiv Berlin (b); 1860–4 averages of prices for May and October from Landesarchiv NRW; 1865–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72).

Martini prices are called *Domkapitularische Fruchttaxe* in Paderborn and were sampled by contemporaries during the time period between *Martini* and Easter (also referred to as winter prices by Gerhard and Kaufhold) to obtain a mean price which was used for monetizing peasant dues (Gerhard and Kaufhold, 1990, 81). *Martini* prices 1676–1833 (barley/oats) and 1640–1833 (rye/wheat) in *Reichstalern* and *Silbergroschen* per *Paderborner Scheffel* from Gerhard and Kaufhold (1990, 78–80, 129–31, 191–4, 242–4).

Quedlinburg

Currency and volume conversion
Conversion of currency and volumes follows Pfister (2017) with one difference. In addition to Pfister (2017), we adjusted prices for the Prussian vellon inflation 1808–21 using adjustment factors for Berlin (see Berlin).

Barley, oats, rye, wheat (1750–1855)
All prices in *Taler* per *Wispel* from Pfister (2017); original sources as follows. Calendar year prices until 1830 are calculated as averages of the prices in January and November from

Schulze (1965, 268–70) and Schulze (1967, 327–9) (price 1). Calendar year prices 1831–55 are extrapolated from another series (price 2), which is calculated as the mean of the minimum and maximum prices from Schulze (1967, 325–6). The relationships 1800–30 for extrapolation are as follows (dependent variable: price 1; explanatory variable: price 2).

Barley: $\alpha = 0.0579^{**}$, $\beta_1 = 0.8218^{***}$, $R^2 = 0.89$. Oats: $\alpha = 0.0602^{***}$, $\beta_1 = 0.7205^{***}$, $R^2 = 0.8$. Rye: $\alpha = 0.0221$, $\beta_1 = 0.9761^{***}$, $R^2 = 0.92$. Wheat: $\alpha = 0.04591$, $\beta_1 = 0.8977^{***}$, $R^2 = 0.91$.

Speyer

Currency and volume conversion
Conversion of currency and volumes as in Pfister (2017).

Rye (1530–1855), barley, oats (1821–1855)
Crop year prices for rye in *Denar* per *Malter* from Pfister (2017); originally from Elsas (1940, 550–4). Calendar year prices 1530–1820 were extrapolated from crop year prices using the general relationship in eq. S8 (parameters in Table S18), because prices from Elsas and Seuffert overlap for only one year. Calendar year prices 1821–55 from Seuffert (1857, 323); accessed through Pfister (2017).

Strasbourg

Currency and volume conversion
Conversion according to Allen (2001, 439).

Barley, oats, rye (1350–1681), wheat (1351–1681)
As in Pfister (2017) prices are from Allen's Database (2001) (Strasbourg file; Allen, 2001, 439; original in *Francs* per hectolitre from Hanauer, 1878, 91–101). Originally a free Imperial city, Strasbourg was conquered by the French crown in 1681. Hence, the data are used only until 1681. Some scattered observations before 1350 were omitted.

Trier

Currency and volume conversion
Currency conversion of *Albus* until 1796 as carried out by Irsigler (1988, 172-173); we use his prices in g Ag. Volume conversion until 1796 follows Irsigler (1988, 190). Volume and currency of data from 19th century converted as for Berlin.

Barley, oats (1550–1871), rye (1567–1871), wheat (1665–1871)

Prices until 1796 in g Ag per *Malter* from Irsigler (1988, 185–9). 1797–1815: gap. Spelt series (1550–1646) is not used because there is no overlap with wheat series so that extrapolation of missing values is impossible. Data for 19th century in *Gute Groschen/Silbergroschen* per *Berliner Scheffel* are from several sources: 1816–59 from Geheimes Staatsarchiv Berlin (b); rye price in 1824 and 1860–4 from Geheimes Staatsarchiv Berlin (a); 1865–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72).

Würzburg

Currency and volume conversion

Prices from Elsas (1936) and Christoforatou (2010) until 1799 are in *Rechengeld* (money of account). For the conversion we applied 168 *Denar* per *Fränkischen Gulden* (cf. discussion by Metz, 1990, 167–8, 309). *Denar* (=*Pfennig*) are converted to g Ag following Metz (1990, 436–43) as in Pfister (2017).

Currency conversion of prices 1815–55 applies the g Ag per *Denar* series for Munich from Pfister (2017) while 1 *Gulden* = 240 *Denar*.

Conversion of volumes until 1799 follows Elsas (1936, 157) (cf. Christoforatou, 2010, 294 and Verdenhalven, 1993, 30 for very similar rates); following Elsas, a different volume applies to oats. *Scheffel* for 1815–55 are converted following Witthöft (1993, 76).

Rye (1490–1855), wheat (1502–1855)

Crop year prices for rye and wheat (1463/1500–1799) from Pfister (2017); original source is Elsas (1936, 634–40) (prices in *Denar* per *Malter*).

Calendar year prices 1490/1502–1655 were obtained by applying the local time series relationship for rye and wheat in table S17 (Elsas' prices) to the crop year series. The relationship rests on calendar year averages based on monthly data from Christoforatou (2010) (see below).

1656–1777: prices for rye and wheat rely on arithmetic averages of monthly data in *Fränkischen Gulden* per *Malter* from Christoforatou (2010, 262–93). The gap in the series 1685–1700 is filled by extrapolation based on crop year prices from Elsas as above.

1778–99: Calendar year prices extrapolated from crop year prices from Elsas as above. 1800–1814 is a gap.

1815–55: Calendar year prices in *Gulden* and *Kreuzer* per *Scheffel* from Seuffert (1857, 282–3).

Oats (1464–1799)

Crop year prices (1462–1799) from Elsas (1936, 634–40); conversion to calendar year prices applies eq. S8 using the parameters from Table S18.

Xanten

Currency and volume conversion

Basic currency system: 1 *Taler* = 26 *solidi*. 1 *solidus* (=*Schilling*) = 12 *Pfennig* (= *denar*) (Beissel, 1889, 75, 91). Intrinsic silver content for *Schilling* from Metz (1990, 416–9). At the beginning, rates of intrinsic content are given for partly overlapping periods; the more recent rate is chosen. In case of gaps the earlier value was continued until a new value is recorded. From 1771 the value of Abel (1978, 302)(cited by Metz, 1990, 419–20, 425) is recorded. This is consistent with contemporary devaluation in Aachen and the intrinsic value of *solidus* in ca. 1827 given by (Beissel, 1889, 100).

Until about 1510 prices are given either in *albus*, *solidus* or a multitude of other coins, the latter becoming frequent from 1436. Main groups of other coins have been converted to silver equivalents using the exchange rates to *Schilling* given by Beissel (1889, 75–100). These include *Stüwer*/*Stuver*, *flems.* (=0.8 *Stüwer*), *albus* of Köln (see Köln for silver content), *Krummstert*, *flor. ren. curr.* (unambiguous rates only from 1480s), *flor. horn.*

From 1511–1585 prices are in *Albus*. Silver content is assessed using the rate of *Albus* per *denar* from Beissel (1889, 83–4). A value given for a particular year or period is continued until a new value is recorded. From 1586 values are in *Taler*; from 1826 Prussian currency in *Graumannscher Fuß*; see Berlin.

Volume conversion until 1799 follows Beissel (1889, 448); from 1800 as for Berlin: 1 *Berliner Malter* = 4 *Berliner Scheffel* (Beissel, 1889, 116).

Prices from Kopsidis (1994) are converted as for Berlin.

Barley (1370–1800), oats, rye, wheat (1370-1819)

Until 1800 calendar year prices from Beissel (1889, 118–33) refer to Xanten; from then on to Goch, a small town ca. 25 km west of Xanten. We do *not* extend data from Xanten with Goch (as Beussel did) but with data from Kopsidis (1994) (for Xanten) until 1819 (except barley: only until 1800). Reason: Fit of prices for Xanten and Goch is relatively weak. The correlations of rye from Beussel (until 1800 Xanten; from 1801 Goch) and rye from Kopsidis (Xanten) for different periods are: 0.9582 (1784–1800), 0.5696 (1801–1819).

Until 1549 prices are in diverse local currency per *Xantener Stiftsmalter*; 1550–85 in *Albus* per *Xantener Stadtmalter*; 1586–1799 in local *Taler* per *Xantener Stadtmalter* (rye, barley and oats 1550–1799 from Jacks' Database (checked with Beissel).

Prices for 1800–19 from Kopsidis (1994, table Va/6) (accessed through GESIS Köln, 2015) are in *Reichstaler* per *Scheffel* of Berlin.

Beissel gives two variants of oats prices until 1529: *avena* and *havena*. The price of the latter is higher than the price of the former. *Hafer* (=oats) seems to correspond to *havena*, so this variety is used.

SA3 Rules for extrapolating Martini and crop year prices to calendar year prices

This section documents the relationships between calendar year prices and *Martini* prices on the one hand and crop year prices on the other hand. We use the results to develop extrapolation rules for cases where calendar year prices are not available but *Martini* or crop year prices can be obtained.

SA3.1 Martini prices

We first discuss the empirical association between calendar year and *Martini* prices. Second we develop the extrapolation rules for *Martini* prices.

SA3.1.1 The relationship between calendar year and Martini prices

For several cities a part of the available data (for Osnabrück even all data) consist of so called *Martini* prices (see Section SA1.2 for the definition). To convert *Martini* into calendar year prices, we distinguish between two cases: (1) a local relationship between calendar year and *Martini* prices can be estimated; (2) it is impossible to estimate this relationship due to a lack of calendar year price data. The latter case is important for other researchers who face the problem of converting *Martini* prices to calendar year prices.

To derive the extrapolation rules, we collect an unbalanced panel of calendar year and *Martini* prices. We specify a reduced form relationship for an unbalanced panel as follows:

$$p_{it} = \alpha_0 + \beta_1 m_{it} + \beta_2 m_{it-1} + \sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S1}$$

Herein p_{it} denotes the calendar year price; m_{it} : *Martini* price, $i = 1, ..., I$ is the city index ($I = 8$ are included for which both type of prices are availabe)[20], t: time index for the respective calendar year (T ranges between 25 and 99), city dummy variables c_i to account for individual fixed effects quantified by α_i (α_0 is the constant; hence, we apply the least

[20]This includes the cities Emden, Göttingen, Halle, Hanover, Lüneburg, Minden, Münster, Paderborn.

Table S3: Relationship calendar year price and *Martini* price, rye 1692–1863 (part A)

	(1) Pooled	(2) city FE	(3) as (2), + trend	(4) as (2), no lag	(5) as (2), only 1st lag	(6) as (2), add 2nd lag
Intercept	0.0129	0.0031	−0.0527	−0.0332	−0.0442	0.0306
	(0.0137)	(0.0213)	(0.0536)	(0.0889)	(0.0859)	(0.0214)
Martini price	0.4885***	0.4833***	0.4813***	0.7812***		0.4767***
	(0.0306)	(0.0305)	(0.0310)	(0.0356)		(0.0324)
Lag 1 Martini price	0.5024***	0.5005***	0.4984***		0.8025***	0.5557***
	(0.0447)	(0.0453)	(0.0446)		(0.0436)	(0.0530)
Year			0.0001	0.0003	0.0003*	
			(0.0001)	(0.0002)	(0.0002)	
Lag 2 Martini price						−0.0871***
						(0.0252)
R^2	0.8978	0.9108	0.9111	0.7710	0.7798	0.9155
Adj. R^2	0.8973	0.9089	0.9090	0.7664	0.7754	0.9135
Num. obs.	442	442	442	460	459	432

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price. SCC standard errors in (). FE: fixed effects.

squares dummy variable estimator), and error term u_{it}. The parameters of interest, β_1 and β_2, are estimated for four data sets, one for each commodity (barley, oats, rye, wheat).

The idea behind this setup is that *Martini* prices are driven by the harvest of the present year, whereas annual prices are driven by the harvests of the present and the past year. Hence, if the estimate of eq. S1 obtains a good fit, calendar year prices can be approximated with a weighted average of present year and past year Martini prices (cf. Phelps Brown and Hopkins, 1959, 31; Bateman, 2011, 451). In the case of rye, parameter estimates turn out as roughly $\beta_1 = 0.5$ and $\beta_2 = 0.5$ in the regression according to equation (S1) (Table S3, model 2, eq. S1; similar results obtain if no city fixed effects are included, model 1). The p-values are based on spatial correlation consistent (SCC) standard errors (Driscoll and Kraay, 1998; Millo, 2017). The latter are also robust to (cross-) serial correlation.

The model fit is lower if only calendar year and contemporaneous *Martini* price are used as predictor (model 4); a similar result obtains, if only the first lag of the *Martini* price enters on the right-hand-side (model 5). An additional second lag turns out as statistically significant but does not improve the model fit substantially (model 6). The Bayesian Information Criterion (BIC) is lower with only one lag (for all cereals), and hence, we opted for the simpler specification.[21]

To pin down a simple extrapolation rule for rye prices in the case when no calendar year prices are available but *Martini* prices can be obtained, we develop and test the following restrictions. First, we tested the hypothesis that $\beta_1 + \beta_2 = 1$. This hypothesis was not

[21] E.g., for rye, model 2 (Table S3), the BIC is -1177; with 2 lags, model 6, the BIC is -1160.

rejected at any conventional level ($p \leq 0.10$) irrespective of whether city dummy variables were included in the model, except for oats. In the latter case, the hypothesis was only *not* rejected, if city dummy variables were included in the model.[22]

Second, we tested the hypothesis that $\beta_1 = 0.5$ and $\beta_2 = 0.5$. Again, this hypothesis was not rejected at any conventional level irrespective of whether city dummy variables were included in the model, except for oats (where we failed to reject after inclusion of city dummy variables).

We then specified a model which incorporates the following restriction: $\beta_2 = 1 - \beta_1$:

$$p_{it} = \alpha_0 + \beta_1 m_{it} + (1 - \beta_1) m_{it-1} + \sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S2}$$

The latter eq. can be written as:

$$p_{it} - m_{it-1} = \alpha_0 + \beta_1 [m_{it} - m_{it-1}] + \sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S3}$$

The left-hand-side is the dependent variable of the restricted model. In the estimated restricted model (with/without city dummy variables), β_1 is close to 0.5 and we cannot reject the hypothesis that $\alpha_0 = \alpha_0$ and $\beta_1 = 0.5$ at any conventional level. The results for the restricted models are shown in Table S4.

Table S4: Relationship calendar year price and *Martini* price, rye 1692–1863 (part B)

	(7) restr. model pooled	(8) as (7) + city FE	(9) as (2), w/o 5% highest/lowest obs.	(10) FD estimator	(11) as (2), in logs	(12) as (2), syn. Martini prices
Intercept	0.0085**	−0.0067	0.0136	0.0003	−0.0058	−0.0066
	(0.0040)	(0.0139)	(0.0233)	(0.0029)	(0.0234)	(0.0078)
FD Martini price	0.4928***	0.4908***				
	(0.0341)	(0.0341)				
Martini price			0.4590***	0.5058***	0.4860***	0.5257***
			(0.0270)	(0.0350)	(0.0272)	(0.0178)
Lag 1 Martini price			0.4991***	0.5639***	0.5026***	0.4620***
			(0.0358)	(0.0514)	(0.0263)	(0.0269)
R^2	0.6200	0.6680	0.8604	0.7222	0.9374	0.9474
Adj. R^2	0.6191	0.6619	0.8568	0.7210	0.9361	0.9468
Num. obs.	442	442	356	434	442	1015

***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Dependent variable: difference of calendar year price and 1st lag of Martini price (restricted models 7,8; see text for details), calendar year price (models 9,12), 1st difference of calendar year price (model 10), log(calendar year price) (model 11). SCC standard errors in (). Model 2 is in Table S1. Model 12 refers to period 1490–1863. FD: first difference; FE: fixed effects; restr.: restricted, syn.: synthetic; w/o: without.

In addition, Table S4 contains several other specifications as robustness checks that yield overall similar results. Dropping very small and very large calendar year prices (model 9) and

[22] Additionally, we tested the following hypothesis. We estimated a pooled model without city dummy variables and tested $\alpha_0 + \beta_1 + \beta_2 = 1$. This hypothesis was rejected for barley and oats; it was not rejected for rye and wheat.

applying the first-differenced estimator (Wooldridge, 2009, 458) yields roughly similar results. Furthermore, increasing the sample by synthetic[23] *Martini* prices, which we calculated from monthly data (model 12), does not alter the coefficients very much either.

We also considered an alternative version of eq. S1, where the dependent and explanatory variables enter in logs. We followed the approach in Wooldridge (2009, 213) to calculate an R^2 for a log-model which can be compared to the R^2 from the regression based on the model in levels (eq. S1). The fit of the log-model (for the prices in levels) was practically the same as for the model in levels: $R^2 = 0.91$. Hence, we worked with the dependent variable in levels. This also simplifies extrapolation, because with a log-model additional adjustment factors are necessary for predicting the variable of interest in levels (Wooldridge, 2009, 210–2). The results for the log-model are shown in Table S4 (model 11).

We ran all 12 specifications for the other grain types as well (barley, oats, wheat). We report only the two most relevant specifications for each cereal in Table S5, because the results were very similar to those for rye.

Table S5: Relationship calendar year price and Martini price for other cereals.

	Barley		Oats		Wheat	
	(1) Pooled	(7) restr. pooled	(1) Pooled	(7) restr. pooled	(1) Pooled	(7) restr. pooled
Intercept	−0.0024	0.0161***	0.0103**	0.0217***	0.0155	0.0123**
	(0.0102)	(0.0034)	(0.0045)	(0.0026)	(0.0129)	(0.0049)
Martini price	0.5304***		0.5439***		0.4893***	
	(0.0619)		(0.0364)		(0.0251)	
Lag 1 Martini price	0.5239***		0.5083***		0.5057***	
	(0.0599)		(0.0380)		(0.0425)	
FD Martini price		0.5041***		0.5185***		0.4918***
		(0.0590)		(0.0370)		(0.0330)
R^2	0.8901	0.5350	0.8469	0.4739	0.9023	0.5804
Adj. R^2	0.8895	0.5339	0.8462	0.4727	0.9018	0.5794
Num. obs.	429	429	432	432	449	449

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price (model 1); difference of calendar year price and 1st lag of Martini price (restricted model 7; see text for details). SCC standard errors in (). FD: first difference; restr.: restricted.

Additionally, we performed time series regressions for each city and commodity. The model obtains directly after dropping individual index and city dummy variables from eq.

[23] To increase the number of observations, synthetic *Martini* prices were calculated as averages of monthly prices of November and December for the cities Cologne, Munich, Nuremberg, Würzburg.

S1: $p_t = \alpha_0 + \beta_1 m_t + \beta_2 m_{t-1} + u_t$. These time series regressions show that the weights are sometimes different for particular cities. However, interpretation of the results is sometimes hampered by very large standard errors due to the limited number of observations. We report these results in Tables S6– S11. Whenever we have both calendar year and *Martini* prices for a city, we use the parameters of these time series regression results to fill gaps in calendar year series. This allows preserving as much local information as possible.

Table S6: Time series relationship calendar year price and Martini price (part A)

	Emden				Göttingen	
	rye	barley	oats	wheat	wheat	rye
	1797–1850	1797–1825	1797–1850	1797–1825	1765–1863	
Intercept	0.0177	−0.0136	0.0406**	0.0149	−0.0337	0.0036
	(0.0214)	(0.0255)	(0.0169)	(0.0408)	(0.0304)	(0.0267)
Martini price	0.3970***	0.4799***	0.5439***	0.6157***	0.5571***	0.5390***
	(0.1107)	(0.0317)	(0.0466)	(0.0880)	(0.0607)	(0.0511)
Lag 1 Martini price	0.5640***	0.5426***	0.3291***	0.3243***	0.5012***	0.4544***
	(0.0975)	(0.0406)	(0.0407)	(0.0862)	(0.0745)	(0.0741)
R^2	0.9220	0.9296	0.8278	0.9160	0.9100	0.9063
Adj. R^2	0.9149	0.9222	0.8114	0.9072	0.9081	0.9043
Num. obs.	25	22	24	22	99	99

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price. Newey and West (1987) standard errors in ().

Table S7: Time series relationship calendar year price and Martini price (part B)

	Göttingen		Halle			
	barley	oats	wheat	rye	barley	oats
	1765–1850		1692–1834			
Intercept	0.0005	−0.0018	0.0281***	0.0129	0.0126*	0.0147**
	(0.0378)	(0.0152)	(0.0067)	(0.0110)	(0.0065)	(0.0058)
Martini price	0.5521***	0.5947***	0.4698***	0.5142***	0.4970***	0.4602***
	(0.0956)	(0.0774)	(0.0200)	(0.0172)	(0.0305)	(0.0674)
Lag 1 Martini price	0.4729***	0.5268***	0.4590***	0.4273***	0.4600***	0.5012***
	(0.1376)	(0.0816)	(0.0241)	(0.0478)	(0.0398)	(0.0637)
R^2	0.8082	0.7824	0.9593	0.9461	0.9081	0.8964
Adj. R^2	0.8036	0.7772	0.9583	0.9448	0.9057	0.8938
Num. obs.	86	86	82	83	82	82

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price. Newey and West (1987) standard errors in ().

Table S8: Time series relationship calendar year price and Martini price (part C)

	Hanover				Lüneburg	
	wheat	rye	oats	barley	barley	oats
		1765–1850			1766–1850	
Intercept	0.0052	0.0061	0.0055	−0.0198	−0.0428	0.0200
	(0.0290)	(0.0226)	(0.0105)	(0.0179)	(0.0257)	(0.0195)
Martini price	0.3956***	0.4067***	0.4968***	0.4763***	0.7640***	0.4756***
	(0.0498)	(0.0376)	(0.0540)	(0.0511)	(0.1303)	(0.0507)
Lag 1 Martini price	0.6311***	0.6116***	0.5868***	0.6106***	0.4139**	0.5164***
	(0.0692)	(0.0676)	(0.0740)	(0.0780)	(0.1540)	(0.0863)
R^2	0.8795	0.9184	0.8762	0.8914	0.8864	0.8898
Adj. R^2	0.8761	0.9160	0.8726	0.8883	0.8791	0.8826
Num. obs.	73	72	72	73	34	34

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price. Newey and West (1987) standard errors in ().

Table S9: Time series relationship calendar year price and Martini price (part D)

	Lüneburg		Minden			
	rye	wheat	barley	oats	rye	wheat
	1766–1850			1777–1850		
Intercept	−0.0421	0.0680	−0.0100	0.0020	0.0205	0.0485*
	(0.0333)	(0.0414)	(0.0248)	(0.0136)	(0.0228)	(0.0260)
Martini price	0.5663***	0.4045***	0.5330***	0.6126***	0.4155***	0.4178***
	(0.1245)	(0.0533)	(0.0572)	(0.0770)	(0.0669)	(0.0806)
Lag 1 Martini price	0.5507***	0.5173***	0.5661***	0.4699***	0.5887***	0.5452***
	(0.1498)	(0.0687)	(0.0994)	(0.0929)	(0.0812)	(0.0738)
R^2	0.8366	0.8789	0.8634	0.7117	0.8670	0.8717
Adj. R^2	0.8245	0.8711	0.8591	0.7027	0.8628	0.8677
Num. obs.	30	34	67	67	67	67

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price. Newey and West (1987) standard errors in ().

Table S10: Time series relationship calendar year price and Martini price (part E)

	Münster				Paderborn	
	barley	oats	rye	wheat	barley	oats
		1816–1863			1811–1833	1809–1833
Intercept	0.0780**	0.0542**	0.0294	0.0380	−0.0061	0.0077
	(0.0356)	(0.0223)	(0.0311)	(0.0601)	(0.0377)	(0.0193)
Martini price	0.3474***	0.3358***	0.5519***	0.5160***	0.7250***	0.7580***
	(0.0974)	(0.0709)	(0.0590)	(0.0552)	(0.1795)	(0.1569)
Lag 1 Martini price	0.5655***	0.6054***	0.5073***	0.5475***	0.3743**	0.3695**
	(0.0780)	(0.0705)	(0.0654)	(0.0660)	(0.1495)	(0.1437)
R^2	0.8400	0.8353	0.9192	0.8937	0.8051	0.7035
Adj. R^2	0.8318	0.8268	0.9149	0.8881	0.7856	0.6766
Num. obs.	42	42	41	41	23	25

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price. Newey and West (1987) standard errors in ().

Table S11: Time series relationship calendar year price and Martini price (part F)

	Paderborn	
	rye	wheat
	1809–1833	1803–1833
Intercept	0.0679	−0.0011
	(0.0580)	(0.0465)
Martini price	0.4559***	0.5283***
	(0.1092)	(0.0795)
Lag 1 Martini price	0.3434***	0.4245***
	(0.0788)	(0.0877)
R^2	0.6878	0.7544
Adj. R^2	0.6594	0.7369
Num. obs.	25	31

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price. Newey and West (1987) standard errors in ().

SA3.1.2 A simple extrapolation rule when calendar year prices are not available

Based on the panel regression estimates, we derive an extrapolation rule for each grain type which can be applied if no calendar year prices are available to estimate a local relationship. The extrapolation rules always use equally weighted contemporaneous and lagged *Martini*

prices ($\beta_{1r} = \beta_{2r} = 0.5$) but differ in the constant α_r; index r indexes the parameters applied for extrapolation across types of grain. The extrapolation rule is:

$$p_t = \alpha_r \cdot \bar{m}_{\text{time series Martini}} + 0.5 \cdot m_t + 0.5 \cdot m_{t-1}. \tag{S4}$$

Note that $\alpha_r \neq \alpha_0$ from the restricted panel regression (eq. S3; Tables S4 and S5, model 7); α_r is the constant from the panel regression in percent. Thus, in the case of rye, $\alpha_r = \alpha_1 = 0.0185$ and calendar year prices are extrapolated by adding 1.85% of the average of the given *Martini* price time series to the equally weighted given contemporaneous and lagged *Martini* prices.

To derive the factors α_r the constant in the panel regression α_0 is related to the average *Martini* price of the full sample used in the panel regressions. This is necessary because the constant in the regression is not independent of the level of the *Martini* prices used in the panel regression. The constant α_0 quantifies a markup on the *Martini* price and is measured in the unit of the dependent variable, that is, in grams silver per litre. To allow application of the rule to a sample of *Martini* prices, where calendar year prices are unknown, the constant used for extrapolation must be scaled to the *Martini* prices at hand. We do this by setting $\alpha_r = \alpha_0 / \bar{m}_{\text{panel}}$ where $\bar{m}_{\text{panel}}$ denotes the arithmetic mean over all cities and years which are included in the restricted model 7. In the case of rye this yields $\alpha_1 = 0.0185$. For other cereals the values are: $\alpha_2 = 0.0467$ (barley) and $\alpha_3 = 0.0202$ (wheat). We then multiply this factor α_r with the mean of the available *Martini* price series.[24]

How sensitive is the rule to variations of the sample and how well does it work out-of-sample? To answer these questions, we perform a cross-validation, a method used to evaluate the out-of-sample performance of regression models (e.g., Roberts et al., 2013, 241). In our context, we drop the time series of one city, the test city (or testing set), from the commodity specific panel data set, which yields the training set. We then train model 7, the restricted model without city dummy variables, on this panel data set (that is, without the test city). We save the parameters α_0, β_1 and β_2. We calculate the constant α_r in percent as described above and round β_1 and β_2 to two digits, which yields β_{1r} and β_{2r}. Thus, the parameters α_r, β_{1r} and β_{2r} are not based on observations from the test city. We then use the extrapolation rule (eq. S4) and predict a calendar year price series using the *Martini* prices of the test city. As a final step we calculate the R^2 of predicted and actual calendar year time series for the test city. We repeat this procedure for all 32 time series with overlapping information.

On average, the rule achieves an out-of sample fit of $R^2 = 0.85$ (without oats: 0.87). The out-of-sample performance is very similar for all cereals except for oats, which performs

[24] We do not use or recommend the rule for oats ($\alpha_4 = 0.1024$) due to the relatively low out-of-sample performance.

worse: $R^2 = 0.81$. This is consistent with the observation that the restriction $\beta_1 = 0.5$ and $\beta_2 = 0.5$ was rejected for oats in the pooled panel regression without city dummy variables (model 1). Hence, we do not recommend to apply the rule for oats. The parameters from the trained models are close to 0.5: $0.47 \leq \beta_{1r} \leq 0.53$ (equivalent for β_{2r}). That is, the parameters of the proposed extrapolation rule are not sensitive to variations of the sample. The model predicts out-of-sample time series fairly well.

It should be noted that the in-sample fit of the unrestricted time series regression is—unsurprisingly—better (on average $R^2 = 0.86$) than what our rule can achieve out-of-sample; but the difference is not large. Still, whenever overlapping information on calendar year and *Martini* prices are available a local relationship should be used to preserve local information. To summarize, our extrapolation rule performs not much worse than estimating a local relationship and far better than treating *Martini* prices as if they were calendar year prices.

For our sample, we proceed as follows. We use the extrapolation rule (eq. S4) for those series for which we have no or insufficient parallel information on annual prices (Osnabrück and Celle[25]), and the results of the unrestricted model estimated separately for each market and cereal where both calendar year and *Martini* prices are available.

To close this discussion, we ask how relevant the data transformation is for market integration studies. In fact, price synchronization between cities might be underestimated, if we use *Martini* prices instead of extrapolated calendar year prices. Table S12 shows two examples, the rye series for Halle and Münster. The correlation of the extrapolated calendar year prices with the rest of the stable sample is on average higher compared to simply using contemporaneous *Martini* prices instead (Halle *Martini*: $r = 0.59$ and calendar year price $r = 0.64$; Münster: $r = 0.49$ and $r = 0.59$). In a very few cases price synchronization is also overestimated if *Martini* prices are used and the correlation of converted calendar year prices with the rest of the stable sample decreases (Paderborn, Xanten).

[25]We use the rule also for Celle because only 11 observations would be available to estimate a local relationship.

Table S12: Improvement of inter-urban correlation by extrapolation to calendar year prices

	Halle *Martini*	Halle calendar year extrapolated	Münster *Martini*	Münster calendar year extrapolated
Augsburg	0.66	0.70	0.26	0.32
Berlin	0.78	0.85	0.50	0.64
Braunschweig	0.60	0.64	0.48	0.63
Cologne	0.56	0.56	0.63	0.69
Dresden	0.79	0.81	0.40	0.52
Gdansk	0.49	0.55	0.54	0.66
Göttingen	0.55	0.60	0.38	0.66
Halle	1.00	1.00	0.43	0.60
Hamburg	0.48	0.64	0.45	0.70
Munich	0.54	0.66	0.20	0.34
Münster	0.47	0.60	1.00	1.00
Osnabrueck	0.64	0.69	0.74	0.87
Paderborn	0.55	0.55	0.75	0.70
Wuerzburg	0.66	0.75	0.37	0.49
Xanten	0.44	0.33	0.79	0.41
Average correlation	0.59	0.64	0.49	0.59

Pearson correlations of price series in column with all price series in row. Rye, stable sample, 1651–1790. Columns 1 and 3 contain correlations of *Martini* prices with the series in the rows; columns 2 and 4 the correlations of the respective calendar year prices extrapolated from the *Martini* prices with the series in the rows. For both cities all 140 observations are available as *Martini* and calendar year price.

SA3.2 Crop year prices

To employ grain prices as an indicator of harvest fluctuations, earlier scholarship has sometimes aggregated prices at the level of crop years (e.g., Elsas, 1933). We first discuss the empirical association between calendar year and crop year prices. Second, we develop the extrapolation rules for crop year prices.

SA3.2.1 The relationship between calendar and crop year prices

The underlying concept of the crop year is that the supply of the recent harvest had a considerable impact on the price. According to this view, the respective time base for the annual mean should refer to the period (that is, the year) between two harvests. In this way, the resulting time series would exhibit the fluctuations resulting from different harvests in a better way. The validity of this argument depends on the degree of market integration, however. The more a market is integrated with others, the less important local supply is in determining the price.

We find two versions of the crop year in the literature. Elsas defined the crop year from August to July of the following year. (Elsas, 1933, 224–5; 1936, 92–3). Bauernfeind (1993, 63) follows this rule. Data from Elsas (1936; 1940; 1949) used by this study include: Augsburg, Frankfurt, Munich, Speyer, Würzburg, and Leipzig (until 1820). In contrast, data for Aachen and Düren (both rye) refer to July until June as the crop year (reported by Rahlf, 1996).

In order to explore the relationship between crop year data and calendar year prices we employ a model that is analogous to the one that we used in the case of Martini prices above (eq. S1):

$$p_{it} = \alpha_0 + \beta_1 h_{it} + \beta_2 h_{it-1} + \beta_3 h_{it-2} + \sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S5}$$

Herein p_{it} denotes the calendar year price, h_{it} the crop year or synonymously the harvest year price; i indexes cities and t years. As in the eq. for the *Martini* prices (S1), c_i denotes city dummy variables. We draw on cities where both crop and calendar years are available or where we can calculate both types of data from monthly data. For rye, there are seven cities (Augsburg, Berlin, Cologne, Hanover, Munich, Nuremberg, Würzburg; see section SA2). Again, we focus on rye and present results for the other cereals only briefly. The resulting tables are organized like those for the *Martini* prices. What follows refers to the crop year from August to July.

The most important difference with respect to the results obtained for *Martini prices* is that it proved necessary to include a second lag. The hypothesis that $\beta_1 + \beta_2 = 1$ was rejected in a model with one lag (oats and rye), while $\beta_1 + \beta_2 + \beta_3 = 1$ was not rejected at $p \leq 0.10$ (based on SCC standard errors; irrespective of whether city dummy variables were included or not). In addition, the BIC was always lower for the model with contemporaneous crop year price and two lags of the latter compared to the model with one lag. E.g., for rye model 2 (one lag), BIC = −3910; for model 6 (two lags) BIC = −3936 (Table S13). In all models the first lag is the most important predictor and the additional second lag in model 6 enters with a negative sign.

Table S13: Relationship between calendar and crop year price, rye 1490–1871 (part A)

	(1) Pooled	(2) city FE	(3) as (2), + trend	(4) as (2), no lag	(5) as (2), only 1st lag	(6) as (2), add 2nd lag
Intercept	−0.0106***	0.0129***	0.0274***	0.0104	0.0292**	0.0191***
	(0.0023)	(0.0046)	(0.0058)	(0.0151)	(0.0146)	(0.0046)
Crop y. price	0.4253***	0.4242***	0.4268***	0.8573***		0.4107***
	(0.0120)	(0.0119)	(0.0120)	(0.0212)		(0.0127)
Lag 1 Crop y. price	0.6033***	0.6049***	0.6076***		0.9244***	0.6687***
	(0.0119)	(0.0119)	(0.0120)		(0.0169)	(0.0181)
Year			−0.0000***	0.0001***	0.0001***	
			(0.0000)	(0.0000)	(0.0000)	
Lag 2 Crop y. price						−0.0715***
						(0.0104)
R^2	0.9742	0.9754	0.9755	0.8190	0.8941	0.9784
Adj. R^2	0.9741	0.9752	0.9753	0.8174	0.8932	0.9782
Num. obs.	902	902	902	915	916	885

***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Dependent variable: calendar year price. SCC standard errors in (). FE: fixed effects. The crop year series for Augsburg ends in 1800, because of a change in the underlying source.

Model fit in Table S13 is generally higher than for estimates based on Martini prices (Table S3). This is because crop year prices are already annual average prices although with the wrong time base. Hence, crop year prices contain more information than *Martini* prices, which translates into a better fit of calendar year regression estimates.

Turning to the fit of individual regressions in Table S13, choosing the lagged crop year price as predictor yields a higher model fit than employing the contemporaneous crop year price (models 4 and 5, Table S13). This finding demonstrates that considering crop year prices as proxies for contemporaneous calendar year prices is not appropriate. Although this result is only a correlation, one could conjecture that the strong lag results from the high labour intensity of the threshing of grain prior to mechanization. It lasted considerable time until the current harvest became available for consumption so that storage of the former crop and expectations rather than the actual new harvest influenced the calendar year price (e.g., Brunt and Cannon, 2017).

In the model with two lags (model 6, Table S13), the calendar year price can be regarded as a weighted average, where the weights sums to one. To determine the weights empirically, we estimated a restricted model ($\beta_2 = 1 - \beta_1 - \beta_3$):

$$p_{it} = \alpha_0 + \beta_1 h_{it} + (1 - \beta_1 - \beta_3) h_{it-1} + \beta_3 h_{it-2} + \sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S6}$$

The latter eq. can be written as:

$$p_{it} - h_{it-1} = \alpha_0 + \beta_1[h_{it} - h_{it-1}] + \beta_3[h_{it-2} - h_{it-1}] + \sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \quad \text{(S7)}$$

The results for model 8 in Table S14 directly refer to eq. S7.

Table S14: Relationship between calendar and crop year price, rye 1490–1871 (part B)

	(7) restr. model pooled	(8) as (7) + city FE	(9) as (6), w/o 5% high. est/lowest obs.	(10) FD estimator	(11) as (6), in logs	(12) as (6), pooled
Intercept	0.0001	0.0222***	0.0266***	0.0000	0.0755***	−0.0033
	(0.0008)	(0.0040)	(0.0060)	(0.0008)	(0.0117)	(0.0023)
FD crop y.	0.4083***	0.4069***				
	(0.0120)	(0.0120)				
Lag 2 - lag 1	−0.0762***	−0.0756***				
(crop y.)	(0.0103)	(0.0103)				
Crop y.			0.4185***	0.4657***	0.4082***	0.4127***
			(0.0182)	(0.0187)	(0.0116)	(0.0128)
Lag 1 crop y.			0.6142***	0.6714***	0.6713***	0.6678***
			(0.0374)	(0.0194)	(0.0140)	(0.0181)
Lag 2 crop y.			−0.0538***	−0.0686***	−0.0659***	−0.0715***
			(0.0152)	(0.0149)	(0.0088)	(0.0105)
R^2	0.7881	0.7977	0.9531	0.8631	0.9799	0.9774
Adj. R^2	0.7876	0.7959	0.9525	0.8626	0.9797	0.9773
Num. obs.	885	885	739	878	885	885

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: difference of calendar year price and 1st lag of crop year price (restricted models 7,8; see text for details), calendar year price (models 9,12), 1st difference of calendar year price (model 10), log(calendar year price) (model 11). Model 6 is in Table S13. SCC standard errors in (). FD: first difference; FE: fixed effects; restr.: restricted, syn.: synthetic; w/o: without.

The results are very similar for log prices (model 11, Table S14). As for the *Martini* prices, we followed the approach in Wooldridge (2009, 213) to calculate an R^2 for a log-model which can be compared to the R^2 from the regression based on the model in levels. The fit of the log-model (for the prices in levels) was practically the same as for the model with two lags in levels: $R^2 = 0.98$. Hence, we used the dependent variable in levels.

We ran all 12 specifications for grain types apart from rye (barley, oats, wheat[26]). We report only the two most relevant specifications for each cereal in Table S15, because the

[26]Two influential observations (Cook's distance > 1; Kleiber and Zeileis, 2008, 96–100) were excluded for wheat from Munich. These influential observations include the crop year prices from the *Heilig Geist* hospital for the years 1713 and 1805, for which Elsas (1936, 270) excluded some underlying observations from the here used crop year averages for 1713 and 1805.

results were very similar to those for rye, although small differences in the particular weights are evident.

Table S15: Relationship calendar year price and crop year price for other cereals.

	Barley		Oats		Wheat	
	(12) Pooled	(7) restr. pooled	(12) Pooled	(7) restr. pooled	(12) Pooled	(7) restr. pooled
Intercept	0.0071	0.0000	−0.0025	−0.0005	0.0063	0.0027**
	(0.0049)	(0.0009)	(0.0016)	(0.0003)	(0.0050)	(0.0013)
Crop y. price	0.4201***		0.3858***		0.3406***	
	(0.0325)		(0.0191)		(0.0224)	
Lag 1 crop y. price	0.6996***		0.7136***		0.7376***	
	(0.0292)		(0.0260)		(0.0239)	
Lag 2 crop y. price	−0.1417***		−0.0899***		−0.0848***	
	(0.0278)		(0.0163)		(0.0146)	
FD crop y. price		0.4325***		0.3817***		0.3435***
		(0.0349)		(0.0175)		(0.0211)
Lag 2 - lag 1		−0.1298***		−0.0942***		−0.0814***
(crop y.)		(0.0211)		(0.0146)		(0.0133)
R^2	0.9693	0.7488	0.9787	0.7424	0.9700	0.6117
Adj. R^2	0.9691	0.7478	0.9786	0.7415	0.9698	0.6103
Num. obs.	518	518	564	564	594	594

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price (model 12); difference of calendar year price and 1st lag of crop y. price (restricted model 7; see text for details). SCC standard errors in (). FD: first difference; restr.: restricted.

In Tables S16–S17 we report the results for individual local time series regressions. These results were used to extrapolate calendar year prices when both calendar year and crop year prices were available.

Table S16: Time series relationship calendar and crop year price (part A)

	Augsburg rye	Munich oats	Munich rye	Munich barley	Munich wheat
	1750–1799	1692–1773	1693–1773	1698–1820	1691–1797
		Kammerrechnungen		*Heilig-Geist*	
Intercept	0.0126	−0.0182**	−0.0196	0.0646***	0.0328
	(0.0227)	(0.0079)	(0.0135)	(0.0203)	(0.0446)
Crop y.	0.5336***	0.5743***	0.4421***	0.3626***	0.1415*
	(0.0525)	(0.0487)	(0.0434)	(0.0803)	(0.0722)
Lag 1 crop y.	0.5252***	0.5769***	0.6704***	0.7705***	0.8139***
	(0.0914)	(0.0481)	(0.0576)	(0.0584)	(0.1787)
Lag 2 crop y.	−0.0395	−0.0568	−0.0526	−0.3228***	0.0642
	(0.0311)	(0.0500)	(0.0524)	(0.0373)	(0.0937)
R^2	0.9450	0.8885	0.9617	0.9660	0.8788
Adj. R^2	0.9408	0.8843	0.9599	0.9619	0.8694
Num. obs.	44	82	69	29	43

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price. Newey and West (1987) standard errors in ().

Table S17: Time series relationship calendar and crop year price (part B)

	Würzburg rye	Würzburg wheat	Munich oats	Munich rye
	1656–1777	1658–1776	1691–1799	1692–1800
	Elsas		*Heilig-Geist*	
Intercept	−0.0083	−0.0074	0.0142	0.0231
	(0.0130)	(0.0244)	(0.0247)	(0.0343)
Crop y.	0.4385***	0.2696***	0.6287***	0.4680***
	(0.0507)	(0.0357)	(0.1720)	(0.1265)
Lag 1 crop y.	0.5218***	0.6563***	0.2023	0.5475***
	(0.0595)	(0.0514)	(0.2235)	(0.1331)
Lag 2 crop y.	0.0445	0.0968*	0.1727	−0.0296
	(0.0316)	(0.0498)	(0.2357)	(0.0971)
R^2	0.9242	0.9115	0.8069	0.9503
Adj. R^2	0.9219	0.9071	0.7827	0.9379
Num. obs.	103	65	28	16

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: calendar year price. Newey and West (1987) standard errors in (). The latter two relationships from the *Heilig-Geist* hospital are not part of the panel regressions.

SA3.2.2 A simple extrapolation rule when calendar year prices are not available

As for the extrapolation based on *Martini* prices, regional specific parameters from time series regressions are preferred. Only if no local calendar year prices were available, we applied

parameters from the commodity-specific panel regressions. In contrast to the *Martini* prices, a common set of parameters does not work for crop year prices. The hypothesis that average coefficients of the restricted pooled regressions (model 7) apply for each single commodity is rejected for rye and wheat at least at the 10% level in the unrestricted model 12.[27] Thus, we applied the the commodity specific parameters from the restricted pooled model 7 rounded to two digits (Table S18). These parameters cannot be rejected in the commodity-specific models 6 and 12. Neither can we reject the hypothesis that $\alpha_0 = 0$ and $\beta_1 = \beta_{1r}$ and $\beta_3 = \beta_{3r}$ in model 7; index r denotes the rounded parameters from the restricted model 7.[28] Thus, the equation for extrapolation that is applied if no local time series relationship is available is:

$$p_{it} = \beta_{1r} \cdot h_{it} + \beta_{2r} \cdot h_{it-1} + \beta_{3r} \cdot h_{it-2}. \tag{S8}$$

Table S18: Parameters for extrapolation of calendar from crop year prices (August–July)

Parameter	barley	oats	rye	wheat
β_{1r}	0.43	0.38	0.41	0.34
β_{2r}	0.70	0.71	0.67	0.74
β_{3r}	-0.13	-0.09	-0.08	-0.08

Note: β_{1r} and β_{3r} are the rounded parameters from the restricted model 7 (Tables S14 and S15); $\beta_{2r} = 1 - \beta_{1r} - \beta_{3r}$.

We explored the robustness of the above extrapolation rule using a cross-validation for all 20 time series (like for the *Martini* prices). The achieved out-of-sample fit of actual and predicted values is $R^2 = 0.94$. The ranges of the estimated parameters from the trained models are very similar to the values reported in Table S18. E.g., for rye the values are: $0.40 \leq \beta_{1r} \leq 0.42$; $0.64 \leq \beta_{2r} \leq 0.68$; $-0.09 \leq \beta_{3r} \leq -0.06$. Using the contemporaneous crop year price to predict the calendar year price achieves a considerably worse out-of-sample fit: $R^2 = 0.71$.

Recall that the foregoing analysis refers to the crop year running from August through July. Two additional data sets relate to prices in crop years beginning in July and ending in June, namely, those for Aachen and Düren (Rahlf, 1996). Unfortunately, it proved impossible to derive a robust extrapolation rule for this variant of the crop year. Therefore, data for Düren and Aachen (before 1784) cannot be used in this study.

[27] We tested both arithmetic and weighted averages. In the latter case, the weights were chosen as the share of the number of observations from the commodity specific panel regression (model 12) in the total number of observations from all four regressions. The weighted average gives a larger weight to the parameters for rye for which the largest number of observations are available. But the parameters were still rejected for wheat in the unrestricted model 6 included city dummy variables at $p < 0.10$.

[28] Model 7 for wheat (Table S15) is the only case were a statistically significant intercept was estimated. Nevertheless the hypothesis that $\alpha_0 = 0$ and $\beta_1 = \beta_{1r}$ and $\beta_3 = \beta_{3r}$ cannot be rejected ($p = 0.19$) in the latter model.

SA4 Plots of data coverage and nominal rye price series

Figure S1 shows the number of cross-sectional observations over time.

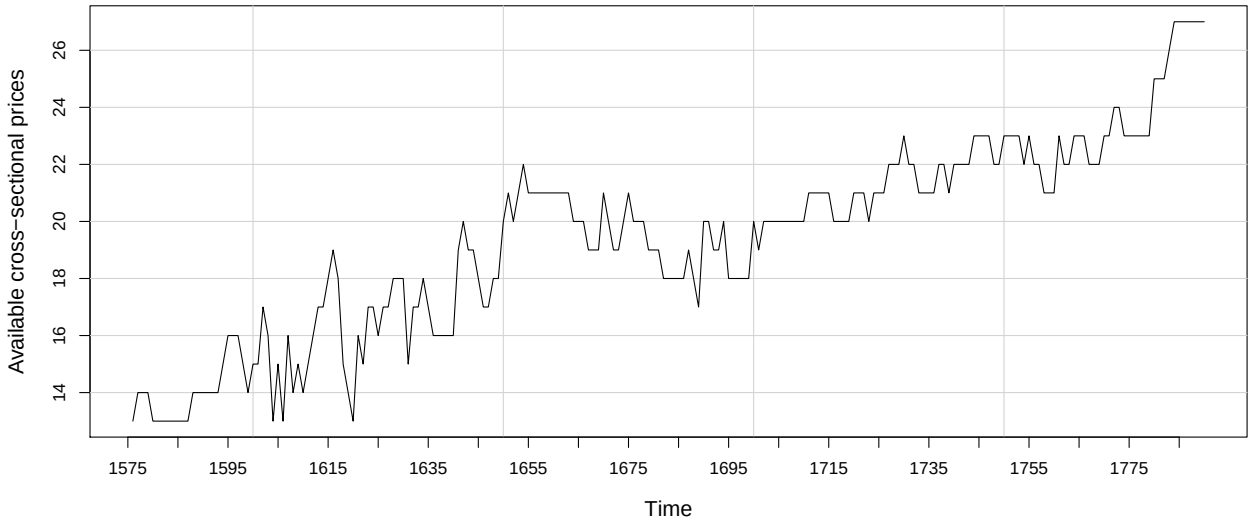

Figure S1: Data coverage over time, rye prices in individual cities, unbalanced sample 1575–1790

The following Figures S2–S4 plot the rye price series described in SA2 for the period 1601-1850.

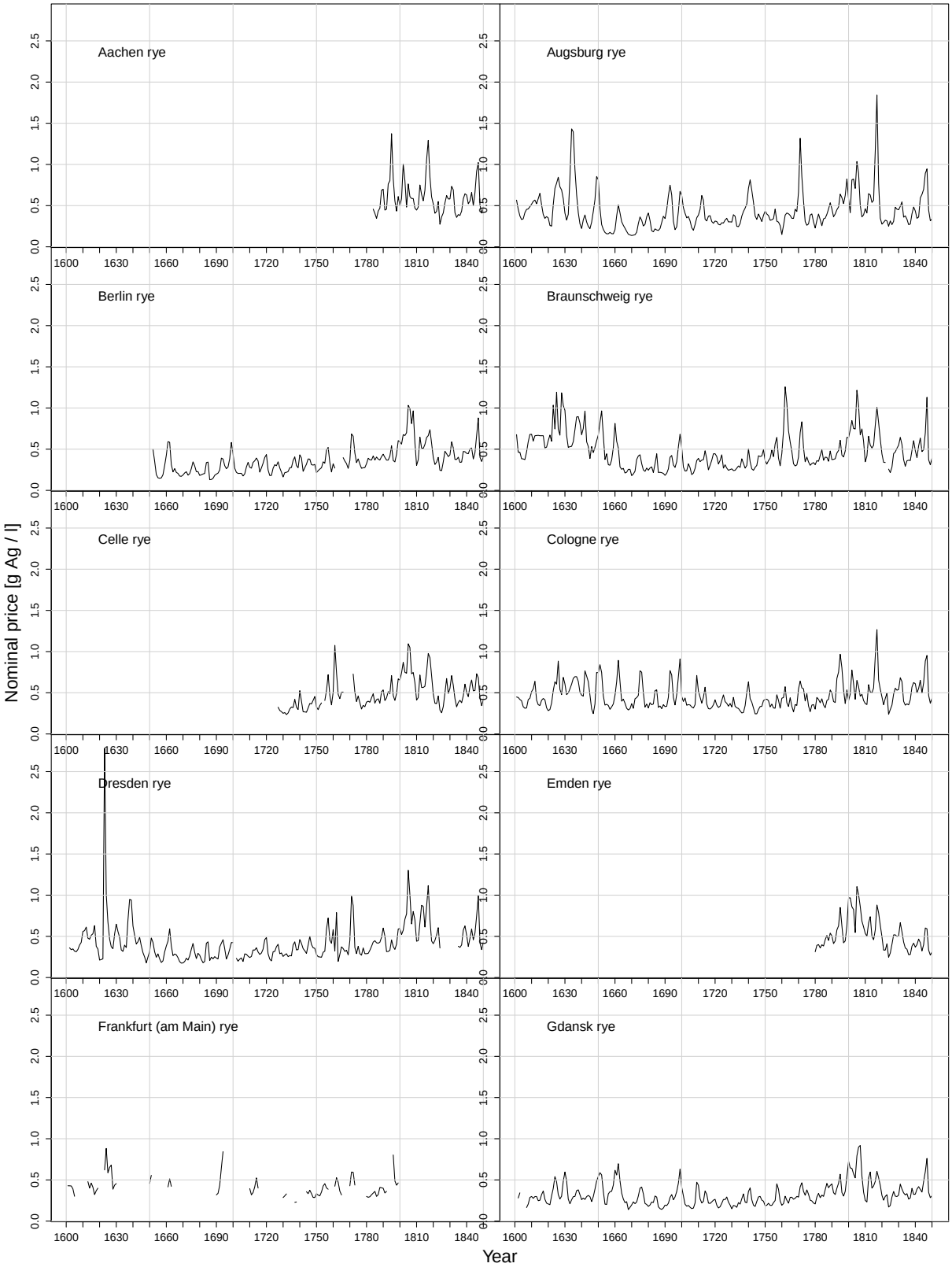

Figure S2: Nominal rye prices in grams of silver per litre, 1601–1850, part 1.

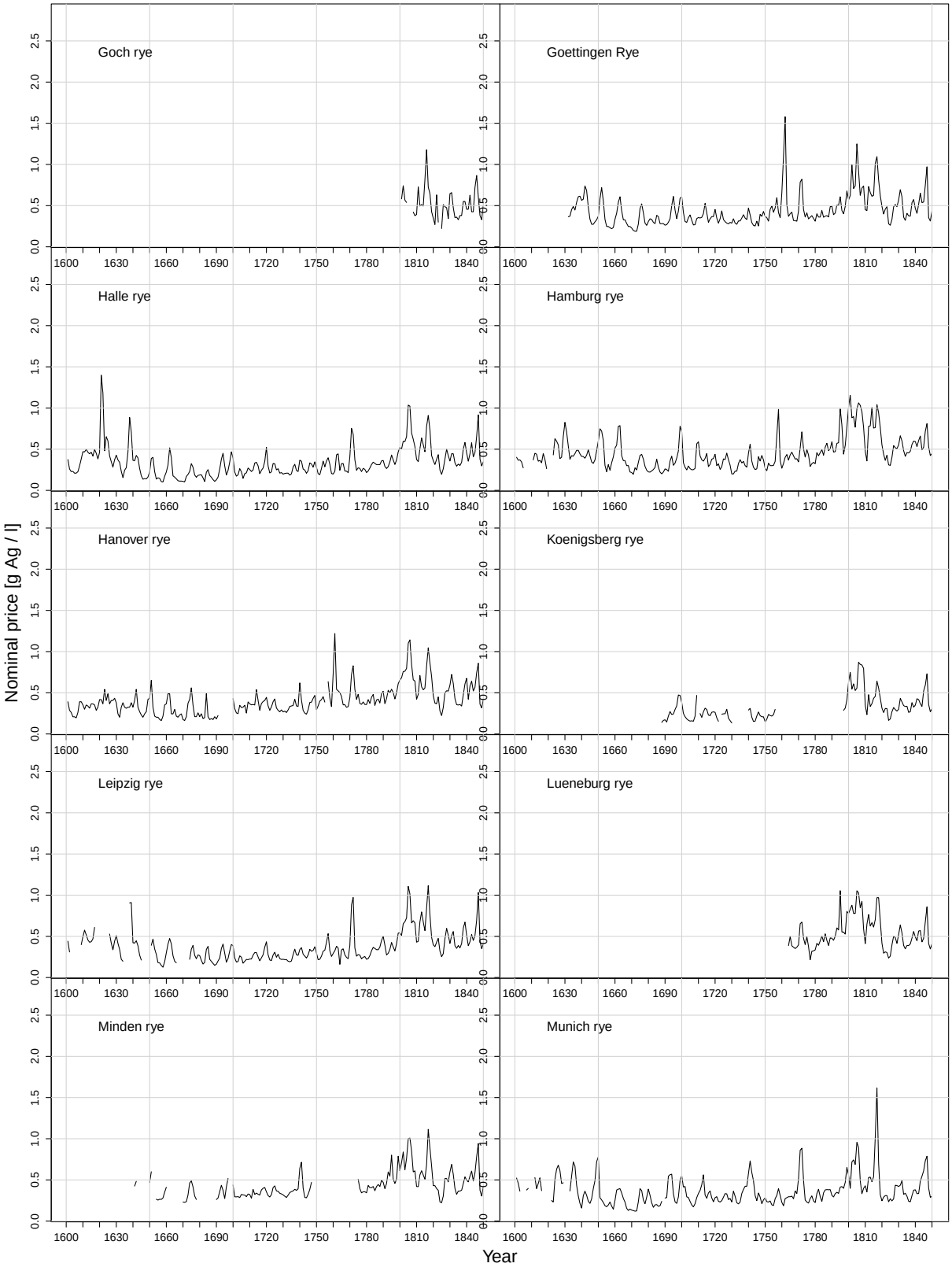

Figure S3: Nominal rye prices in grams of silver per litre, 1601–1850, part 2.

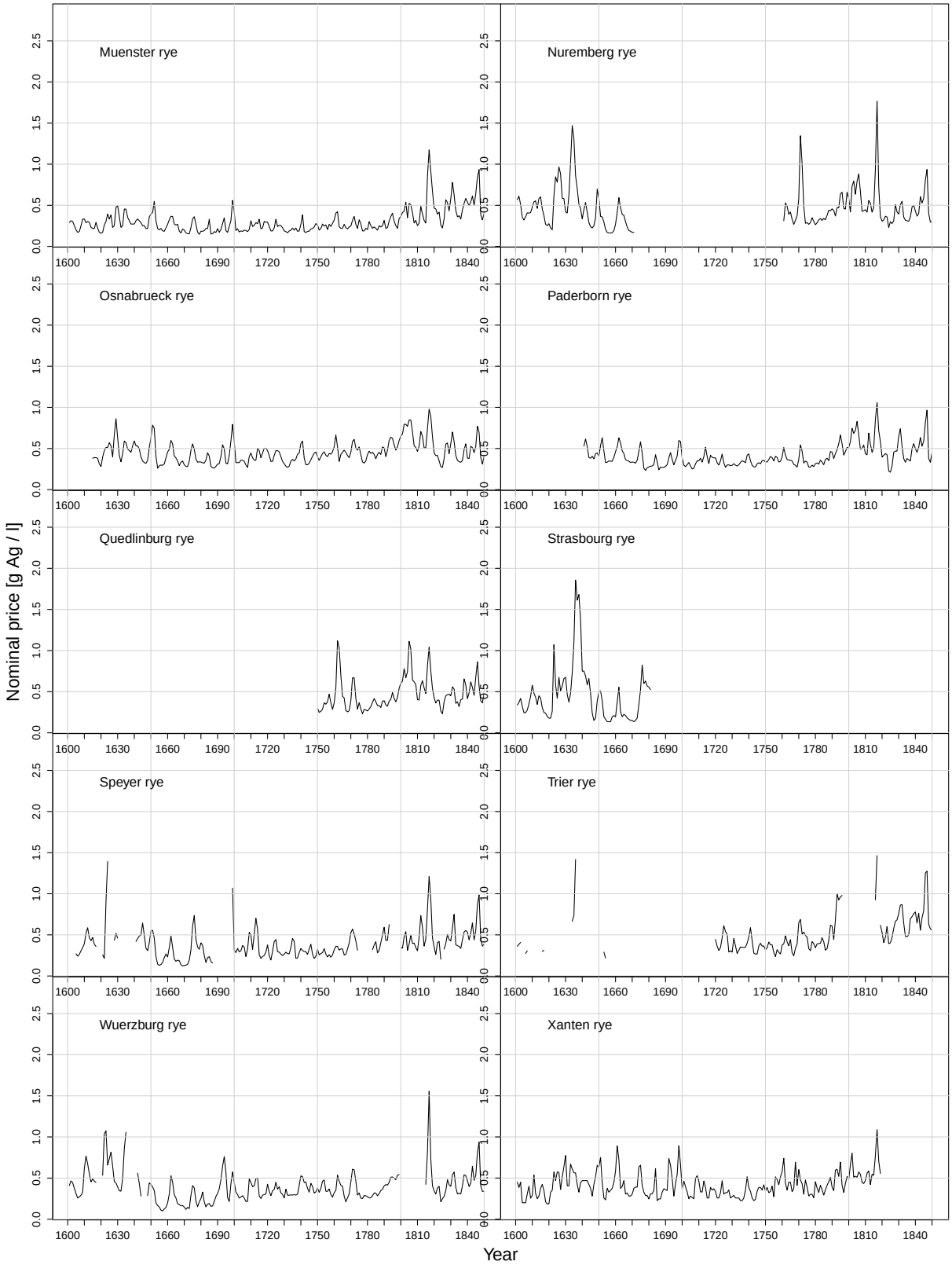

Figure S4: Nominal rye prices in grams of silver per litre, 1601–1850, part 3.

SA5 Time series properties of rye prices

Section SA5.1 reports results for Augmented Dickey Fuller (ADF) tests and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) tests for all rye price series of the stable sample. Similarly, Section SA5.2 contains corresponding results for the aggregate rye price.

SA5.1 Rye price series of stable sample

Table S19 shows the results of ADF tests for all rye prices of our stable sample 1651–1790. We test for the presence of a unit root in nominal and real prices including their log form (Dobado-Gonzàlez et al., 2012). The null hypothesis of a unit root can be rejected for all prices series with $p < 0.01$.

Table S19: Augmented Dickey Fuller tests, rye prices, stable sample 1651–1790

	Nominal		log(nominal)		real		log(real)	
	t-statistic	ρ	t-statistic	ρ	t-statistic	ρ	t-statistic	ρ
Augsburg	-6.69	0.60	-4.76	0.73	-6.10	0.66	-4.67	0.73
Berlin	-5.88	0.48	-7.45	0.51	-6.63	0.35	-6.76	0.34
Braunschweig	-6.57	0.59	-6.43	0.60	-6.01	0.62	-6.03	0.61
Cologne	-7.23	0.43	-6.84	0.48	-7.27	0.39	-6.92	0.43
Dresden	-7.00	0.36	-6.76	0.41	-6.80	0.36	-6.82	0.36
Gdansk	-5.48	0.67	-4.32	0.72	-5.57	0.65	-5.91	0.64
Goettingen	-7.13	0.46	-5.21	0.60	-5.18	0.57	-6.71	0.54
Halle	-6.10	0.42	-6.10	0.47	-8.23	0.40	-6.13	0.44
Hamburg	-6.73	0.49	-6.30	0.56	-5.94	0.39	-6.52	0.51
Munich	-6.92	0.61	-4.98	0.68	-6.16	0.63	-6.21	0.64
Muenster	-7.30	0.39	-7.08	0.42	-7.16	0.34	-7.09	0.36
Osnabrueck	-4.84	0.57	-6.11	0.53	-6.44	0.42	-6.47	0.42
Paderborn	-4.20	0.63	-4.07	0.66	-5.46	0.49	-4.23	0.58
Wuerzburg	-5.04	0.65	-6.20	0.65	-4.46	0.70	-5.47	0.69
Xanten	-6.24	0.47	-5.85	0.52	-6.81	0.38	-6.61	0.41

Note: Nominal prices in g Ag / l. The t-statistic is from the test regression including drift term, trend and up to 4 lags (Wooldridge, 2009, 633; lag length selection based on Akaike Information Criterion). Some series contain a few missing values which were linearly interpolated for this test. The critical value to reject the null hypothesis of a unit root at the 1% (5%; 10%) level is -3.99 (-3.43; -3.13). ρ is the autoregressive parameter deduced from the test regression. The null hypothesis of a unit root can be rejected for all prices series with $p < 0.01$.

If we split the sample, the pattern is very similar for the period 1651–1720: The null hypothesis of a unit root is rejected in all cases at least at the 10% level (in only two out of fifteen cases not at the 5% level; Table S20). Once we restrict the ADF test to the last 70 years of our stable sample (1721–1790), a unit root cannot be rejected for two out of fifteen cities in at least one specification (Table S21) at the 5% level. Only for one city, namely Paderborn, a unit root cannot be rejected at the 10% level.

Table S20: Augmented Dickey Fuller tests, rye prices, stable sample 1651–1720

	Nominal		log(nominal)		real		log(real)	
	t-statistic	ρ	t-statistic	ρ	t-statistic	ρ	t-statistic	ρ
Augsburg	*-3.78*	0.66	-4.19	0.69	*-3.77*	0.68	-4.27	0.63
Berlin	-5.13	0.55	-4.89	0.58	-5.15	0.51	-5.01	0.52
Braunschweig	-4.73	0.55	-4.67	0.57	-4.73	0.50	-4.72	0.52
Cologne	-5.43	0.39	-5.21	0.45	-5.25	0.38	-5.04	0.42
Dresden	-4.91	0.47	-4.93	0.48	-4.76	0.40	-4.91	0.40
Gdansk	*-4.04*	0.66	-4.26	0.67	-4.08	0.64	-4.36	0.64
Goettingen	-4.17	0.59	-4.26	0.61	-4.49	0.46	-4.61	0.49
Halle	*-3.47*	0.59	-5.08	0.54	-4.94	0.48	-4.70	0.38
Hamburg	-4.33	0.60	*-3.85*	0.65	-4.13	0.58	*-3.83*	0.61
Munich	-4.87	0.56	-4.16	0.53	-4.62	0.55	-4.80	0.57
Muenster	-5.02	0.41	-4.77	0.45	-5.06	0.31	-4.97	0.33
Osnabrueck	-4.17	0.52	-4.06	0.54	-4.61	0.41	-4.54	0.39
Paderborn	*-3.74*	0.58	**-3.17**	0.68	-4.09	0.47	*-3.92*	0.51
Wuerzburg	**-3.33**	0.66	-5.01	0.58	**-3.38**	0.64	-4.83	0.56
Xanten	-4.94	0.41	-4.58	0.46	-5.23	0.32	-5.02	0.35

Note: Nominal prices in g Ag / l. The t-statistic is from the test regression including drift term, trend and up to 4 lags (Wooldridge, 2009, 633; lag length selection based on Akaike Information Criterion). Some series contain a few missing values which were linearly interpolated for this test. The critical value to reject the null hypothesis of a unit root at the 1% (5%; 10%) level is -4.04 (-3.45; -3.15). *Italic* city names indicate that the null hypothesis was not rejected in at least one specification at the 1% level; **bold** (***bold italic***) for failure to reject at 5% (10%) level. ρ is the autoregressive parameter deduced from the test regression.

Table S21: Augmented Dickey Fuller tests, rye prices, stable sample 1721–1790

	Nominal		log(nominal)		real		log(real)	
	t-statistic	ρ	t-statistic	ρ	t-statistic	ρ	t-statistic	ρ
Augsburg	-4.30	0.59	-4.37	0.61	*-3.92*	0.63	*-4.03*	0.62
Berlin	-4.49	0.32	-5.25	0.40	-5.18	0.02	-5.52	-0.06
Braunschweig	-4.43	0.60	-4.13	0.62	*-3.76*	0.69	*-3.46*	0.69
Cologne	-5.52	0.14	-5.89	0.11	-5.47	0.07	-5.46	0.12
Dresden	-4.80	0.31	-4.35	0.38	-4.38	0.36	-4.25	0.37
Gdansk	*-3.60*	0.44	-5.21	0.43	-5.14	0.41	-5.18	0.42
Goettingen	-4.82	0.44	-4.60	0.51	-4.60	0.51	-4.10	0.57
Halle	-4.76	0.32	-4.74	0.27	-4.93	0.22	-6.50	0.27
Hamburg	-6.05	0.22	-6.09	0.29	-5.95	0.23	-5.61	0.05
Munich	**-3.28**	0.69	-4.37	0.62	-4.30	0.65	-4.26	0.62
Muenster	-5.10	0.34	-5.05	0.35	-4.79	0.39	-4.75	0.39
Osnabrueck	-4.50	0.44	-4.50	0.47	-5.59	0.44	-5.67	0.46
Paderborn	*-3.83*	0.47	*-3.82*	0.51	***-2.96***	0.60	***-2.82***	0.63
Wuerzburg	-4.52	0.56	-4.31	0.58	-4.80	0.21	-4.38	0.52
Xanten	*-3.75*	0.49	*-3.74*	0.52	-4.18	0.40	-4.23	0.42

Note: Nominal prices in g Ag / l. The t-statistic is from the test regression including drift term, trend and up to 4 lags (Wooldridge, 2009, 633; lag length selection based on Akaike Information Criterion). Some series contain a few missing values which were linearly interpolated for this test. The critical value to reject the null hypothesis of a unit root at the 1% (5%; 10%) level is -4.04 (-3.45; -3.15). *Italic* city names indicate that the null hypothesis was not rejected in at least one specification at the 1% level; **bold** (***bold italic***) for failure to reject at 5% (10%) level. ρ is the autoregressive parameter deduced from the test regression.

We also conducted KPSS tests of real prices for all individual cities (including their logged version and a sample split). The KPSS test reports more series to be neither trend nor level stationary and suggests a unit root in more cases than the ADF test. Nevertheless, the majority of price series is stationary for most of the time.

Table S22: KPSS tests, real rye prices, stable sample 1651–1790

	real rye price, p-values		log(real rye price), p-values	
Null-hypothesis	level st.	trend st.	level st.	trend st.
1651–1790				
Augsburg	<0.01	0.0292	<0.01	<0.01
Berlin	<0.01	>0.1	<0.01	>0.1
Braunschweig	>0.1	0.017	0.0707	0.0201
Cologne	<0.01	>0.1	<0.01	>0.1
Dresden	<0.01	>0.1	<0.01	>0.1
Gdansk	0.0601	<0.01	>0.1	<0.01
Goettingen	>0.1	>0.1	>0.1	>0.1
Halle	<0.01	>0.1	<0.01	>0.1
Hamburg	>0.1	0.0613	>0.1	0.0479
Munich	0.065	0.0575	0.0247	0.0225
Muenster	>0.1	>0.1	>0.1	>0.1
Osnabrueck	>0.1	>0.1	>0.1	>0.1
Paderborn	<0.01	>0.1	<0.01	>0.1
Wuerzburg	<0.01	<0.01	<0.01	<0.01
Xanten	0.0953	>0.1	>0.1	0.094

Note: KPSS test: Kwiatkowski-Phillips-Schmidt-Shin test; st.: stationarity. **Bold** city names indicate that the series cannot be regarded as either trend or level stationary. Short truncation lag parameter used.

Table S23: KPSS tests, real rye prices, stable sample 1651–1720

	real rye price, p-values		log(real rye price), p-values	
Null-hypothesis	level st.	trend st.	level st.	trend st.
Augsburg	<0.01	>0.1	<0.01	>0.1
Berlin	>0.1	>0.1	>0.1	>0.1
Braunschweig	0.0486	<0.01	0.091	<0.01
Cologne	>0.1	>0.1	>0.1	>0.1
Dresden	>0.1	>0.1	>0.1	>0.1
Gdansk	0.0358	0.0553	0.0468	0.0766
Goettingen	> 0.1	>0.1	>0.1	>0.1
Halle	0.0177	>0.1	<0.01	>0.1
Hamburg	>0.1	0.0163	>0.1	0.012
Munich	<0.01	>0.1	<0.01	>0.1
Muenster	>0.1	>0.1	>0.1	0.0791
Osnabrueck	>0.1	>0.1	>0.1	>0.1
Paderborn	0.0448	0.0816	0.043	0.0472
Wuerzburg	<0.01	>0.1	<0.01	>0.1
Xanten	>0.1	>0.1	>0.1	>0.1

Note: KPSS test: Kwiatkowski-Phillips-Schmidt-Shin test; st.: stationarity. **Bold** city names indicate that the series cannot be regarded as either trend or level stationary. Short truncation lag parameter used.

Table S24: KPSS tests, real rye prices, stable sample 1721–1790

	real rye price, p-values		log(real rye price), p-values	
Null-hypothesis	level st.	trend st.	level st.	trend st.
Augsburg	>0.1	>0.1	>0.1	>0.1
Berlin	<0.01	>0.1	<0.01	>0.1
Braunschweig	0.0576	<0.01	0.033	<0.01
Cologne	>0.1	>0.1	>0.1	>0.1
Dresden	>0.1	>0.1	>0.1	>0.1
Gdansk	<0.01	>0.1	<0.01	>0.1
Goettingen	>0.1	0.0162	0.0861	<0.01
Halle	>0.1	>0.1	>0.1	>0.1
Hamburg	0.0252	>0.1	<0.01	>0.1
Munich	>0.1	>0.1	>0.1	>0.1
Muenster	>0.1	0.066	>0.1	0.0579
Osnabrueck	>0.1	>0.1	>0.1	>0.1
Paderborn	>0.1	0.0891	>0.1	0.0886
Wuerzburg	<0.01	0.0664	<0.01	0.0648
Xanten	0.055	<0.01	0.0449	<0.01

Note: KPSS test: Kwiatkowski-Phillips-Schmidt-Shin test; st.: stationarity. **Bold** city names indicate that the series cannot be regarded as either trend or level stationary. Short truncation lag parameter used.

SA5.2 Aggregate rye price

Table S25 shows the results of ADF-tests for the aggregate rye price (for the construction of this variable, see SA6) for three periods 1651–1790, 1651–1720, and 1721–1790. The null hypothesis of a unit root can be rejected for all specifications covering the years 1651–1790 with $p < 0.01$. In all specifications for the years 1651–1720, a unit root is rejected at the 5% level. But for 1721–1790 a unit root cannot be rejected for the real price of the stable sample.

Table S25: Augmented Dickey Fuller tests, aggregate rye price

	Nominal		log(nominal)		real		log(real)	
	t-statistic	ρ	t-statistic	ρ	t-statistic	ρ	t-statistic	ρ
1651–1790								
Stable sample	-5.42	0.59	-5.39	0.62	-5.53	0.57	-5.63	0.57
Unbalanced	-5.72	0.57	-5.72	0.60	-5.83	0.55	-5.94	0.56
Sub-period 1651–1720								
Stable sample	*-3.75*	0.64	*-3.80*	0.66	-5.39	0.58	-5.16	0.60
Unbalanced	*-3.89*	0.60	*-4.02*	0.63	-4.25	0.55	-4.38	0.55
Sub-period 1721–1790								
Stable sample	-5.08	0.49	-4.69	0.54	***-3.11***	0.64	-4.12	0.58
Unbalanced	-4.17	0.50	*-3.98*	0.52	*-3.66*	0.57	*-3.68*	0.56

Note: Nominal prices in g Ag / l. The t-statistic is from the test regression including drift term, trend and up to 4 lags (Wooldridge, 2009, 633; lag length selection based on Akaike Information Criterion). The critical value to reject the null hypothesis of a unit root at the 1% (5%; 10%) level is -3.99 (-3.43; -3.13) for the period 1651–1790; -4.04 (-3.45; -3.15) for 1651–1720 and 1721–1790. *Italic* row names indicate that the null hypothesis was not rejected in at least one specification at the 1% level; **bold** (***bold italic***) for failure to reject at 5% (10%) level. ρ is the autoregressive parameter deduced from the test regression.

These results for ADF tests of the aggregate rye price are corroborated by the KPSS test results reported in Table S26 (all for the real rye price). The KPSS test clearly rejects level or trend stationarity for the period 1720–90 but not for the years 1651–1720. If the full sample period is included (1651–1790), level stationarity is rejected for the unbalanced sample, whereas trend stationarity is not rejected.

Table S26: KPSS tests, aggregate real rye price

	real rye price, p-values		log(real rye price), p-values	
Null-hypothesis	level st.	trend st.	level st.	trend st.
1651–1790				
Stable sample	> 0.1	> 0.1	> 0.1	> 0.1
Unbalanced	0.0515	> 0.1	0.0229	> 0.1
Sub-period 1651–1720				
Stable sample	> 0.1	> 0.1	> 0.1	> 0.1
Unbalanced	> 0.1	> 0.1	> 0.1	> 0.1
Sub-period 1721–1790				
Stable sample	0.0774	0.0227	0.0699	0.0229
Unbalanced	0.0426	0.0476	0.0364	0.0509

Note: KPSS test: Kwiatkowski-Phillips-Schmidt-Shin test; st.: stationarity. **Bold** row names indicate that the series cannot be regarded as either trend or level stationary. Short truncation lag parameter used.

SA6 The aggregate real price of rye in Germany, 1600–1850

In SA6.1, we first provide more details concerning the construction and deflation of the aggregate real price for rye in pre-industrial Germany. Section SA6.2 discusses an alternative approach to calculate the aggregate price (panel regression instead of arithmetic mean). In Section SA6.3, we review the role of grain prices in the consumer price index (CPI), which is used to deflate the aggregate rye price. To provide a broader context, we additionally include price data from 1601 and up to 1850 in our unbalanced sample.

SA6.1 Deflation of the aggregate nominal rye price

In view of the analysis of price volatility, we constructed the aggregate price series in year t as the cross-sectional arithmetic mean of the annual rye prices in individual cities in year t. To isolate grain price shocks from inflationary shocks that affected the price level as a whole, we consider real rather than nominal grain prices. Deflation of the aggregate series is based on the CPI constructed by Pfister (2017).[29] This index is the silver price of a basket of fixed quantities of eleven goods presumably consumed annually by an adult town dweller. Thus, we calculate the real price as the ratio of the silver price of 1000 litres of a particular grain

[29]Pfister (2017) shows that the loss of information from using a national CPI as deflator instead of local deflators makes little difference when deflating wages.

type to the annual silver price of a consumer basket. In addition to deflating the grain price, this ratio preserves the information of how much these 1000 litres of grain cost relative to the consumer basket.

SA6.2 Alternative method of aggregation

In what follows, we discuss an alternative approach to derive an aggregate real rye price based on a panel regression. For the unbalanced sample, this alternative approach leads to a very similar estimate of the aggregate real price. We thus consider the approach in the main paper robust to this alternative.

Since the price series for individual towns cover different time periods, the alternative version is based on an unbalanced panel regression with fixed effects for cities and years. The idea is that the city-specific dummy variables absorb unobserved characteristics of individual markets. Thus, the advantage of the panel regression is that the estimate of the aggregate price for individual years is not biased due to unobserved city-specific effects when single cities enter or leave the sample. A potential shortcoming is that the panel regression extracts a particular type of variance: common shocks to all cities. Idiosyncratic shocks instead are part of the error term and do not become part of the aggregate price. The estimated panel model is:

$$p_{it} = \alpha_0 + \sum_{i=1}^{N-1} \alpha_i c_i + \sum_{t=1}^{T-1} \beta_t y_t + u_{it}. \tag{S9}$$

Herein, p_{it} denotes the mean price in city $i = 1, ..., N$ $(N = 30)$[30] in year $t = 1, ..., T$ $(T = 250$; period: 1601–1850); c_i are city-specific dummy variables, y_t are year dummy variables and u_{it} is the error term. The aggregate price in year t is then defined as: $\bar{p}_t = \alpha_0 + 1/(N-1) \sum_{i=1}^{N-1} \alpha_i + \beta_t$. Thus, the aggregate price for an individual year t is calculated as the sum of the constant α_0, the mean of the city effects α_i (i.e. the average deviation from the constant), and the corresponding parameter β_t of the time dummy variable. Like the aggregate price based on the cross-sectional mean, we deflated this alternative nominal price series based on the panel regression with the CPI from Pfister (2017).

We estimated eq. S9 by applying the least squares dummy variable estimator on prices both in levels and in logs. It turned out that the fit of the regression in levels ($R^2 = 0.65$) is slightly higher than of the variant in logs ($R^2 = 0.64$; see Wooldridge, 2009, 213 on calculation of comparable R^2). Hence, we only consider the results for prices in levels. Figure

[30] Inclusion of data for the 19th century increases the unbalanced sample from 29 to 30 cities as the city of Goch enters the sample.

S5 shows the trajectory of the aggregate real price; the latter is very similar irrespective of the methodology used.

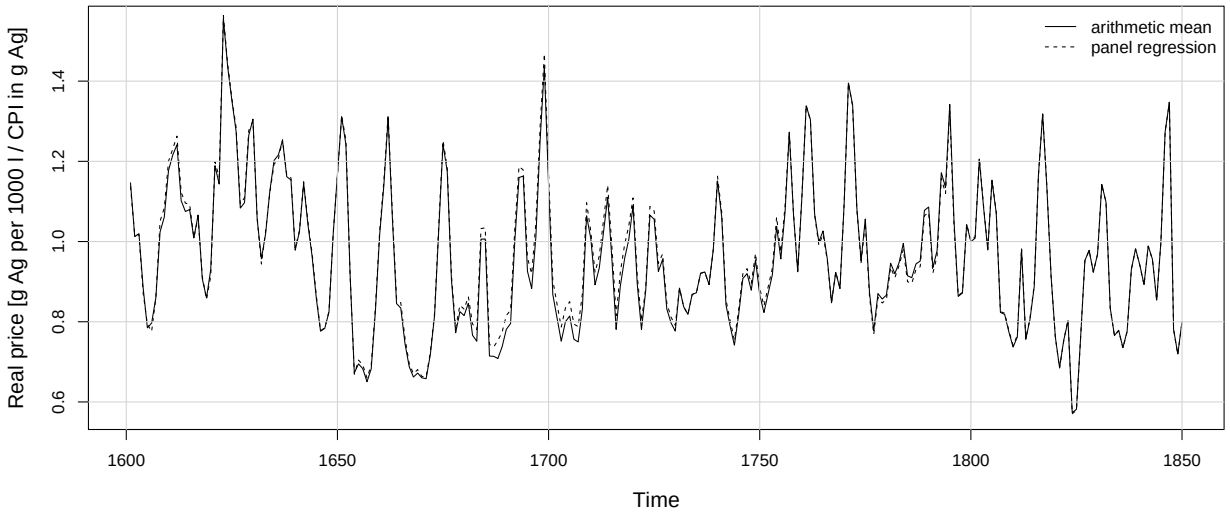

Figure S5: Two approaches to calculate the aggregate real rye price Germany. Unbalanced sample, 30 cities 1601–1850. Source: own representation.

SA6.3 The role of cereal prices in the CPI

The CPI contains much variation from cereal prices. One might argue that this eliminates real trends and fluctuations in the aggregate real price. Earlier scholarship has argued that there was a secular increase of cereal prices in the second half of the 18th century (Abel, 1978, 196–8). For Germany, Abel's observation is based on the mean of 10-year-average *nominal* rye prices from 13 cities in grams of silver per 100 kg (Abel, 1978, 196–8). According to Abel, the rise of the grain price induced land expansion and increased the intensity of land use (Abel, 1978, 196–8, 200–2). Abel discusses two reasons for increasing cereal prices. First, a real factor, that is, increasing demand as a result of increasing population and second, a monetary factor, namely, an expansion of the monetary base due to an increase in the world production of silver. Figure S6 shows the nominal rye price, 1651–1850 for the unbalanced sample. There is a statistically significant upward trend in 1731–92, indicating that Abel's observation is preserved in our nominal data (1792 is chosen as the last year, because the Revolutionary Wars began in that year and silver contents are not reliable during the subsequent years due to real shocks from war and monetary shocks resulting from war finance). The secular price increase, however, is not visible in our deflated rye price series: the size of the trend estimate is reduced very much and statistical significance is lost.

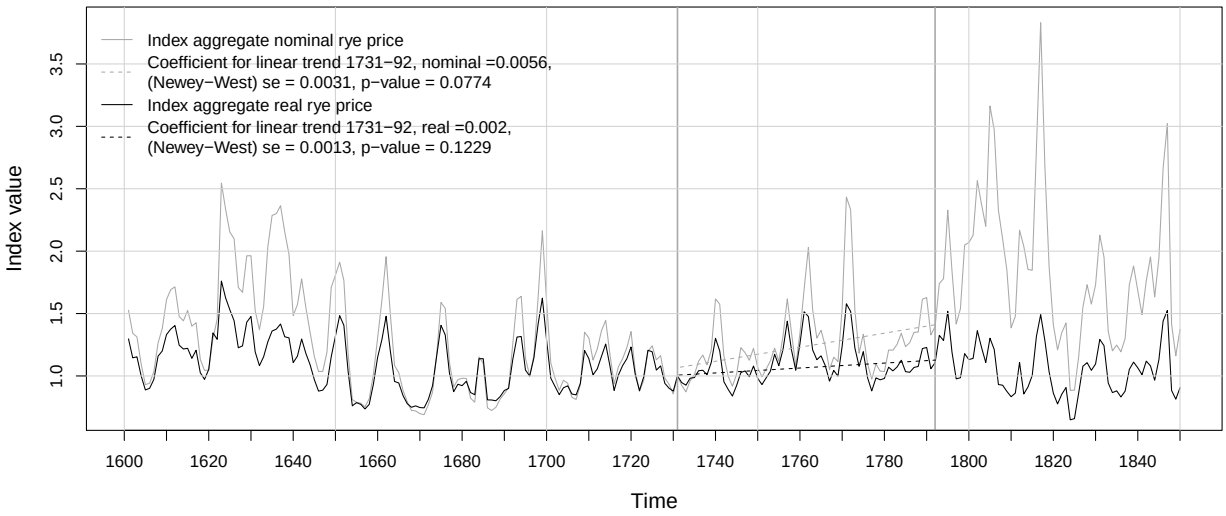

Figure S6: Index of nominal and real rye price, 1731 = 1; unbalanced sample, 30 cities 1601–1850. Vertical lines highlight sub-period of upward trend in nominal price discussed by Abel (1978, 196–8); 1731 was chosen by Abel; in 1792 the Napoleonic Wars began. Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Source: own representation.

One concern with this finding might be that the consumer basket that Pfister (2017) uses to construct the CPI is driven by cereal prices to a large extent. Thus, (real) trends and even much of the annual fluctuations could be eliminated when deflating nominal rye prices, because the CPI is positively correlated with the nominal rye price. The slope parameter from a regression of the CPI on the nominal aggregate rye price (both variables log-differenced) is 0.3 ($R^2 = 0.75$; stable sample 1651-1790; slope parameter with data in logs but no differences: 0.35; $R^2 = 0.8$). That is, a 1% increase of the nominal rye price is associated with a 0.3% increase of the CPI.

In fact, cereal prices affect a large fraction of expenditure in the CPI (53%: bread, beer, eggs via grain as feed; Pfister, 2017, S1, p. 1)[31] but do not enter the CPI directly. This might explain the relatively small slope coefficient from the regression of the CPI on the nominal aggregate price in comparison to the high share of prices related to cereals. For example, rye enters the bread price but the latter is also determined by wage costs (similarly for beer where barley is the underlying cereal). This is relevant because the ratio of the prices for rye over bread increases by ca. 0.13% per year (7 cities 1650–1797, contains two shorter series for Göttingen and Munich; data from Pfister, 2017). In other words, rye becomes relatively more expensive over time than bread. Thus, using bread instead of rye in the CPI in fact

[31]Meat and butter production might potentially also be affected (grain as feed) but Pfister (2017, S1) used beef prices for meat. Cows are ruminants and thus, their advantage is to convert feed which is not suitable for human consumption (forage such as grass and hay) into food for human beings.

works against the elimination of real trends from the nominal rye series. In addition, bread prices fluctuate less than cereal prices (Pfister, 2017, S2, pp. 4–5). Furthermore, a budget share of 17% in the basket underlying the CPI relates to non-food commodities (e.g., candles and firewood). Among the non-food commodities we observe an important upward trend for firewood prices, a land intensive good which has an average share of 5.5% in the CPI (energy crisis, Pfister, 2017, S2, p. 8). The increase of the real price for firewood my constitute the principal reason why the real rye price remained stable.

We conclude that the CPI should be used to account for economy-wide monetary inflation and that there is only a modest variability dampening effect by cereal-sensitive-goods in the denominator of the real rye price. Much of the increase in the aggregate rye price which Abel (1978) discussed was rather a monetary phenomenon, at least in the German case.

SA7 Definition of North-Western and continental Germany

Using continentality as an indicator of local climate, we divide the markets included in this study into two sub-regions, namely, a north-western sub-sample and the remainder located in the interior of the mainland, which we relate to as 'continental Germany'.

According to the Köppen-Geiger classification by Kottek et al. (2006), Germany has a temperate climate, fully humid with warm summer (Cfb; data for 1951–2000).[32] A version with a higher spatial resolution by Peel et al. (2007) (which also applies a slightly different threshold between temperate (C) and cold (D) climates), confirms this classification for Northern, South-Western and Western Germany, but classifies large parts of South-Eastern and Eastern Germany as cold, fully humid, with warm summer (Dfb; precipitation: 1909–1991, temperature: 1923–1991).

Consistent with the latter classification, the continentality of the climate (measured by the within-year temperature difference between the hottest and the coldest month) within Germany increases in south-eastern direction (based on 1961–1990 climate normal; Müller-Westermeier et al., 2001, map 7). Additionally, other research shows that while the spatial correlation of temperature within Germany is very high (50% common variance at a distance of 1050 km; 1956–1995), the one of precipitation is much lower (120 km; 1896–1995) (Rapp, 2000, 30–35). Thus, cities are allocated to North-Western Germany if continentality is measured as ≤ 17.5°C (Müller-Westermeier et al., 2001, map 7). Gdansk is not part of the latter classification. The WMO climate normal 1961–90 is only available for the official

[32] A Köppen-Geiger map showing regional variations of climates for the studied period would be optimal, because climatic regions might change over time. To our knowledge, such a map is not available.

station in Hel (ca. 35 km from Gdansk) and yields a continentality of 17.6°C (WMO, 2017b).[33] Thus, Gdansk is allocated to continental Germany.

SA8 Relationship between the aggregate real rye price and weather variables

We first analyze the time series properties of several weather variables, which were reconstructed by previous research (Luterbacher et al., 2004; Xoplaki et al., 2005; Pauling et al., 2006). Second, we estimate short-run relationships between the aggregate real rye price and these weather variables.

SA8.1 Time series properties of weather variables

Table S27 shows unit root and stationarity tests for the weather variables that we use to estimate a short-run relationships with the real rye price. Note that both temperature and precipitation variables are *European* averages and thus only a proxy variable for conditions in Germany. The ADF-test rejects the null-hypothesis of a unit root for all but one series at the 1% level; for winter precipiation in the sub-period 1721–90 at the 5% level. For winter and spring temperature, the KPSS-test rejects level and trend stationarity for the entire period 1651–1790. In all other cases the series are level or trend stationary based on the KPSS-test. The spring temperature series and winter precipitation for 1720–90 show the highest first order serial correlation (ρ, Table S27).

For spring temperature 1651–1790 from Xoplaki et al. (2005), we also performed an additional test based on the specification used by Kelly and Ó Gráda (2014, 1384). Also in this specification the series shows statistically significant first order autocorrelation of 0.27, and a significant positive trend (heteroscedasticity consistent standard error; $R^2 = 0.13$).

[33] Similarly the continentality for Gdansk based on a later climate normal (1981–2010) is 19.1°C (WMO, 2017a).

Table S27: Time series properties weather variables 1651–1790

	ADF-test		*p*-values KPSS-test	
	t-statistic	ρ	level st.	trend st.
1651–1790				
Winter temp.	-5.47	0.14	0.0351	<0.01
Spring temp.	-6.98	0.26	<0.01	0.0149
Autumn temp.	-7.9	0.02	>0.1	>0.1
Winter prec.	-6.45	0.25	>0.1	>0.1
Sub-period 1651–1720				
Winter temp.	-7.04	-0.4	>0.1	>0.1
Spring temp.	-5.75	0.03	0.0217	>0.1
Autumn temp.	-5.15	0.14	>0.1	>0.1
Winter prec.	-4.86	0.02	>0.1	>0.1
Sub-period 1721–1790				
Winter temp.	-6.7	-0.32	<0.01	>0.1
Spring temp.	-5.07	0.21	>0.1	>0.1
Autumn temp.	-5.46	-0.02	>0.1	>0.1
Winter prec.	-3.77	0.38	>0.1	<0.01
Δ Spring temp.	-7.36	-1.13	>0.1	>0.1
Δ Autumn temp.	-7.56	-3.23	>0.1	>0.1
Δ Winter temp.	-7.29	-2.53	>0.1	>0.1
Δ Winter prec.	-5.79	-2	>0.1	>0.1

Note: The t-statistic is from the Augmented Dickey Fuller (ADF) test regression including drift term, trend and up to 4 lags (Wooldridge, 2009, 633; lag length chosen based on Akaike Information Criterion). The critical value to reject the null hypothesis of a unit root at the 1% (5%; 10%) level is -3.99 (-3.43; -3.13) for the period 1651–1790; -4.04 (-3.45; -3.15) for 1651–1720 and 1721–1790. The null hypothesis is rejected in all specifications at the 1% or 5% level. ρ is the autoregressive parameter deduced from the test regression. Δ indicates that the series is first-differenced. Temp.: temperature; prec.: precipitation. KPSS test: Kwiatkowski-Phillips-Schmidt-Shin test; st.: stationarity. **Bold** row names indicate that the series cannot be regarded as either trend or level stationary. Short truncation lag parameter used. Data sources: temp.: Luterbacher et al. (2004), Xoplaki et al. (2005); prec.: Pauling et al. (2006); data retrieved from NOAA (2017; 2018).

Figure S7 shows the two temperature variables which exhibit a changing pattern during the period 1651–1720 corroborating the results of the KPSS tests from Table S27.

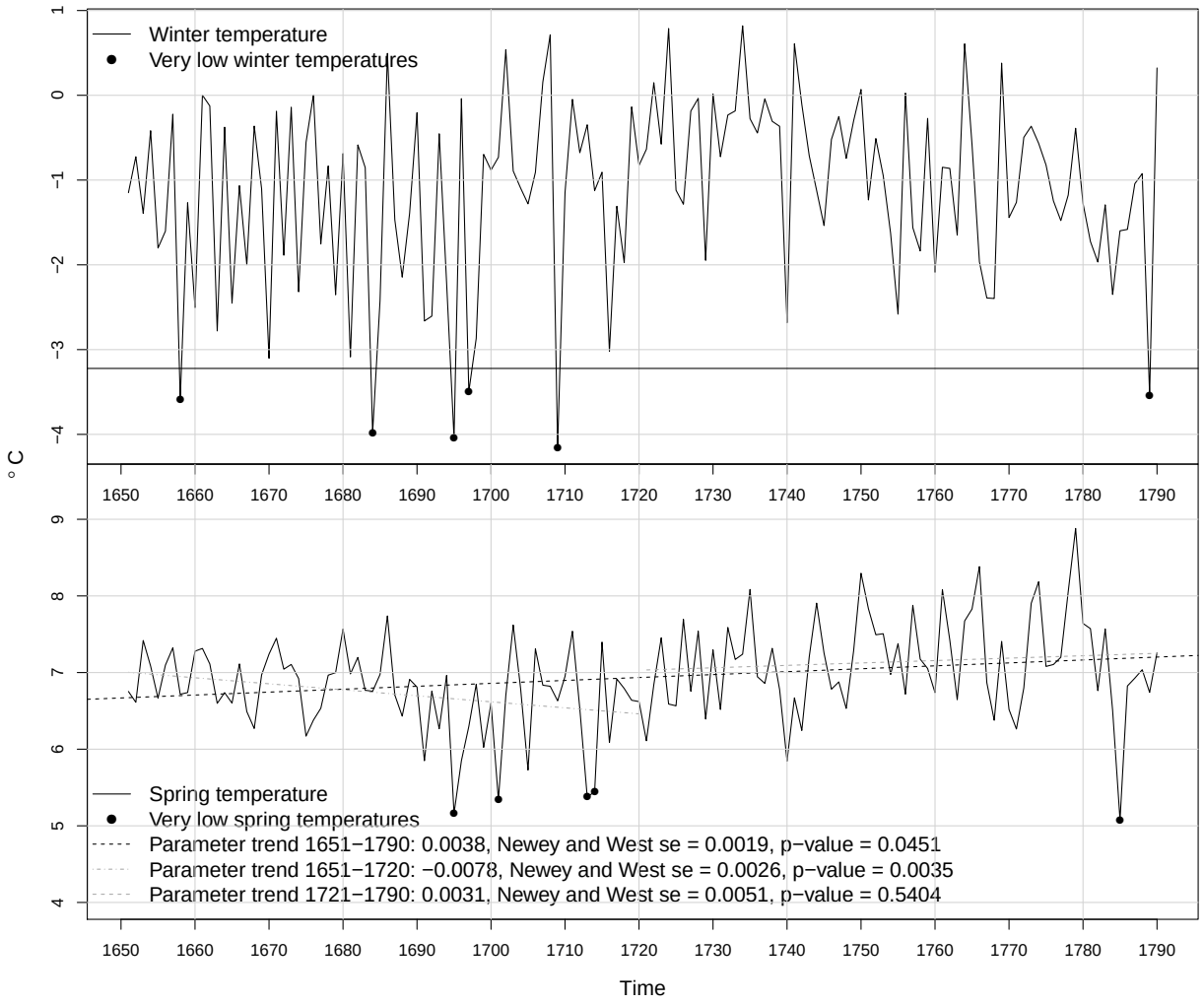

Figure S7: Temperature reconstructions for Europe 1651–1790. Very low: < mean minus two times the standard deviation. Own representation. Horizontal line in upper panel is drawn at the average winter temperature of 1963 (-3.22°C), when major rivers (Elbe, Rhine, Main) were frozen. Data sources: Luterbacher et al. (2004); Xoplaki et al. (2005); data retrieved from NOAA (2017).

SA8.2 Empirical relationship of real rye price and weather variables based on linear regression models

The time series properties of the real rye price and the weather variables are complex and partly changing. The rye price is stationary until 1720 but not anymore after 1720 (Tables S25 and S26). Spring temperature shows a positive trend, winter temperature a changing variance patterns as discussed above and visibile in Figure S7. Thus, we specify two regression models: one for the sub-period 1651–1720; the other for the sub-period 1721–1790. In

this way, we also achieved residuals free of significant serial correlation while using lagged dependent variables (Wooldridge, 2009, 412).

For the first sub-period, all variables are at least trend stationary and hence we specify a regression model where the logarithm of the aggregate real rye price p_t is regressed on a linear trend t and the following i weather variables x_{it}: winter temperature x_{1t}, winter precipitation x_{2t}, an interaction term of the latter two variables, and spring temperature x_{3t}. Additionally, we include two lags of the dependent variable (to account for serial correlation): $\log(p_t) = \alpha + \rho_1 \log(p_{t-1}) + \rho_2 \log(p_{t-2}) + \sum_{i=1}^{3} \beta_i x_{it} + \beta_{12}(x_{1t}x_{2t}) + u_t$. Due to the inclusion of an interaction term, winter temperature and precipitation enter in mean-centered form.

The estimate for β_1 and β_2 are statistically significant (Table S28). The results show that higher winter temperatures are associated with lower prices (β_1) and higher winter precipitation with higher prices (β_2). In particular, a one unit (°C) increase of winter temperature is related to a 2% reduction of the aggregate rye price. A one unit (mm) increase in precipitation is related to a 2% increase of the aggregate price. The interaction term is not significant at conventional levels but was necessary to capture non-linearities; otherwise the regression specification error test (RESET) was not passed at 10% level. The higher the winter temperatures, the less positive (negative) is the relationship of winter precipitation and price (yield). Higher spring temperature are associated with lower prices; but the result is not significant at conventional levels ($p = 0.12$). However, the Akaike information criterion (AIC) is lower for model 1 compared to model 3 and thus, we favored the latter.

Table S28: Relation of aggregate real rye price and weather variables 1651–1720

	(1)	(2)	(3)	(4)
	Baseline	w/o interaction	w/o spring temp.	w/o weather variables
Intercept	−1.3792	−1.2478	−2.1973*	−1.9178
	(1.2393)	(1.2342)	(1.1409)	(1.1552)
Lag 1 price	1.1209***	1.1141***	1.1195***	1.0923***
	(0.0991)	(0.0990)	(0.1004)	(0.1032)
Lag 2 price	−0.5766***	−0.5559***	−0.5411***	−0.5308***
	(0.0989)	(0.0971)	(0.0976)	(0.0992)
Winter temp.	−0.0212*	−0.0202*	−0.0252**	
	(0.0112)	(0.0112)	(0.0110)	
Winter prec.	0.0018*	0.0017*	0.0016*	
	(0.0009)	(0.0009)	(0.0009)	
Winter temp. x winter prec.	−0.0009		−0.0006	
	(0.0009)		(0.0009)	
Spring temp.	−0.0421	−0.0365		
	(0.0266)	(0.0260)		
Trend	0.0010	0.0009	0.0013*	0.0011
	(0.0007)	(0.0007)	(0.0007)	(0.0007)
R^2	0.7302	0.7251	0.7188	0.6825
Adj. R^2	0.6987	0.6981	0.6912	0.6676
Num. obs.	68	68	68	68

Note: ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$. Dependent variable: logged aggregate real rye price. Standard errors in (). For all models, the Breusch-Godfrey test cannot reject the null hypothesis of no serial correlation; the Breusch-Pagan test does not detect heteroscedasticity. temp.: temperature; prec.: precipitation; w/o: without.

For sub-period 1721–1790, we specify the model in log-differences for all variables to ensure stationarity. We also include the crude death rate as an indicator of changing demand conditions:[34]
$\Delta \log(p_t) = \alpha + \rho_1 \Delta \log(p_{t-1}) + \rho_2 \Delta \log(p_{t-2}) + \beta \Delta \log(x_t) + \sum_{i=1}^{4} \gamma_i \log(d_{t-i}) + \Delta u_t$. The notation is similar as for the previous model for the years 1651–1720 but the only included weather variable x_t is winter precipitation. Spring and autumn temperature were significant in some specifications but not robust to the inclusion of lagged mortality. Additionally, we tested a dummy variable for the Seven Years' War (1756–63).

[34]Results for weather variables not in logs but only in first differences are similar. We opted for weather variables in logs due to an easier interpretation of the model parameters as elasticities.

Table S29: Relation of aggregate real rye price and weather variables 1721–1790

	(1) Baseline	(2) Seven Years' War	(3) Drop weather variable	(4) No CDR, more weather variables	(5) add CDR	(6) from 1690
Intercept	0.0017	0.0020	0.0013	0.0028	0.0020	−0.0030
	(0.0125)	(0.0125)	(0.0132)	(0.0115)	(0.0127)	(0.0098)
Lag 1 price	0.1964	0.1636	0.1419	0.0966	0.1978	0.2809***
	(0.1372)	(0.1395)	(0.1424)	(0.1132)	(0.1392)	(0.0967)
Lag 2 price	−0.2152	−0.2316	−0.1572	−0.4402***	−0.2547	−0.4200***
	(0.1642)	(0.1641)	(0.1709)	(0.1134)	(0.1741)	(0.1020)
Winter prec.	0.2843**	0.2852**		0.2378**	0.2762**	0.2211**
	(0.1128)	(0.1124)		(0.1101)	(0.1151)	(0.0901)
Lag 1 CDR	−0.2062	−0.1830	−0.1894		−0.1777	−0.1485*
	(0.1510)	(0.1517)	(0.1586)		(0.1573)	(0.0823)
Lag 2 CDR	−0.2745**	−0.2663**	−0.2646*		−0.2434*	−0.1588*
	(0.1319)	(0.1316)	(0.1386)		(0.1398)	(0.0833)
Lag 3 CDR	0.1218	0.1192	0.1098		0.1121	−0.0141
	(0.1151)	(0.1147)	(0.1209)		(0.1217)	(0.0821)
Lag 4 CDR	−0.1981*	−0.1857	−0.1774		−0.1894	−0.2080***
	(0.1134)	(0.1134)	(0.1189)		(0.1156)	(0.0789)
7 Years' War		0.0800		0.0799		
		(0.0678)		(0.0685)		
Spring temp.				−0.1784*	−0.0887	
				(0.1057)	(0.1204)	
Lag 1 autumn temp.				−0.2079*	−0.0487	−0.1778
				(0.1203)	(0.1368)	(0.1165)
R^2	0.3206	0.3394	0.2343	0.2810	0.3295	0.3670
Adj. R^2	0.2255	0.2315	0.1442	0.2091	0.2038	0.3088
Num. obs.	58	58	58	67	58	96

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: First differences of logged aggregate real rye price. Standard errors in (). The Breusch-Godfrey test cannot reject the null hypothesis of no serial correlation, the Breusch-Pagan test does not detect heteroscedasticity. All variables except dummy in log differences; dummy in first differences. CDR: crude death rate; temp.: temperature; prec.: precipitation.

In line with the results for the previous sub-period (1651–1720; Table S28), Winter precipitation is again significant and of the same sign. In addition, lagged autumn temperature shows a negative association with prices but only when data from 1690 is included; moreover, the coefficient is not significant at conventional levels ($p = 0.131$). Model 1 in Table S29 also passes the RESET procedure. We ran a range of additional robustness checks for both sub-periods (further alterations of functional form such as inclusion of quadratic terms, additional potential weather variables such as summer temperature). However, we found our results not altered in an important way.

Generally, the estimates presented in Tables S28 and S29 should not be interpreted as causal effects of the particular weather variables on the real rye price. This is because additional variables that determine yield such as evapotranspiration etc. (Albers et al., 2017) (and which determine prices through yield) are omitted from this regression due to data limitations. Nevertheless, it is noteworthy that the results of these reduced form regressions are in line with evidence from long-term agronomic field experiments for winter rye yields (Chmielewski, 1992; Chmielewski and Köhn, 2000), which we discuss below. Note that in our reduced form regressions the dependent variable is the real price and thus, all signs are inverted compared to yield effects.

For the period 1651–1720, lower winter temperatures are associated with a higher aggregate real rye price; however, this relationship disappears after 1720. Although the winter period is not part of the growing season, very low temperatures can lead to a reduction of tillers and thus crop density (Chmielewski and Köhn, 2000, 258).

Furthermore, high winter precipitation is related negatively to experimental yield data (Chmielewski, 1992, 27), which corresponds to higher prices in our results in both sub-periods. The negative interaction term of winter precipitation and temperature during 1651–1720 (although not significant) could be explained with warm winters in which continued plant growth (Chmielewski and Köhn, 2000, 254–5) might have reduced yield decreasing (price increasing) effects of nutrient leaching. The latter is one potential explanation for negative effects of additional precipitation on cereal yields (Albers et al., 2017, 57).[35] Overall, the presented empirical evidence, although it is based merely on reduced form regression

[35]In addition, the signs for the variables which are not significant a conventional levels are reasonable. For the years 1721–1790, higher autumn temperatures are related to lower prices (but only with $p = 0.131$); higher autumn temperatures are beneficial for the early development of the plant before winter rest (Chmielewski, 1992, 27; Chmielewski and Köhn, 2000, 258). Higher spring temperatures are associated with a decreasing aggregate price (but $p = 0.12$), reflecting increasing output possibly due to an earlier start of the growing season after winter rest approximately in March (Chmielewski, 1992, 27; Chmielewski and Köhn, 2000, 258). The latter results are also in line with recent work of Esper et al. (2017, 48–9) who relate grain prices and weather variables at a larger geographical scale.

models, shows that weather shocks affected the German aggregate rye price, an important macroeconomic indicator of the pre-industrial era.

SA9 Additional results for unbalanced sample 1651–1790

In what follows, we provide the main results for a much larger unbalanced sample of 29 cities for the period 1651–1790. Both results, price convergence before the French Revolution and the Great Moderation of Grain Price Volatility, remain robust.

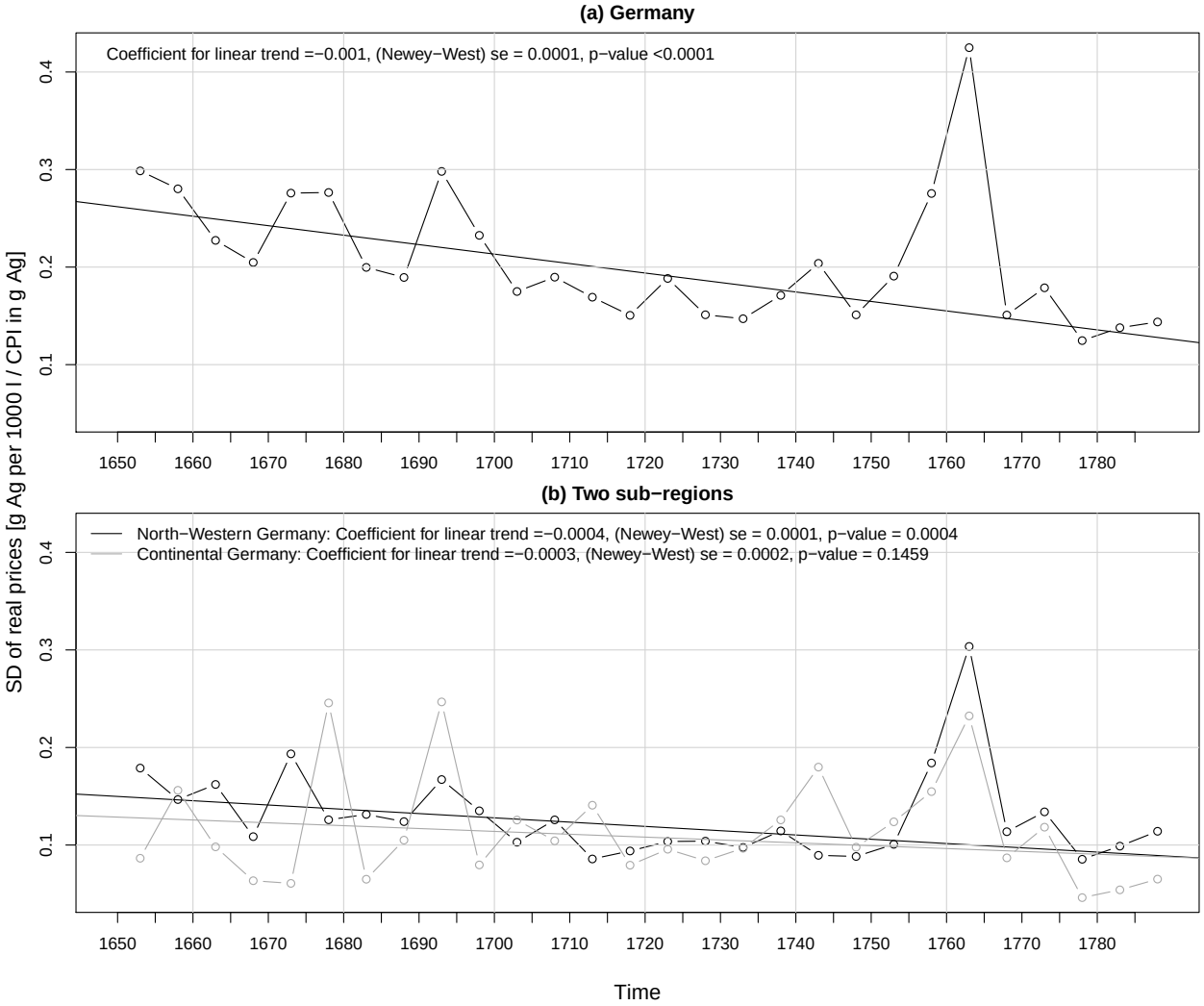

Figure S8: Inter-urban price dispersion 1651/5–1786/90. Cross-sectional standard deviation of real 5-year-mean-prices, rye (unbalanced sample). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

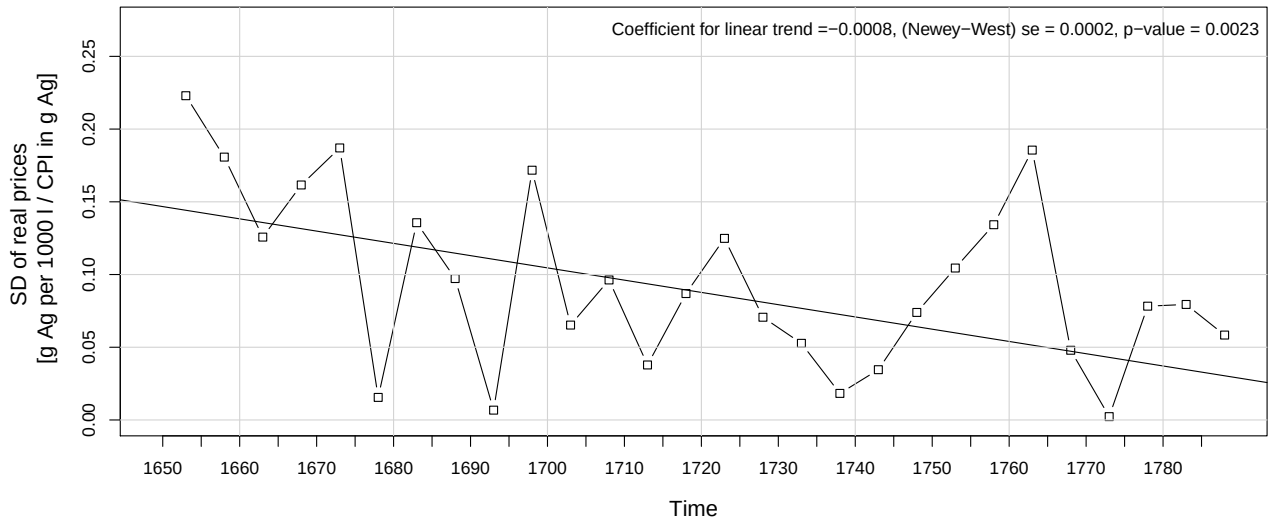

Figure S9: Inter-regional price dispersion 1651/5–1786/90. Between-region standard deviation (square root of variance between North-Western and continental Germany); real 5-year-mean-prices, rye (unbalanced sample). Regression for linear trend includes a dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

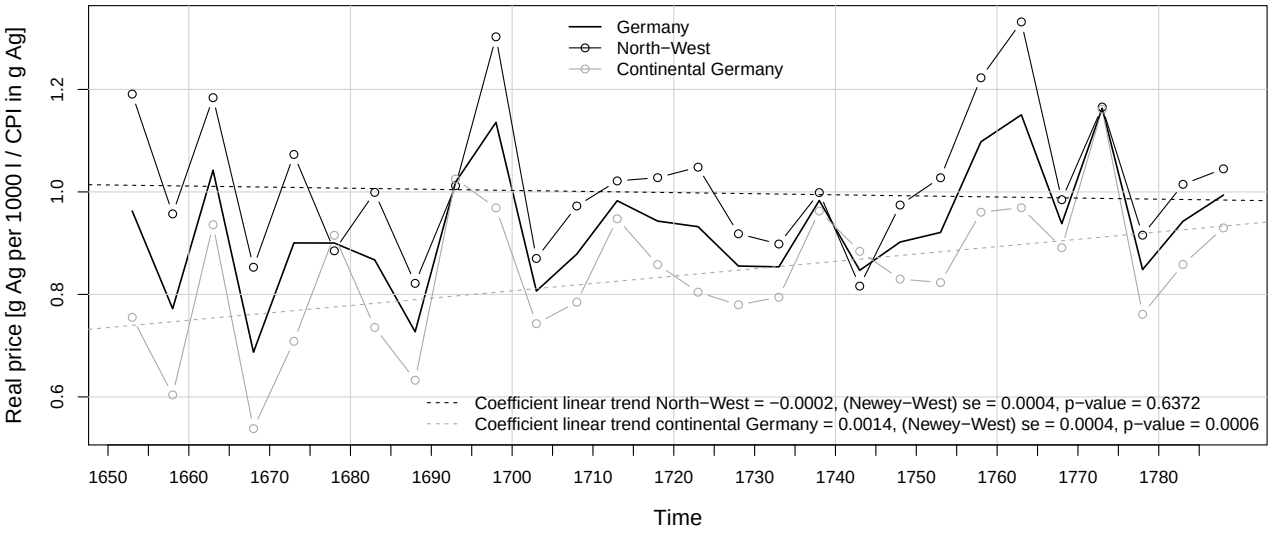

Figure S10: Real 5-year-mean-prices for Germany, North-Western and continental Germany, rye (unbalanced sample) 1651/5–1786/90. Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

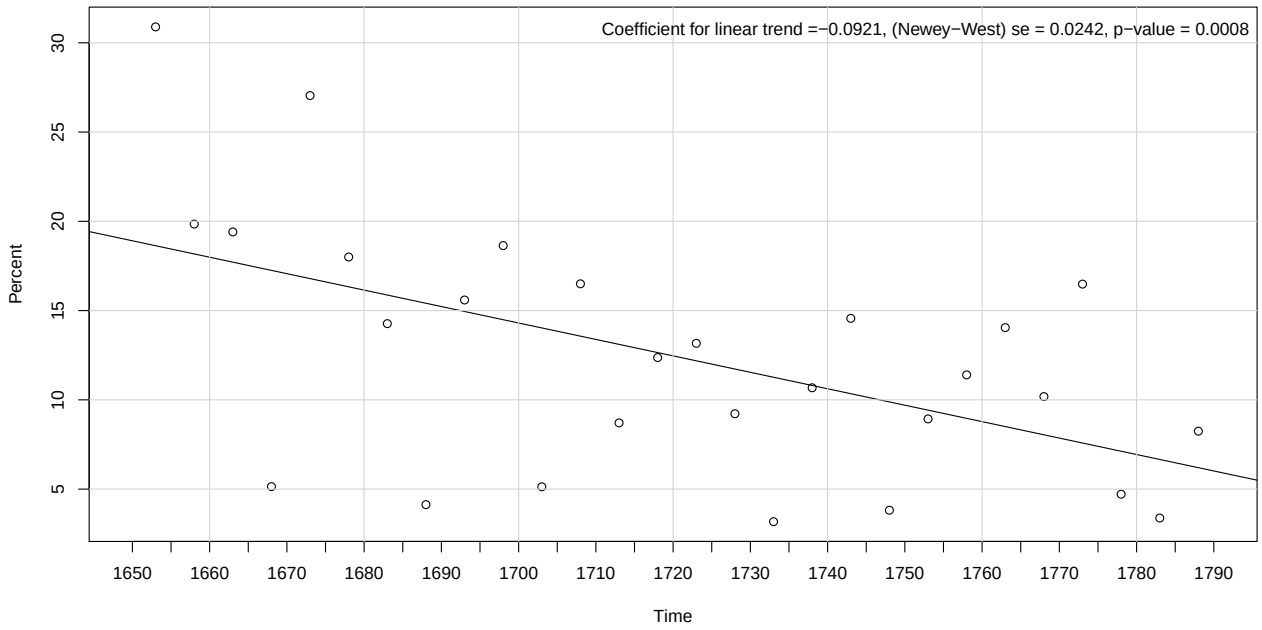

Figure S11: The Great Moderation of Grain Price Volatility 1651/5–1786/90. Volatility of aggregate real rye price Germany (unbalanced sample). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

SA10 Additional results for price convergence at annual frequency

Figure S12 shows price convergence measured as cross-sectional SD in 1651–1790 (stable sample, 15 cities, black solid line).[36] We can exclude analytically that *symmetric absolute* weather shocks or climate change leading to *symmetric absolute* price changes affect the SD.

[36]The strong peak around the year 1760 can be explained in part by monetary financing of the Seven Years' War, which is not correctly reflected in silver conversion rates.

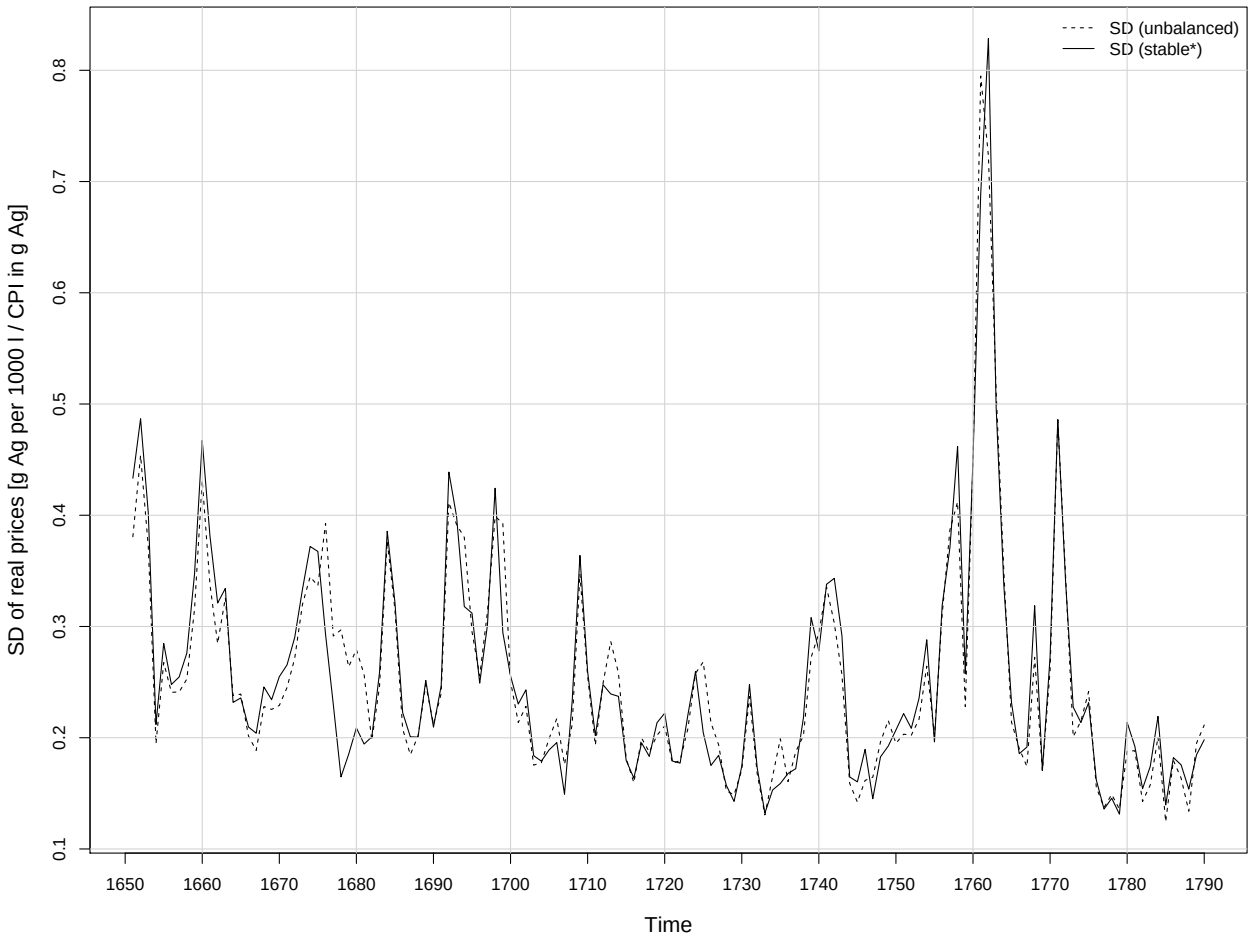

Figure S12: Inter-urban price dispersion: cross-sectional standard deviation (SD) of real rye prices Germany. *: $\leq$5% missing observations per individual series. Each cross-sectional nominal price series is deflated with the national CPI from Pfister (2017).

In addition, Figure S13 shows that the theoretical possibility of a divergence between the trajectories of CV and SD described in our formal analysis (see Section 4.1 in main paper and Table 1) is also borne out in our data. To render SD and CV comparable, each series is calculated for exactly the same data (stable sample). Then we index each series by dividing it by the observation for the year 1651. During several sub-periods SD (black solid line) and CV (grey solid line) move in opposite directions, such as during the 1650s and the 1700s. Several high values of the CV before 1700 lead to a larger negative trend estimate compared to the SD. That is, the CV reports a higher degree of convergence for the period 1651–1790. The trend estimate of the SD is remarkably close to the one based on the time series of the cross-sectional median absolute deviation (MAD), which is a robust measure of dispersion (e.g., Schlittgen, 2012, 248). The empirical differences between the trend estimates of CV and SD underscore the relevance of our methodological discussion.

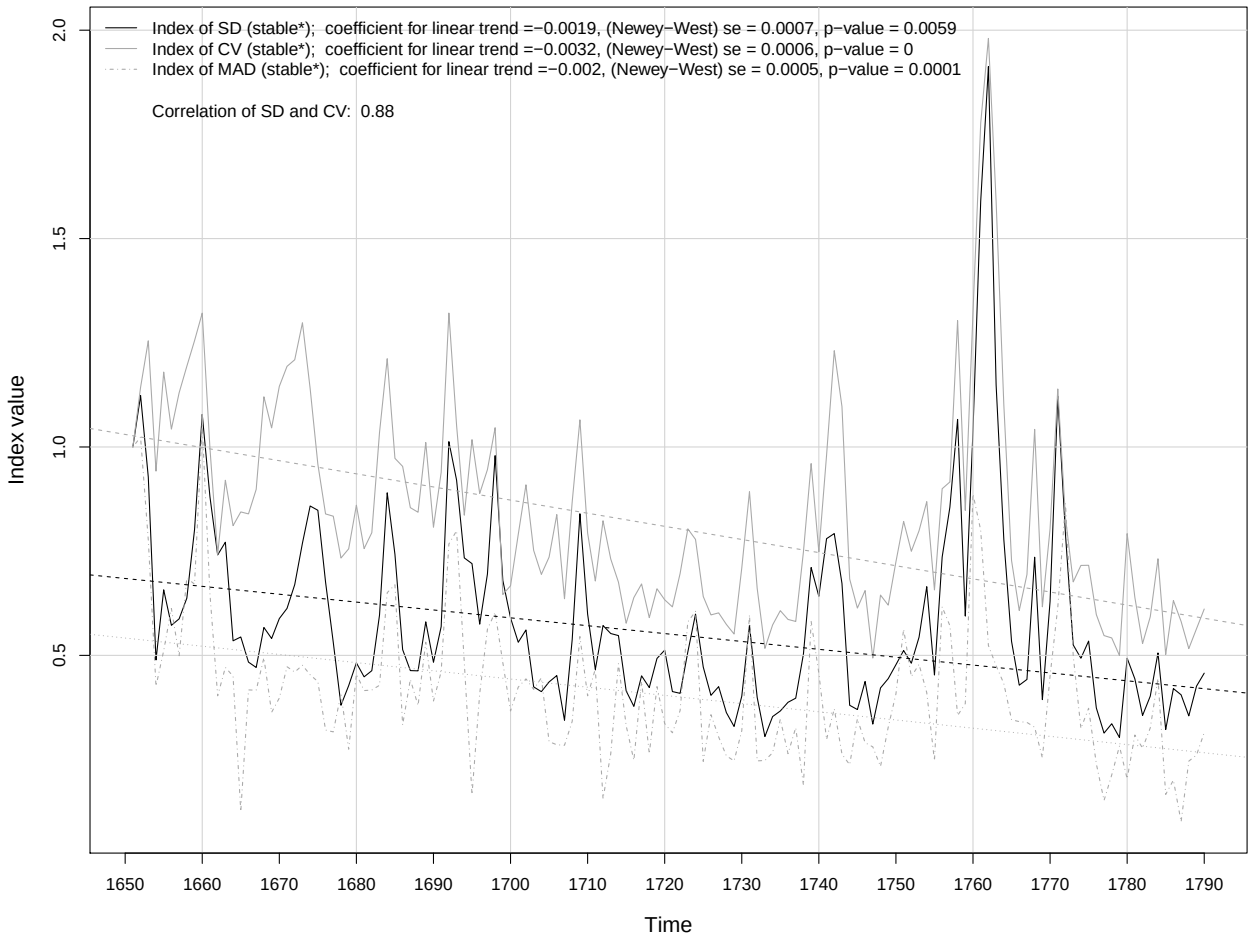

Figure S13: Inter-urban price dispersion: cross-sectional standard deviation, coefficient of variation and median absolute deviation of real rye prices Germany, stable sample ($\leq$5% missing observations per individual series). Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Each cross-sectional nominal price series is deflated with the national CPI from Pfister (2017).

The following plots show the results for not indexed CV (Figure S14) and MAD (Figure S15).

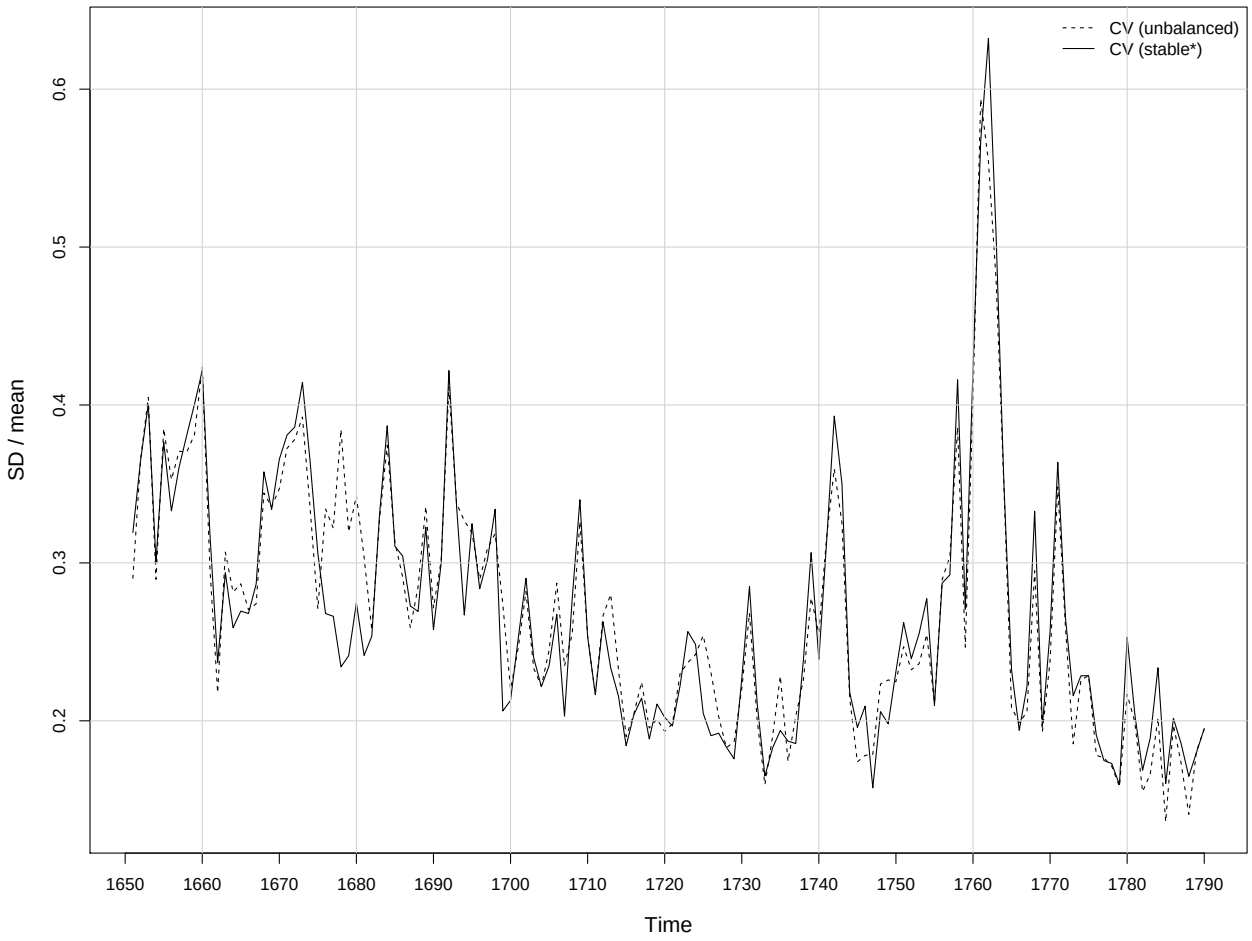

Figure S14: Inter-urban price dispersion: cross-sectional coefficient of variation (CV) of real rye prices Germany. *: $\leq$5% missing observations per individual series. Each cross-sectional nominal price series is deflated with the national CPI from Pfister (2017).

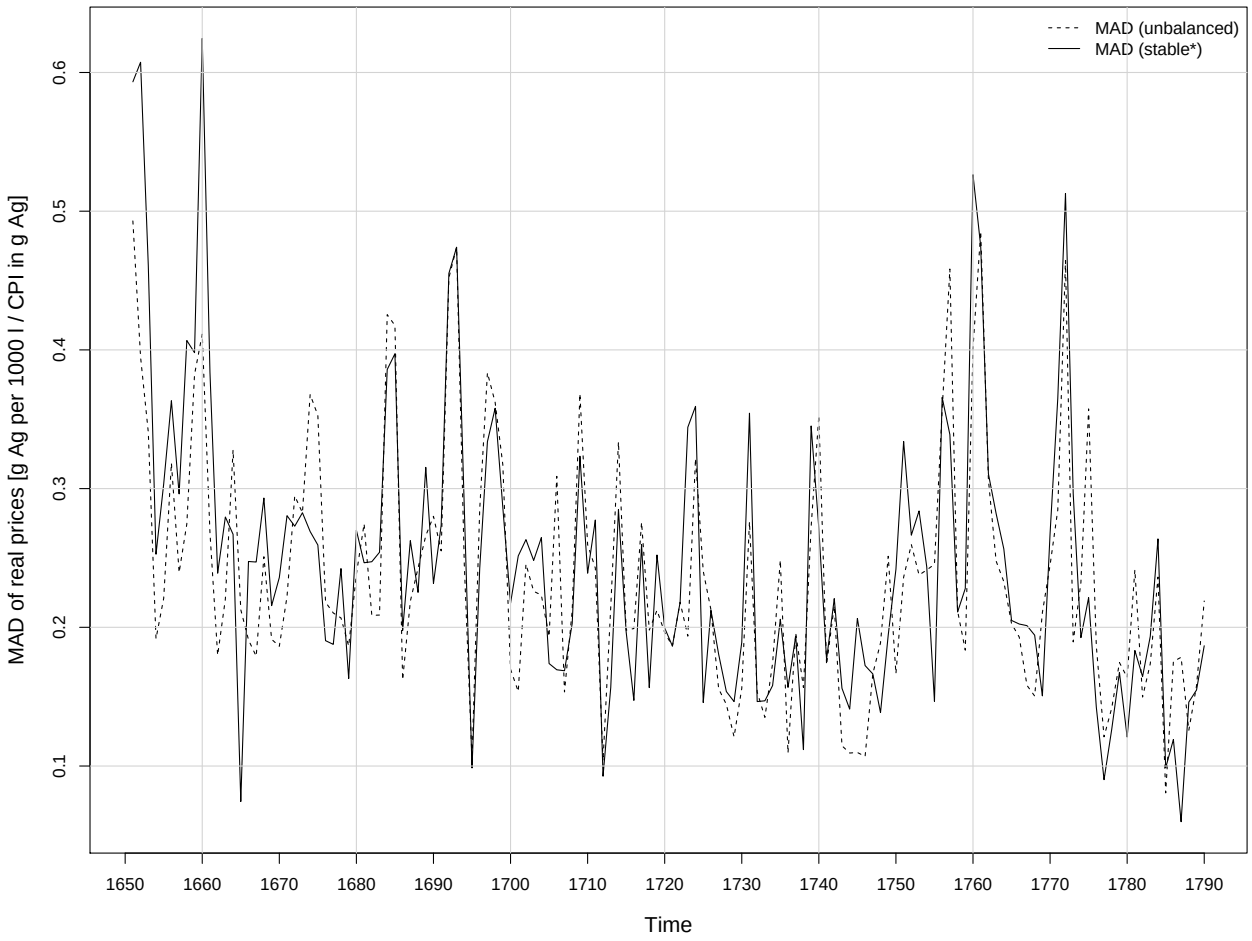

Figure S15: Inter-urban price dispersion: cross-sectional median absolute deviation (MAD) of real rye prices Germany. *: $\leq$5% missing observations per individual series. Each cross-sectional nominal price series is deflated with the national CPI from Pfister (2017).

SA11 Additional panel results for the Great Moderation of Grain Price Volatility

Table S30 shows the results of the following model: $v_{it} = \beta y_t + \alpha_i + u_{it}$. Herein v_{it} denotes the volatility of the real rye price for city i in the five-year-period t; variable y_t is the central year of five year periods (e.g., 1653 for the five-year-period 1651–1655) and thus counts the number of years. Hence, the parameter β quantifies by how many percentage points volatility decreases per year. The city specific error component α_i is removed by within-transformation of the data (Wooldridge, 2009, 481).

Table S30: Panel analysis of real rye price volatility

	Dependent variable: Volatility
Trend	-0.0698 $(0.0254)^{***}$
R^2	0.0852
Adj. R^2	0.0511
Num. obs.	419

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Ordinary least squares on within transformed data. SCC standard error in ().

SA12 Additional results based on the river systems Elbe and Rhine

The following Figures shows the results of the variance decomposition with four sub-regions and two aggregate regions discussed in Section 6 of the main paper.

Additional information on Paderborn: While Paderborn is located at the Lippe, a tributary of the Rhine, the upper part of the Lippe beyond the city of Haltern was separated by natural barriers from the lower part and difficult ship due to about 16 water mills using dams (Bremer, 2001, 19–22). Specifications with Paderborn in the Rhine region and in the region North-East lead to similar results.

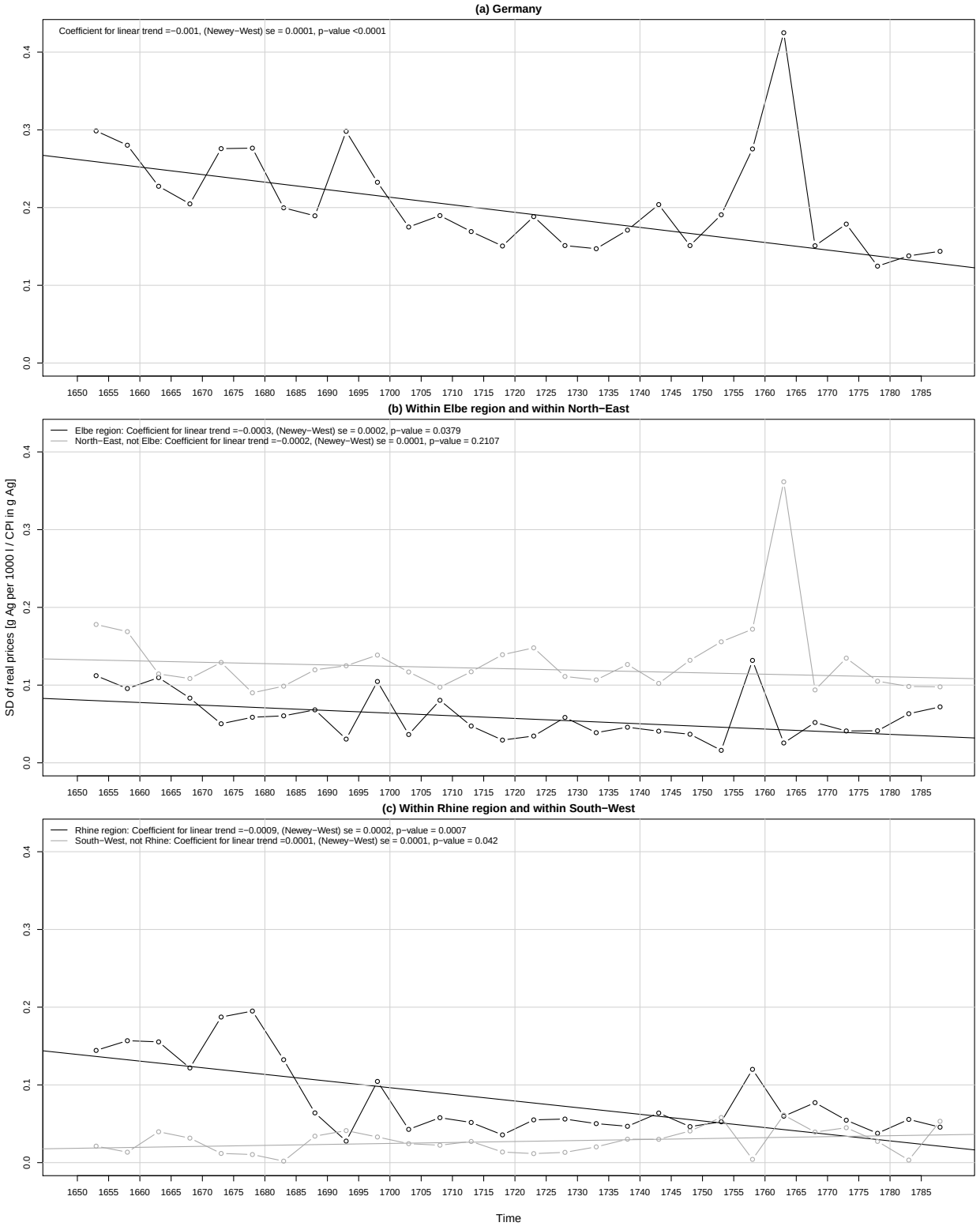

Figure S16: Inter-urban price dispersion 1651/5–1786/90. Cross-sectional standard deviation of real 5-year-mean-prices, rye (unbalanced sample). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Regressions for linear trends include dummy variable for the Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

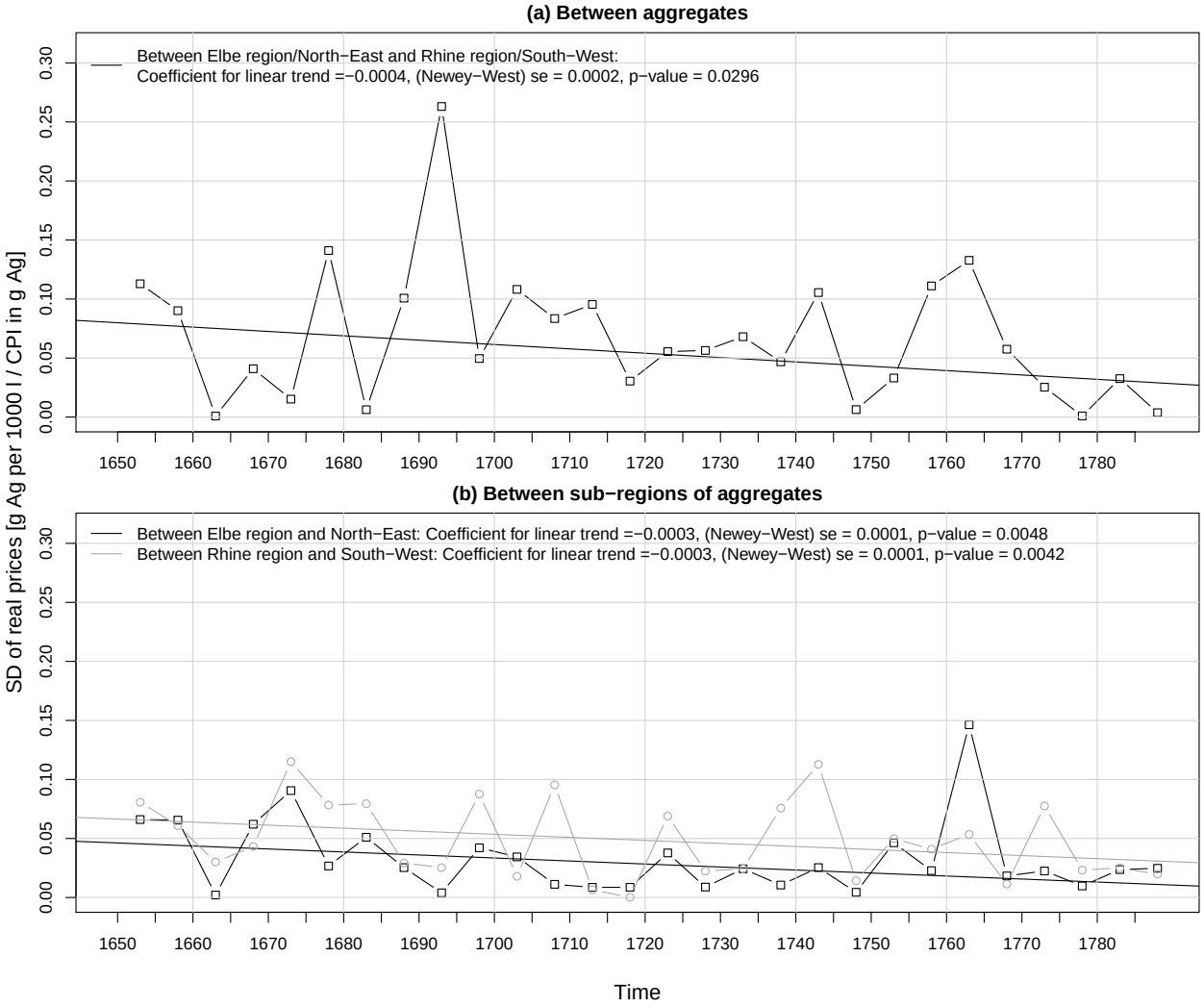

Figure S17: Inter-regional price dispersion 1651/5–1786/90. Between-region standard deviations (square root of between variances); real 5-year-mean-prices, rye (unbalanced sample). Regressions for linear trends include a dummy variable for the Seven Years' War (1756–1763); for panel (a) also for the crisis of 1690/91. Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

SA13 Additional results: split North-Western and continental Germany

In the variance decomposition referring to the river systems of Elbe and Rhine (Section 6 of the main paper), convergence within North-Western Germany, by construction, can neither appear within either of the four redefined sub-regions nor between the Elbe region and

North-East or between the Rhine-Main region and South-West. To quantify the convergence of North-Western Germany while allowing for more regional depth, we performed the four-region-variance-decomposition from the main paper in an alternative way. We revert to North-Western and continental Germany but split North-Western and continental Germany into two sub-regions each: North and West and South and East. All cities in 'South' and 'East' remain the continental cities. The results show a strongly significant downward trend between North and West but no significant convergence between South and East (dummy variables for 1690/91 and the Seven Years' War included).

A plausible trade connection between North and West was the following one. Münster is a city that is part of 'North' and close to the city of Haltern (ca. 50 km). The latter was an important reloading point from waterway to inland transport via road at least around 1600 (Bremer, 2001, 22, 96). The relevant waterway was the Lippe that was throughout navigable for larger ships on its lower part until Haltern and a tributary of the Rhine. The Rhine provided the connection to Xanten or Cologne (both cities in 'West').

The cities of the four regions are listed below for convenience:

1. North: Hamburg, Emden, Lueneburg, Celle, Hannover, Osnabrueck, Minden, Braunschweig, Münster, Paderborn, Goettingen.

2. West: all non-continental cities south of Xanten which are located in Western Germany: Xanten, Aachen, Köln, Trier.

3. South: all continental cities in Southern Germany: Frankfurt, Wuerzburg, Speyer, Nuremberg, Strasboug, Augsburg, München.

4. East: Quedlinburg, Halle, Leipzig, Dresden, Berlin, Gdansk, Königsberg.

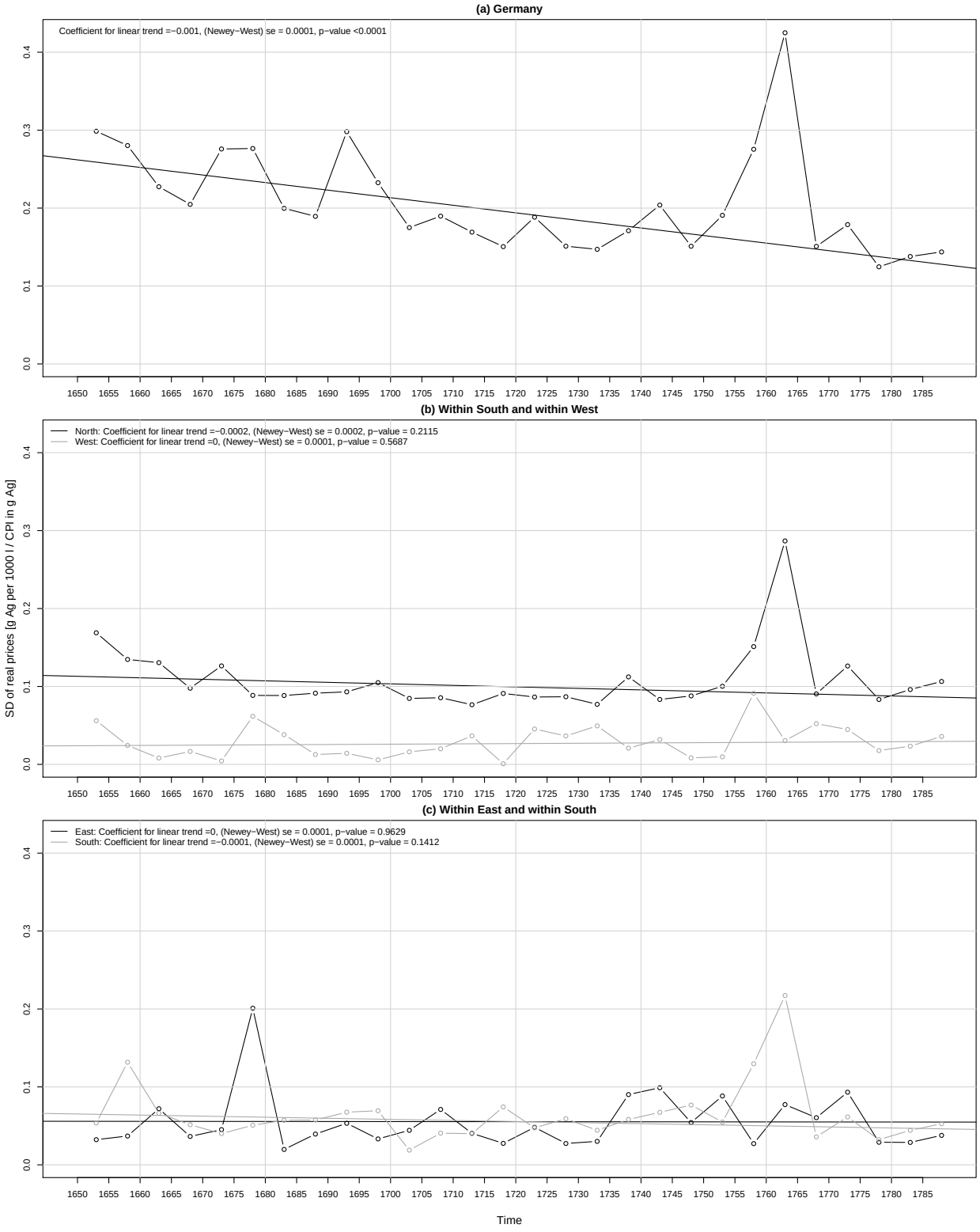

Figure S18: Inter-urban price dispersion 1651/5–1786/90. Cross-sectional standard deviation of real 5-year-mean-prices, rye (unbalanced sample). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

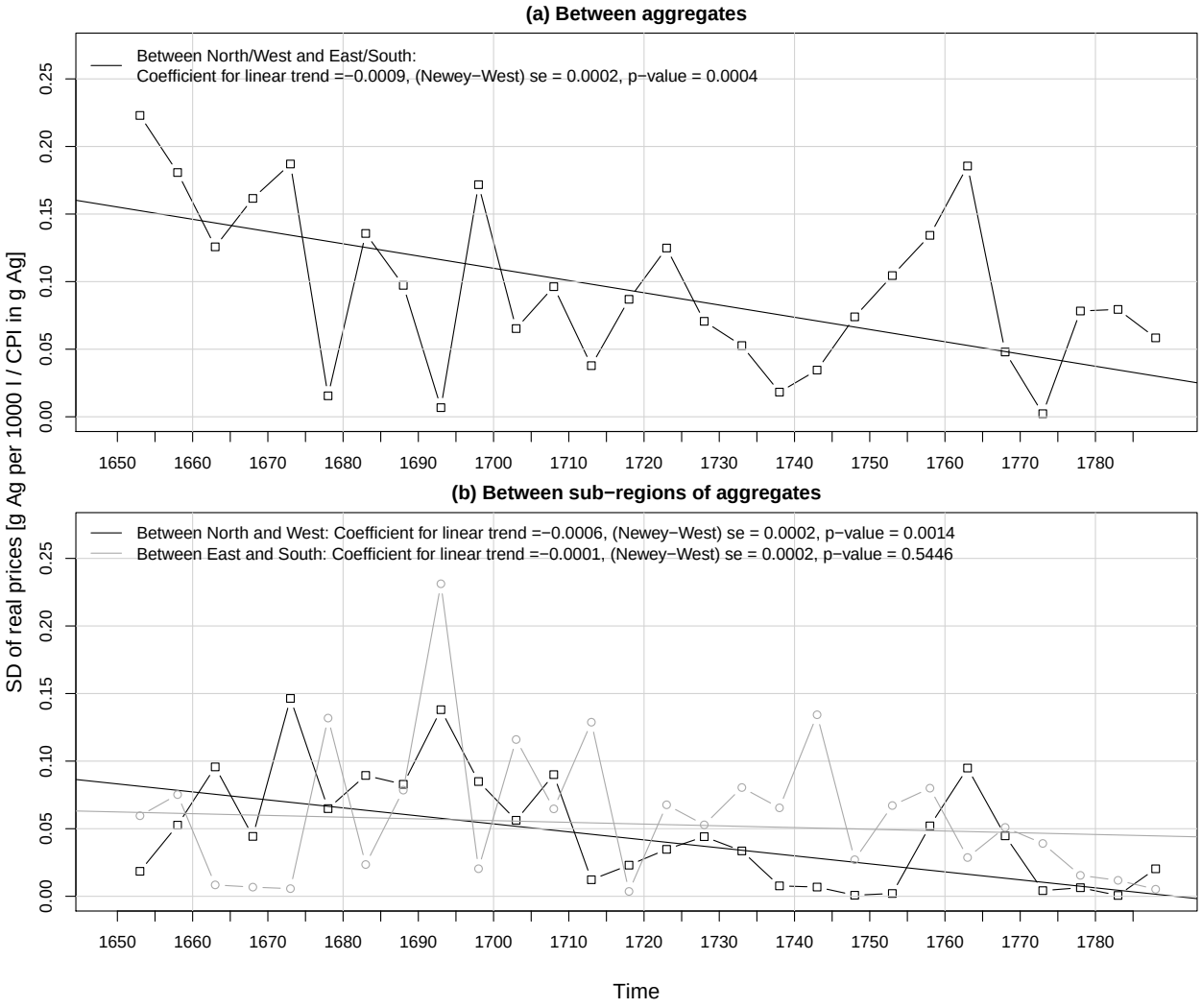

Figure S19: Inter-regional price dispersion 1651/5–1786/90. Between-region standard deviations (square root of between variances); real 5-year-mean-prices, rye (unbalanced sample). Regressions for linear trends include dummy variables for the Seven Years' War (1756–1763) and the crisis of 1690/91. Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

SA14 Additional results: long-run sample of rye prices 1576–1790

This extension explores two questions. First, in how far does price convergence between 1651 and 1790 reflect 'only' the reconstruction of the state of market integration before the Thirty Years' War (1618–1648)? Second, what was the relative importance of climate change vs. market integration in the Great Moderation of Grain Price Volatility? Answers to both

questions rely on the backwards extension of the sample period to 1576 and are discussed in Sections SA14.1 and SA14.2, respectively.

While we can present one of the most extensive grain price data sets for Germany, data coverage before 1650 is clearly very limited. A stable sample ($\leq$5% missing observations per individual series) of rye prices includes seven, mostly North-Western cities[37] for the period 1576–1790. On average there are about eight additional cities in the unbalanced sample for the period 1576–1650 (Figure S1; for comparison: stable sample 1651–1790: 15 cities; unbalanced: additional six cities on average). We provide the results in Section SA14.3.

SA14.1 Reduction of price convergence and volatility relative to the period before the Thirty Years' War

The results for price convergence are less strong but the difference in the trend estimates compared to the baseline result is not very large. On average, prices converge by 0.25% per year (Figure S20; 0.24% per year in unbalanced sample). We estimated this trend while accounting for the increased heterogeneity associated with the Thirty Years' War with additional dummy variables. We use one dummy variable for the entire war period 1618–48 and an additional one for the Kipper and Wipper inflation in 1620–23. The somewhat smaller parameter is compensated by a substantially longer sample period so that a similar reduction of inter-urban price dispersion occurred compared to the baseline period 1651–1790. In relative terms, the cross-sectional dispersion was 31% from 1576–1601 and it decreased to 21% in 1766–1790.

The most important difference with respect to the baseline result is that the volatility moderation is attenuated much and that the estimated trend becomes insignificant at conventional levels ($p = 0.11$). Aggregate volatility is still larger for the 25 years before the Thirty Years' War 1591–1615 (15.6%) than for the years prior to the Napoleonic Wars 1766–1790 (11.6%); but obviously the decrease of four percentage points or ca. 25% is much smaller than in our baseline result where the reduction was 12 percentage points or 46% (1651–75: volatility in the stable sample was 26%; 1766–90: 14%).

SA14.2 Back-of-the-envelope calculation: sources of the Great Moderation of Grain Price Volatility

While a decline of rye price volatility by 25% is still a relevant magnitude, the question about the relative importance of spatial arbitrage in the 'great moderation' of price volatility

[37] Braunschweig, Cologne, Hamburg, Muenster, Xanten, Augsburg, Gdansk.

compared to other factors emerges. Ideally we would like to have regionally disaggregated data on grain output, inputs, weather shocks, crop diseases and grain prices. Such data in combination with a research design like a natural experiment could yield accurate estimates of the relative importance of market integration and climate change; but these data are not available. Nevertheless, the backwards extension of the sample period in combination with our theoretical treatment of volatility in Appendix A of the main paper and SA16.1.3 allows a back-of-the-envelope calculation. The latter can provide a tentative answer about the relative importance of market integration vs. climate change in accounting for the decline in grain price volatility.

Contrary to intuition, markets did not disintegrate following the Thirty Years' War compared to the period before the War: In 1590–1615, the cross-sectional SD of five-year prices was on average 0.32 while it was 0.30 for the years after the War 1651–1675, that is, practically the same (unbalanced sample: 0.25 and 0.26). A similar result obtains in the recent work on European wheat prices by Federico et al. (2018, 11–12, figures 2 and 3). However, rye price volatility in our data set surged to 26.6% in 1651-1675.

Our formal analysis shows that aggregate volatility must increase with decreasing spatial arbitrage, which was apparently not the case. Alternatively, volatility increases with more frequent, larger or more symmetric shocks. Given this framework, we conclude that the surge in price volatility during the period 1651–1675 compared to the pre-War level must be explained by the changing behavior of exogenous shocks. A potential explanation of increased volatility is the Maunder Minimum, the period of much reduced sunspots and associated lower solar irradiance from 1645 to 1715, which contributed to a series of particularly cold winters and springs (Xoplaki et al., 2005, 2; Masson-Delmotte et al., 2013, 389–90; 392; Schönwiese, 2013, 111, 333–5; see also our discussion of time series properties of weather variables in SA8.1). This entails the conclusion that also a substantial amount of the decline of price volatility after the Thirty Years' War must not be attributed to market integration but rather to a decline in shock severity and/or frequency and/or symmetry connected with the end of the Maunder Minimum; and thus the beginning end of the LIA.[38]

Our back-of-the-envelope calculation on the sources of the 'great moderation' rests on two assumptions. First, we assume a similar shock pattern before the Thirty Years' War and during the period 1766–1790; the end of our sample period. This assumption is impossible to

[38]The Maunder minimum (1645–1715) is entirely part of the period the LIA covers. In particular, the LIA is dated 1450–1850 but the spatial and temporal pattern of the LIA is not uniform across the the Northern Hemisphere (Masson-Delmotte et al., 2013, 389, 409; Kelly and Ó Gráda, 2014, 1374). Following the Intergovernmental Panel on Climate Change, a cooler climate prevailed during the LIA but the potential reasons for its occurrence (internal climate variability vs. external forcing such as solar irradiance) are debated (Bindoff et al., 2013, 885, 919).

test empirically, because yield data are not available for the period under study. The second assumption, which rests on empirical support, is the practically unchanged cross-sectional price dispersion before and after the Thirty Years' War reported above. That is, the degree of market integration remained roughly constant. Given these assumptions and our theoretical framework, we tentatively attribute 25% (volatility decline between 1591–1615 and 1766–1790 in long-run sample) of the 46% (volatility decline between 1651–1675 and 1766–1790 in baseline result of main paper) to increased spatial arbitrage. The remaining decline of 21% (= 46% − 25%) is attributed to less severe shocks, the alternative hypothesis in our theoretical framework. In other words, we use the volatility moderation between the pre-War period and the end of our sample period to net out the role of more severe shocks that likely occurred together with the Maunder Minimum (1645–1715) but then presumably petered out during the eighteenth century.

The appropriateness of this guestimate depends on the assumption of similar shock patterns before the Thirty Years' War and at the end of the sample period. If the LIA also manifested itself in harsher shocks before the Thirty Years' War compared to to the years 1766–1790, we overestimate the contribution of market integration. Overall, increasing spatial arbitrage contributed to 'the great moderation of grain price volatility' but a fraction of similar magnitude of the total reduction can be attributed to the reduction of exogenous shocks. Admittedly, while this tentative calculation attributes the reduction in volatility either to market integration or shocks, the set of alternative hypotheses is potentially larger and might also include changes in storage possibilities or behavior (Federico, 2012, 484).

SA14.3 Results 1576–1790

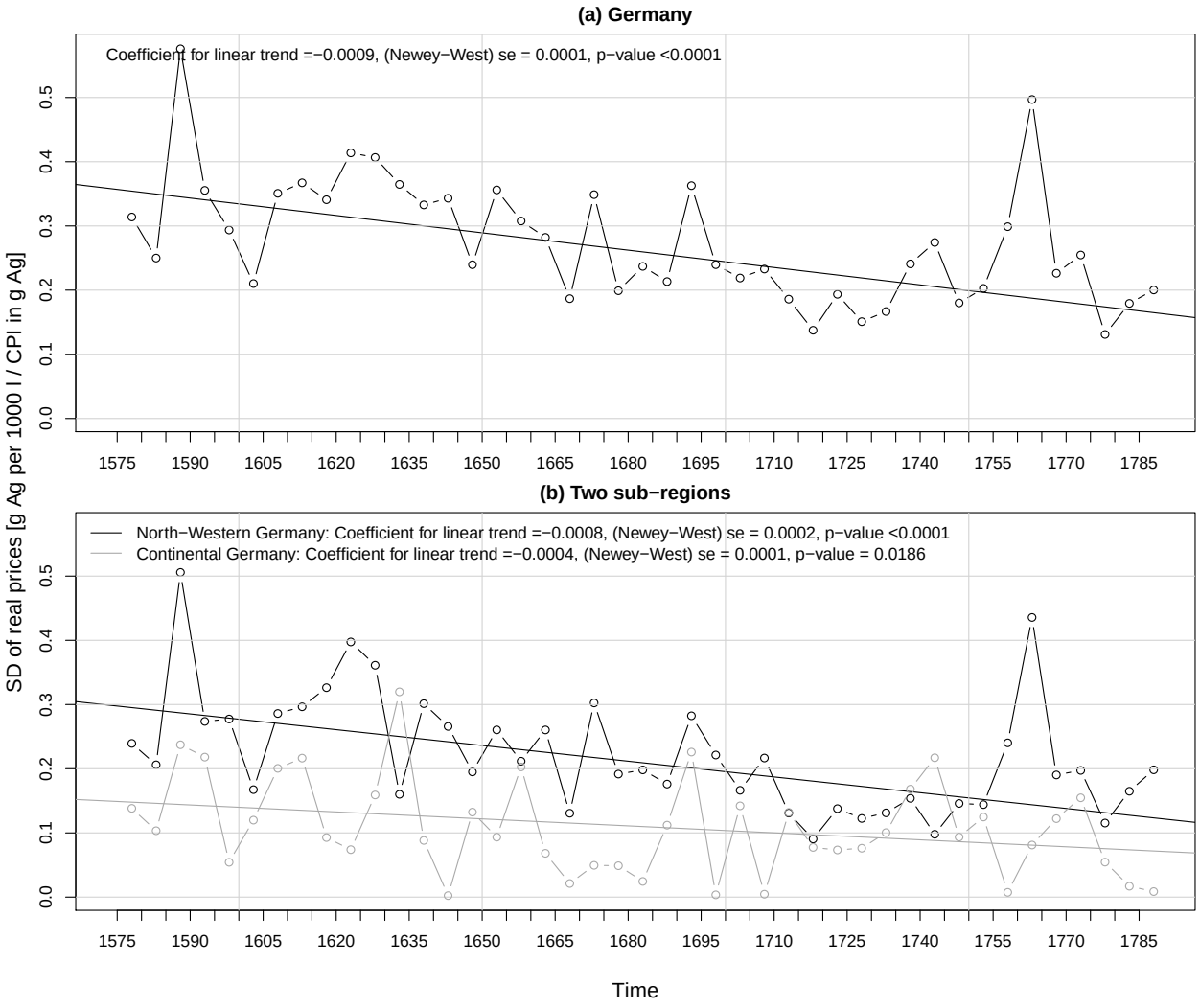

Figure S20: Inter-urban price dispersion 1576/80–1786/90. Cross-sectional standard deviation of real 5-year-mean-prices, rye (stable sample, 7 cities). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Regressions for linear trends include dummy variables for Thirty Years' War (1618–48), the Kipper and Wipper inflation (1620–23) and the Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

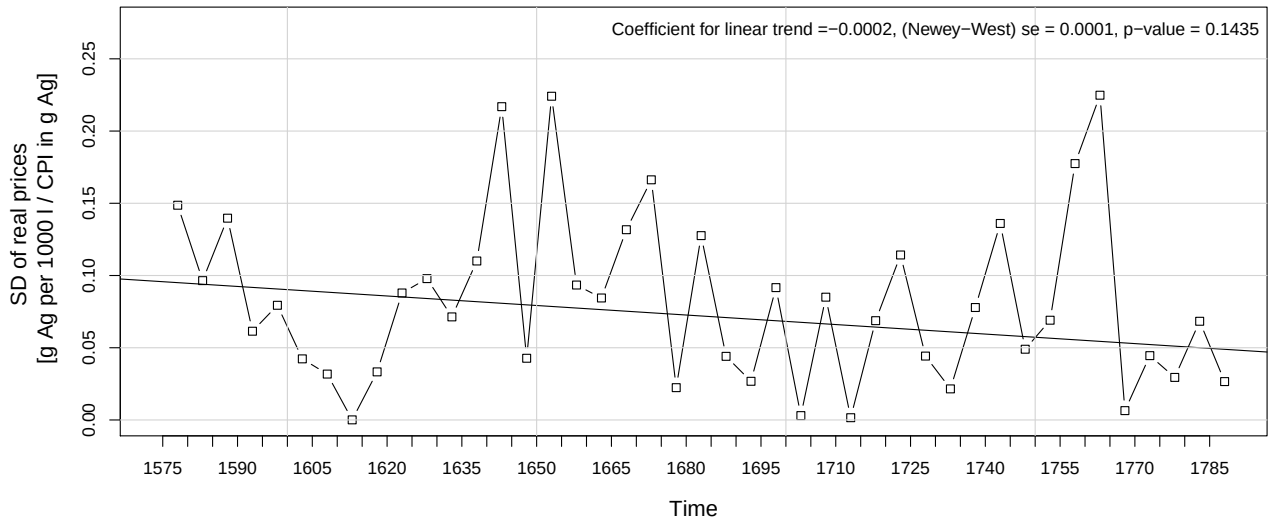

Figure S21: Inter-regional price dispersion 1576/80–1786/90. Between-region standard deviation (square root of variance between North-Western and continental Germany); real 5-year-mean-prices, rye (stable sample, 7 cities). Regression for linear trend includes dummy variables for Thirty Years' War (1618–48), the Kipper and Wipper inflation (1620–23) and the Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

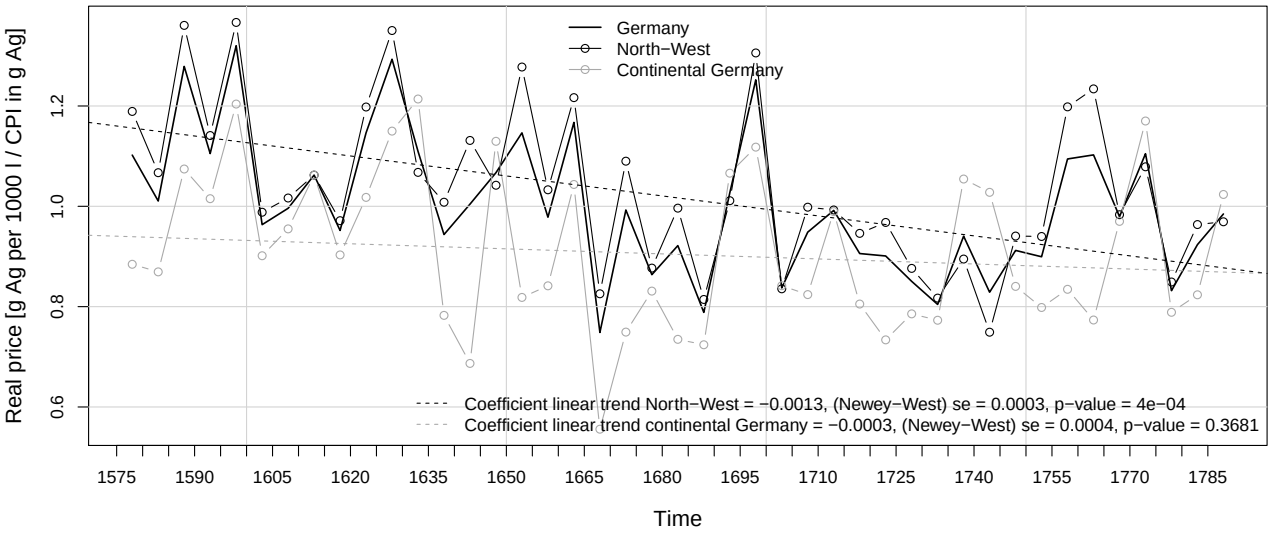

Figure S22: Real 5-year-mean-prices for Germany, North-Western and continental Germany, rye (unbalanced sample) 1576/80–1786/90. Regressions for linear trends include dummy variables for Thirty Years' War (1618–48), the Kipper and Wipper inflation (1620–23) and the Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

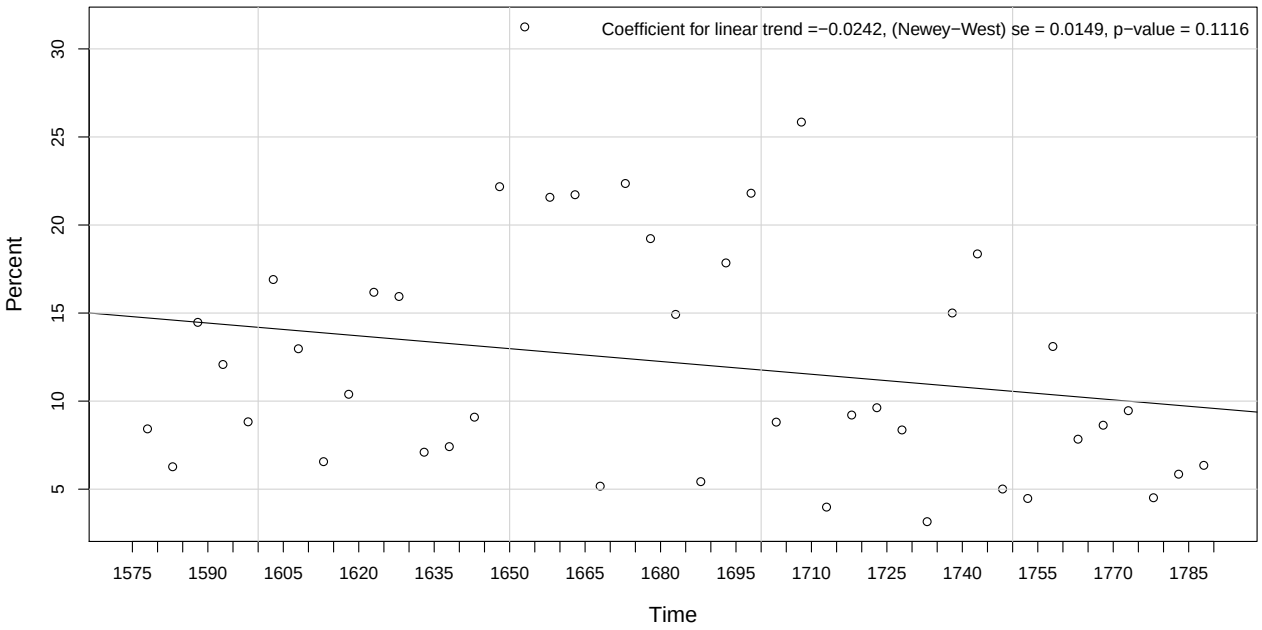

Figure S23: Volatility of aggregate real rye price Germany 1576/80–1786/90 (stable sample, 7 cities). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

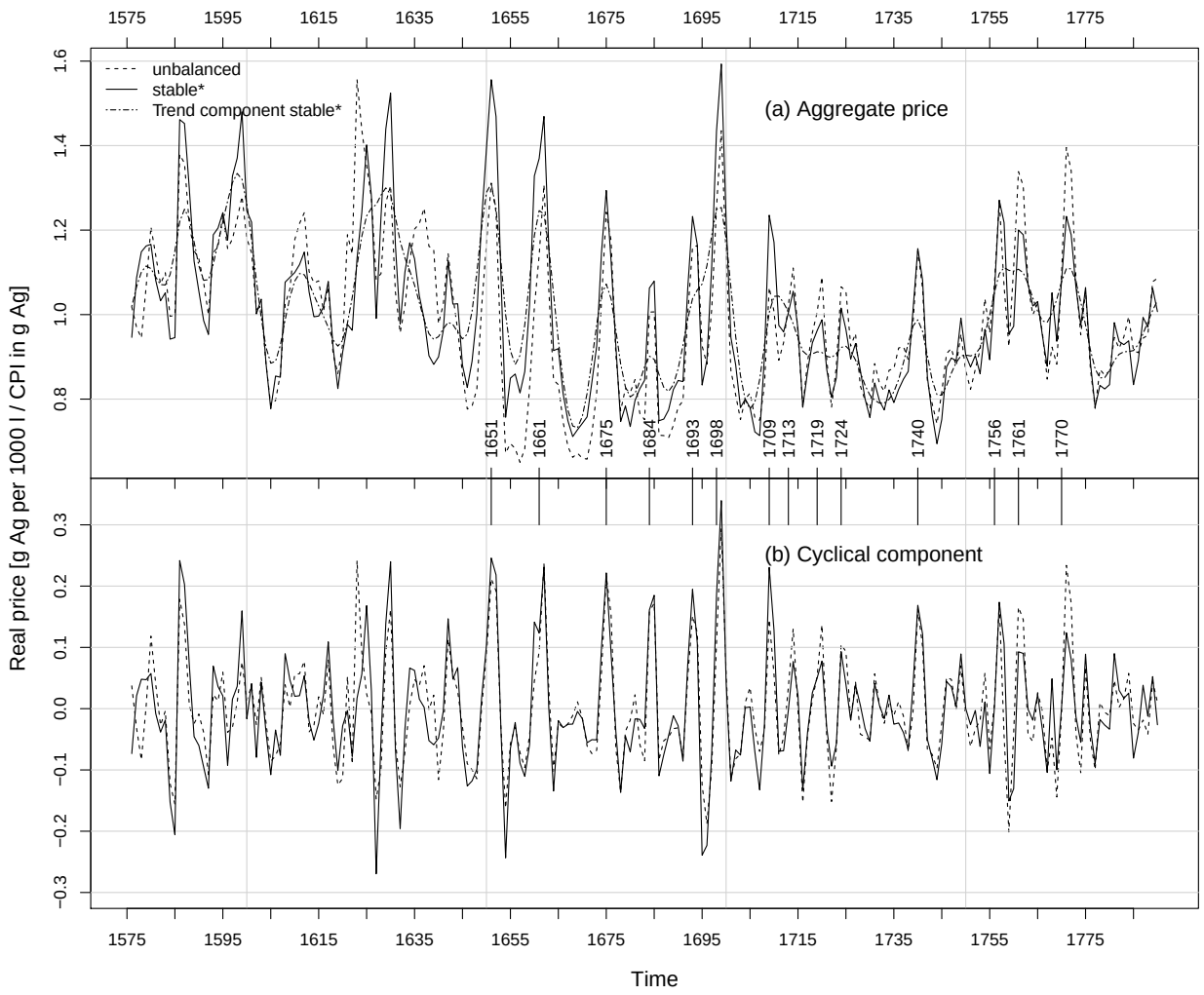

Figure S24: Aggregate real rye price Germany. Stable sample includes 7 cities, 1576–1790. Trend (shown for stable sample series only) and cyclical component from Hodrick-Prescott-Filter, $\lambda = 6.25$ (Ravn and Uhlig, 2002). Vertical lines and given years on upper horizontal axis in panel (b) mark major price peaks associated with subsistence crises. *: $\leq$5% missing observations per individual series. The aggregate nominal rye price is deflated with the CPI from Pfister (2017). Data sources: see SA2.

SA15 Supporting information: anatomy of the CV with proportional shock

We extend the formal analysis of the CV by relaxing the assumption of an additive absolute shock. Instead we assume a shock proportional to the initial price level. The effects depend on whether arbitrage between cities takes place or not and thus, we distinguish four cases: symmetric and asymmetric shock, each one with and without arbitrage.

Symmetric proportional shock

The symmetric proportional shock is defined as the factor $\tau_t = 1 + r_t$, where r_t is the rate at which all prices increase (decrease).

No arbitrage

The effect of a symmetric shock τ_t to the prices in all cities on the mean price is:

$$\begin{aligned} \bar{p}_t^z &= \frac{(p_{1t}\tau_t) + ... + (p_{Nt}\tau_t)}{N} \\ \bar{p}_t^z &= \frac{p_{1t} + ... + p_{Nt}}{N}\tau_t = \bar{p}_t\tau_t \end{aligned} \tag{S10}$$

The shock τ_t does not cancel from the sum of squared deviations:

$$\begin{aligned} \sum_{i=1}^{N}(p_{it} - \bar{p}_t^z)^2 &= (p_{1t}\tau_t - \bar{p}_t\tau_t)^2 + ... + (p_{Nt}\tau_t - \bar{p}_t\tau_t)^2 \\ &= (p_{1t} - \bar{p}_t)^2\tau_t^2 + ... + (p_{Nt} - \bar{p}_t)^2\tau_t^2 \\ &= (p_{1t} - \bar{p}_t)^2 + ... + (p_{Nt} - \bar{p}_t)^2 + N\tau_t^2 \end{aligned} \tag{S11}$$

Consequently, the symmetric shock is in both numerator and denominator of the CV and cancels.

$$CV_t = \frac{\sqrt{\frac{N\tau_t^2}{N-1}}\sigma_t}{\bar{p}_t\tau_t} = \frac{\sqrt{\frac{N}{N-1}}\sigma_t}{\bar{p}_t} \tag{S12}$$

Because $\lim_{N\to\infty}\sqrt{N/(N-1)} = 1$, the initial CV obtains. If the population standard deviation (N in the numerator) is used instead of the empirical standard deviation ($N-1$ in the numerator), the CV obtains directly. A proportional price shock affecting all markets with the same factor τ_t has no effect on the CV. In the context of measuring inequality, this property is known as the relative income principle (Ray, 1998, 178, 188). While historical city samples studying market integration have a relatively small N, the remaining effect of $\sqrt{N/(N-1)}$ is also small and independent of the size of the shock τ_t. For example, with 15 cities (as in our case): $\sqrt{N/(N-1)} \approx 1.035$.

If the the empirical standard deviation is used in the CV, the remaining increase of the CV is very small. Contrary, the SD σ_t increases by factor $\sqrt{N/(N-1)}\tau_t$ for $\tau_t > 1/\sqrt{N/(N-1)}$ and decreases for $0 < \tau_t < 1/\sqrt{N/(N-1)}$. This also explains why symmetric monetary inflation, which can be thought of as a proportional shock τ_t, is accounted for by the CV while this is not the case for the SD.

Perfect arbitrage

The proportional shock τ_t alters all local prices by the same percentage but this results in different absolute price gaps for each city. For example, if $\tau_t > 1$ (positive price shock), prices would increase by the same factor. The altered absolute price gaps, however, provide the incentive for arbitrage. Under perfect arbitrage these local gaps are eliminated to the level of prior existing absolute price gaps (caused by unchanged trade costs), so that in each city the absolute local price shock can be written as $\bar{p}_t(\tau_t - 1) = \bar{p}_t(1 + r_t - 1) = s_t$.

For example, let $r_t = 0.1$. Assume the price in city 1 as 5 units, the price will increase due to the shock by 0.5 units. In city 2 the price is 10 units. The local price will increase by 1 unit. Arbitrage will restore the initial absolute price gap of 5 units. The price in each city will increase by 0.1 multiplied with the mean of 7.5 units, that is, 0.75 units. The new price in city 1 after the shock and arbitrage is 5.75, the price in city 2 is 10.75.

$$\begin{aligned}\sum\nolimits_{i=1}^{N}(p_{it} - \bar{p}_t^z)^2 &= (p_{1t} + \bar{p}_t(\tau_t - 1) - \bar{p}_t\tau_t)^2 + ... + (p_{Nt} + \bar{p}_t(\tau_t - 1) - \bar{p}_t\tau_t)^2 \\ &= (p_{1t} - \bar{p}_t)^2 + ... + (p_{Nt} - \bar{p}_t)^2.\end{aligned} \tag{S13}$$

In this case, compared to the prior situation *no arbitrage* the CV is affected, because the numerator (the SD) does not contain the shock τ_t anymore but the denominator does. The CV decreases for $\tau_t > 1$ and increases for $0 < \tau_t < 1$.

Asymmetric proportional shock

The asymmetric shock to the price in city 1 is defined as factor $\tau_{1t} = 1 + r_{1t}$. In this case, the rate r_{1t} at which the local price increases or decreases is location specific.

Perfect arbitrage

As in the case *symmetric shock, perfect arbitrage*, perfect arbitrage leads to equal distribution of the shock across all cities. The intuition is that the local rate r_{1t} combined with the local price p_{1t} leads to a local absolute price increase which is—due to arbitrage—shared by all cities. That is, each of the N cities experiences the Nth share the local absolute price shock

$p_{1t}(\tau_{1t}-1) = p_{1t}r_{1t} = s_{1t}$ caused by the local proportional shock with factor τ_{1t}.

$$\begin{aligned}\bar{p}_t^z &= \frac{p_{1t}+...+p_{Nt}}{N} + \frac{(p_{1t}[\tau_{1t}-1])}{N} \\ &= \frac{p_{1t}+...+p_{Nt}}{N} + \frac{p_{1t}r_{1t}}{N} = \bar{p}_t + \frac{p_{1t}r_{1t}}{N}\end{aligned} \tag{S14}$$

This shock affects the sum of squared deviations as follows:

$$\begin{aligned}\sum_{i=1}^{N}(p_{it}-\bar{p}_t^z)^2 &= (p_{1t}[1+\frac{\tau_{1t}-1}{N}] - [\bar{p}_t + p_{1t}\frac{\tau_{1t}-1}{N}])^2 \\ &+ ... + (p_{Nt} + p_{1t}\frac{\tau_t-1}{N} - [\bar{p}_t + p_{1t}\frac{\tau_{1t}-1}{N}])^2 \\ &= (p_{1t}[1+\frac{\tau_{1t}-1}{N} - \frac{\tau_{1t}-1}{N}] - \bar{p}_t)^2 + ... + (p_{Nt}-\bar{p}_t)^2.\end{aligned} \tag{S15}$$

Thus, the local shock cancels from the sum of squared deviations and hence, the SD is unchanged. This result is equivalent to the case with the absolute shock s_{1t}. The CV is affected as follows. For positive shocks, that is $r_{1t} > 0$, the effect on the mean price is positive (eq. S14) and the CV decreases as well but in a different way. Now $\frac{p_{1t}r_{1t}}{N}$ enters the denominator additively, not multiplicatively as the factor τ_t in the case *proportional symmetric shock, perfect arbitrage*. Whether the denominator decreases to a lesser extent compared with the case *proportional symmetric shock, perfect arbitrage* depends on the relative size of τ_t vs. the relative effect of adding $\frac{p_{1t}r_{1t}}{N}$ to $\bar{p}_t$ in the denominator.

No arbitrage

In this case, the shock τ_{1t} to city 1 does not spread to any other city and thus affects the sum of squared deviations as follows:

$$\begin{aligned}\sum_{i=1}^{N}(p_{it}-\bar{p}_t^z)^2 &= (p_{1t}\tau_{1t} - [\bar{p}_t + \frac{p_{1t}(\tau_{1t}-1)}{N}])^2 + ... + (p_{Nt} - [\bar{p}_t + \frac{p_{1t}(\tau_{1t}-1)}{N}])^2 \\ &= (p_{1t}[\tau_{1t} - \frac{\tau_{1t}-1}{N}] - \bar{p}_t)^2 + ... + (p_{Nt} - \bar{p}_t - \frac{p_{1t}(\tau_{1t}-1)}{N})^2.\end{aligned} \tag{S16}$$

The shock does not cancel from the sum of squared deviations. Both SD and CV change due to an asymmetric proportional shock. As in the case with an asymmetric absolute shock, the sign of the change depends on whether the local price is moved closer to the sample mean or not relative to the situation without shock.

Let the example be again the prices in three cities: 2, 4, 6. Now the price in city 1 increases by 50% so that $r_{1t} = 0.5$ and $\tau_{1t} = 1 + r_{1t} = 1.5$ and the price in city 1 including the shock is $p_{1t}^z = p_{1t}\tau_{1t} = 3$. Now the squared difference of the price with shock and the mean price with shock is *smaller* compared to the situation without shock. The behavior of the squared difference as well as of the SD and the CV is equivalent to the case *asymmetric absolute shock, no arbitrage*.

SA16 Supporting information: spatial arbitrage reduces aggregate volatility

We first provide additional details on the case of an absolute shock, which we discuss in Appendix A of the main paper. Second, we repeat the analysis for the case of a proportional shock.

SA16.1 Absolute shock

SA16.1.1 Additional details on the flat line assumption

Figure S25 illustrates the 'flat line' assumption.

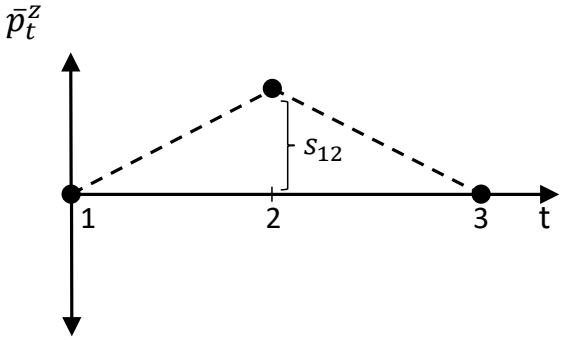

Figure S25: Graphical illustration of analytical framework. $\bar{p}_t^z$: cross-sectional average price in year t inlcuding the shock s_{12} to city 1 in year 2. Source: own representation.

Theoretically, the flat line assumption can be reconciled with the notion of equilibrium. Any deviation from the line, equivalent to the horizontal axis in Figure S25, is a shock. Empirically, we do not observe any long-run trend in the average national real price indicating that the assumption is also reasonable on empirical grounds. If the prices were not equal in all periods without the shock, the algebra is complicated because the squared deviations are different for each sub-period. In this case, the SSD can decrease due to a positive shock (hence, volatility can decrease), if the shocked local price is moved closer to the temporal mean of all T prices (in an analogous manner as in the cross-sectional analysis). But to assume different prices in all periods makes it difficult to define what constitutes a shock.

SA16.1.2 Additional steps on the simplification of $u(.)$.

$$
\begin{aligned}
u(.) &= (\bar{p} - \bar{p} - \frac{as_{12}}{T})^2 + \Big[\bar{p} - \bar{p} + \frac{(Ta-a)s_{12}}{T}\Big]^2 + (\bar{p} - \bar{p} - \frac{as_{12}}{T})^2 + ... \\
&+ (\bar{p} - \bar{p} - \frac{as_{12}}{T})^2 \\
&= (T-1) \cdot (-\frac{as_{12}}{T})^2 + \Big[\frac{(Ta-a)s_{12}}{T}\Big]^2 \\
&= (T-1) \cdot \frac{a^2 s_{12}^2}{T^2} + \frac{(Ta-a)^2 \cdot s_{12}^2}{T^2} \\
&= (T-1) \cdot \frac{a^2 s_{12}^2}{T^2} + \frac{(T^2a^2 - 2Ta^2 + a^2) \cdot s_{12}^2}{T^2} \\
&= (T-1) \cdot \frac{a^2 s_{12}^2}{T^2} + \frac{(T^2 - 2T + 1) \cdot a^2 \cdot s_{12}^2}{T^2} \\
&= (T - 1 + T^2 - 2T + 1) \cdot \frac{a^2 s_{12}^2}{T^2} \\
&= (T^2 - T)\frac{a^2 s_{12}^2}{T^2} \\
&= (T-1)\frac{a^2 s_{12}^2}{T}
\end{aligned}
\tag{S17}
$$

SA16.1.3 Effect of a shock on aggregate volatility

To evaluate the effect of a shock on volatility we need to analyze whether the increase of $u(.)$ is larger than the increase of the denominator $\bar{p}^z$ of the volatility V, so that the shock s_{12} increases volatility. Thus, we take the first derivative of $\frac{u(.)}{\bar{p}^z}$with respect to the shock s_{12} and by use of the Quotient Rule:

$$
\frac{\partial u/\bar{p}^z}{\partial s_{12}} = \frac{\overbrace{2(T-1)\frac{a^2 s_{12}}{T}}^{u'>0} \cdot \overbrace{[\bar{p} + \frac{as_{12}}{T}]}^{\bar{p}^z>0} - \overbrace{[(T-1)\frac{a^2 s_{12}^2}{T}]}^{u>0} \cdot \overbrace{\frac{a}{T}}^{\bar{p}^{z\prime}>0}}{\underbrace{[\bar{p} + \frac{as_{12}}{T}]^2}_{\bar{p}^{z2}>0}} > 0, \tag{S18}
$$

because the numerator can be simplified to:

$$
\begin{aligned}
&2(T-1)\frac{a^2 s_{12}}{T} \cdot \bar{p} + 2(T-1)\frac{a^3 s_{12}^2}{T^2} - (T-1)\frac{a^3 s_{12}^2}{T^2} \\
=&2(T-1)\frac{a^2 s_{12}}{T} \cdot \bar{p} + (T-1)\frac{a^3 s_{12}^2}{T^2} > 0, \text{because } T > 1.
\end{aligned}
$$

Thus, a shock increases volatility. To evaluate a negative price shock $s_{12} < 0$ (which *ceteris paribus* corresponds to a positive output shock, e.g, due to a good harvest) the numerator

is rewritten as:

$$= \overbrace{(T-1)\frac{a^2 s_{12}}{T}}^{<0 \text{ if } s_{12}<0} \cdot \overbrace{(2\bar{p} + \frac{a s_{12}}{T})}^{>0 \text{ if } s_{12} > \frac{-2\bar{p}T}{a}}$$

Since the denominator is squared, it is positive. Thus, $\frac{\partial u/\bar{p}^z}{\partial s_{12}}$ is negative and volatility will decrease as long as the numerator is negative, which is the case as long as $s_{12} > \frac{-2\bar{p}T}{a}$.

The largest s_{12} which can just not satisfy this condition is $s_{12} = -4\bar{p}$ for $T = 2$ (with $T < 2$ we cannot calculate a temporal volatility) and $a = 1$ (no cross-sectional arbitrage; more arbitrage would mean a smaller a and thus a smaller s_{12}) (due to the negative sign). This value, $s_{12} = -4\bar{p}$, is implausibly small, because it means that the shock would reduce the local price by four times the cross-sectional average, because $\bar{p} = \bar{p}_2$. The local price must be smaller than N times the cross-sectional average, however. The number of cross-sectional units is given by $a = \frac{1}{N} = 1$, so $N = 1$. This would mean a reduction of the local price by 400%, which is not possible.[39] Thus, the condition $s_{12} > \frac{-2T\bar{p}}{a}$ is satisfied.

In other words, an increasing s_{12} in case of a negative price shock means that the shock becomes 'less negative' and the price in period 2 with shock, $\bar{p}_2^z$ increases towards the 'flat line' from below. Once the 'flat line' of initial prices is passed, the shock becomes positive and volatility increases as shown above.

SA16.1.4 Effect of spatial arbitrage on aggregate volatility

Additional step on the simplification of the numerator

The numerator can be simplified to:

$$\begin{aligned} &2(T-1)\frac{a s_{12}^2}{T} \cdot \bar{p} + 2(T-1)\frac{a^2 s_{12}^3}{T^2} - (T-1)\frac{a^2 s_{12}^3}{T^2} \\ =&2(T-1)\frac{a s_{12}^2}{T} \cdot \bar{p} + (T-1)\frac{a^2 s_{12}^3}{T^2} > 0, \text{because } T > 1. \end{aligned}$$

SA16.2 Proportional shock

The analysis with a proportional instead of an absolute shock is similar but as in the cross-sectional analysis, the shock τ_{12} to city 1 in year 2 is defined as $\tau_{12} = 1 + r_{12}$ where r_{12} is the shock rate at which the price increases or decreases.

[39] To illustrate further: If the number of cross-sectional units increases to $N = 2$ (and $a = 1/2$), the condition is $s_{12} > -8\bar{p}$. The local price must be smaller than $N = 2$ times the cross-sectional average at $t = 2$ and thus, the shock cannot be 8 time the size of the initial local price.

Effect on mean

The average price over time including the shock is $\bar{p}^z$:

$$\begin{aligned}\bar{p}^z &= \frac{\bar{p}_1 + [\bar{p}_2(1+ar_{12})] + \bar{p}_3 + ... + \bar{p}_T}{T} = \frac{\bar{p}_1 + \bar{p}_2 + \bar{p}_3 + ... + \bar{p}_T}{T} + \frac{ar_{12}\bar{p}_2}{T} \\ &= \bar{p} + \frac{ar_{12}\bar{p}_2}{T}\end{aligned} \tag{S19}$$

Effect on SSD

The SSD including the shock are defined as the function $u(.)$:

$$\begin{aligned}u(.) &= \sum\nolimits_{t=1}^{T}(\bar{p}_t - \bar{p}^z)^2 \\ &= (\bar{p}_1 - \bar{p} - \frac{ar_{12}\bar{p}_2}{T})^2 + \left[\bar{p}_2 + ar_{12}\bar{p}_2 - \bar{p} - \frac{ar_{12}\bar{p}_2}{T}\right]^2 \\ &+ (\bar{p}_3 - \bar{p} - \frac{ar_{12}\bar{p}_2}{T})^2 + ... + (\bar{p}_T - \bar{p} - \frac{ar_{12}\bar{p}_2}{T})^2\end{aligned} \tag{S20}$$

The term for the second period in [.] can be simplified to $\left[\frac{\bar{p}_2\cdot(T+Tr_{12}a-ar_{12})}{T} - \bar{p}\right]^2$. Like in the case absolute shock, the 'flat line' assumption ($\bar{p} = \bar{p}_1 = \bar{p}_2 = \bar{p}_3 = ... = \bar{p}_T$) allows to simplify eq. S20:

$$\begin{aligned}u(.) &= (\bar{p} - \bar{p} - \frac{ar_{12}\bar{p}}{T})^2 + \left[\frac{\bar{p}\cdot(T + Tr_{12}a - ar_{12} - T)}{T}\right]^2 + (\bar{p} - \bar{p} - \frac{ar_{12}\bar{p}}{T})^2 + ... \\ &+ (\bar{p} - \bar{p} - \frac{ar_{12}\bar{p}}{T})^2 \\ &= (T-1)\cdot(-\frac{ar_{12}\bar{p}}{T})^2 + \left[\frac{\bar{p}\cdot ar_{12}(T-1)}{T}\right]^2 \\ &= (T-1)\cdot\frac{a^2r_{12}^2\bar{p}^2}{T^2} + \frac{\bar{p}^2\cdot a^2r_{12}^2(T-1)^2}{T^2} \\ &= \left[(T-1) + (T-1)^2\right]\cdot\frac{a^2r_{12}^2\bar{p}^2}{T^2} \\ &= \left[T - 1 + T^2 - 2T + 1\right]\cdot\frac{a^2r_{12}^2\bar{p}^2}{T^2} \\ &= (T-1)\cdot\frac{a^2r_{12}^2\bar{p}^2}{T}\end{aligned} \tag{S21}$$

To evaluate the effect of a shock on volatility we take the first derivative of $\frac{u(.)}{\bar{p}^z}$ with respect to the shock rate r_{12} and by use of the Quotient Rule:

$$\frac{\partial u/\bar{p}^z}{\partial r_{12}} = \frac{\overbrace{2(T-1)\frac{a^2r_{12}\bar{p}^2}{T}}^{u'>0}\cdot\overbrace{[\bar{p} + \frac{ar_{12}\bar{p}}{T}]}^{\bar{p}^z>0} - \overbrace{[(T-1)\frac{a^2r_{12}^2\bar{p}^2}{T}]}^{u}\cdot\overbrace{\frac{a\bar{p}}{T}}^{\bar{p}^{z\prime}>0}}{\underbrace{[\bar{p} + \frac{ar_{12}\bar{p}}{T}]^2}_{\bar{p}^{z2}>0}} > 0, \tag{S22}$$

because the numerator can be simplified to:

$$\begin{aligned}&2(T-1)\frac{a^2 r_{12}\bar{p}^3}{T}+2(T-1)\frac{a^3 r_{12}^2\bar{p}^3}{T^2}-(T-1)\frac{a^3 r_{12}^2\bar{p}^3}{T^2}\\ =&2(T-1)\frac{a^2 r_{12}\bar{p}^3}{T}+(T-1)\frac{a^3 r_{12}^2\bar{p}^3}{T^2}>0, \text{because } T>1.\end{aligned}$$

Thus, a shock increases volatility. To evaluate a negative price shock $r_{12}<0$ (which *ceteris paribus* corresponds to a positive output shock, e.g, due to a good harvest) the numerator is rewritten as:

$$=\overbrace{(T-1)\frac{a^2 r_{12}\bar{p}^3}{T}}^{<0 \text{ if } r_{12}<0}\cdot\overbrace{(2+\frac{ar_{12}}{T})}^{>0 \text{ if } r_{12}>\frac{-2T}{a}}$$

Since the denominator is squared, it is positive. Thus, $\frac{\partial u/\bar{p}^z}{\partial r_{12}}$ is negative and volatility will decrease as long as the numerator is negative, which is the case as long as $r_{12}>\frac{-2T}{a}$.

The largest r_{12} which can just not satisfy this condition is $r_{12}=-4$ for $T=2$ (with $T<2$ we cannot calculate a temporal volatility) and $a=1$ (no cross-sectional arbitrage; more arbitrage would mean a smaller a and thus a smaller r_{12}). This value, $r_{12}=-4$, is implausibly small, because it means that the shock would reduce the local price by 400%, which is not possible. Thus, the condition $r_{12}>\frac{-2T}{a}$ can be regarded as satisfied.

The role of arbitrage can be analyzed by taking the first derivative of $\frac{u(.)}{\bar{p}^z}$with respect to a:

$$\frac{\partial u/\bar{p}^z}{\partial a}=\frac{\overbrace{2(T-1)\frac{ar_{12}^2\bar{p}^2}{T}}^{u' \text{different}}\cdot[\bar{p}+\frac{ar_{12}\bar{p}}{T}]-[(T-1)\frac{a^2r_{12}^2\bar{p}^2}{T}]\cdot\overbrace{\frac{r_{12}\bar{p}}{T}}^{\bar{p}^{z\prime} \text{different}}}{[\bar{p}+\frac{ar_{12}\bar{p}}{T}]^2}>0, \tag{S23}$$

because the numerator can be simplified to:

$$2(T-1)\frac{ar_{12}^2\bar{p}^3}{T}+(T-1)\frac{a^2r_{12}^3\bar{p}^3}{T^2}>0, \text{because } T>1.$$

Thus, a larger number of cities N, which are available for spatial arbitrage, means that $a=\frac{1}{N}$ decreases and reduces volatility.

SA17 Relationship of aggregate rye price and death rate

The *level* of mortality decreases in Germany during the 18th century (Figure S26; data from Pfister and Fertig, 2010). Additionally, the death rate's volatility decreases—once the mortality peak in 1756–1763 is attributed to the Seven Years' War (Figure S26, right axis). The main exception to this pattern is the outstanding subsistence crisis around 1770.

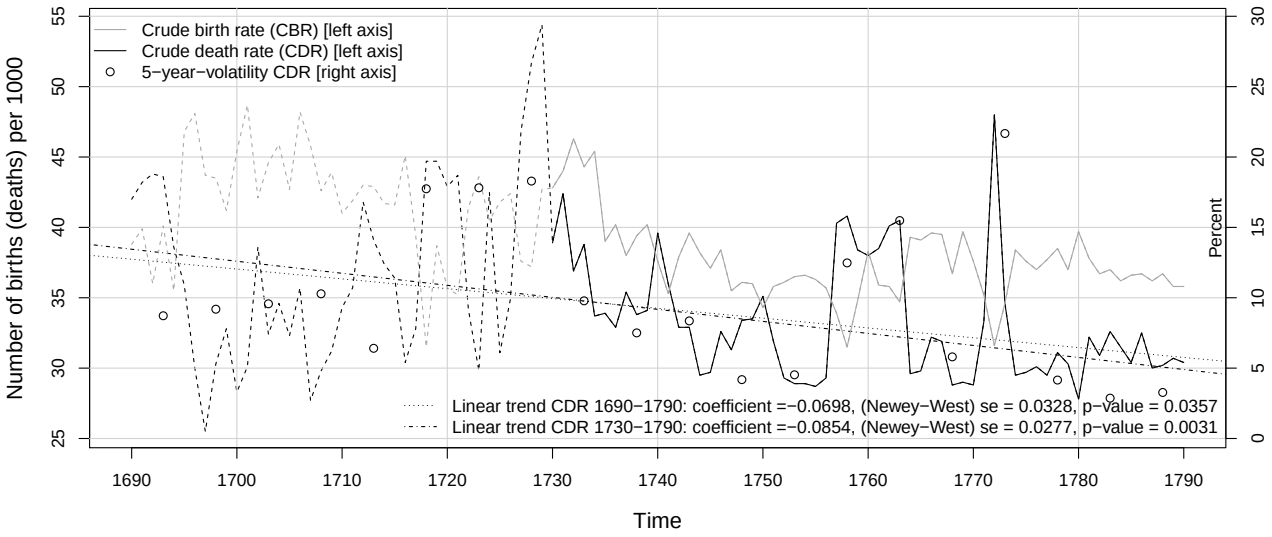

Figure S26: Crude birth and death rates, Germany 1690/1730–1790. Dashed lines refer to periods with unreliable data which are included for a broader picture. Source: data from Pfister and Fertig (2010); similar as their figure 4, p. 31. Regressions for linear trends include a dummy variable for Seven Years' War (1756–1763). Results for trend estimates are similar without dummy. Volatility is plotted as circle for centered year of 5-year period (e.g., 1733 for 1731–35).

The correlation of log-differenced rye price and death rate reported in the main paper is robust to inclusion or exclusion of dummy variables. Dropping the war dummy increases the parameter to 0.38% and reduces $R^2 = 0.14$. Including an additional dummy variable for 1772, where the highest CDR of 48 deaths per 1000 and a very high real price are observed, decreases the relationship to still substantial 0.26% ($p < 0.01$).

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