Empirical Essays on Agriculture in Germany: Past and Present

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vorgelegt von
Hakon ALBERS, M.Sc.

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ERSTGUTACHTER: Prof. Dr. Wolf-Heimo GRIEBEN

ZWEITGUTACHTER: Prof. Dr. Michael KOPSIDIS

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Abstract

Empirical Essays on Agriculture in Germany: Past and Present

by Hakon ALBERS

Since the Neolithic Revolution, humans have produced and consumed cereals. This thesis comprises three essays on the past and present of cereal production and trade. Together with my coauthors I analyze how humans and nature shape these economic activities.

Against the background of global warming and increasing cereal yield variability, the first paper develops an approach to decompose wheat yield volatility into input and weather driven categories. We apply this approach to data for Germany in the period 1995–2009 and show that increasing wheat yield volatility stems in parts from changes in weather variables. However, also input adjustments of farmers and macro-level shocks such as policy changes and the price boom in agricultural commodities in 2008 matter for the surge in yield volatility.

Only few people work in agriculture in an industrialized economy like Germany today. The second paper focuses on the Malthusian period when agriculture was the dominant economic activity in Germany. We ask in how far an integrated market for grain existed in Germany in the pre-railway era, in particular in the period 1650–1790. That period also witnessed climate change, namely the disappearance of the Little Ice Age. We create a new grain price data set, which we analyze to learn about trade in those times based on the law of one price. As weather shocks influence grain price data, measuring market integration is a challenge. The coefficient of variation may be susceptible to shocks and lead to incorrect inferences about the improvements in trade. We find advances in market integration for the foodgrain rye in North-Western Germany and along major rivers; our analysis shows that this result is not driven by weather shocks or climate change in a plausible way. An important consequence of improved spatial arbitrage was the increased stability of aggregate foodgrain prices, a key aspect of food security.

The third paper analyzes the impact of the Napoleonic and Revolutionary Wars (1792–1815) on Germany as a natural experiment. Applying a difference-in-differences framework to rye price data from 1780 to 1830, we find that the territorial expansion of states led to larger internal markets and entailed a significant reduction of price gaps after the wars. These effects are evident for Bavaria and a region that was integrated into Prussia after the wars. Both inland regions had been left untouched by earlier improvements in market integration.

This thesis shows how both nature and humans shape economic activity. Weather and climate affect agricultural production; humans form the institutions governing trade in agricultural commodities. Next to the decisions about trade, humans also develop new agricultural technology to arrest the detrimental effects of nature on the production of food, which remains an essential consumption good also in the modern growth regime. Given global challenges such as climate change and population growth, it follows that the determinants of technological change in agriculture are an important area of future research.

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List of Abbreviations

ADF	**A**ugmented **D**ickey and **F**uller
AIC	**A**kaike **I**nformation **C**riterion
BGP	**B**alanced **G**rowth **P**ath
BMEL	**B**undes**m**inisterium für **E**rnährung und **L**andwirtschaft
BMELF	**B**undes**m**inisterium für **E**rnährung, **L**andwirtschaft und **F**orsten
BMELV	**B**undes**m**inisterium für **E**rnährung, **L**andwirtschaft und **V**erbraucherschutz
CAP	**C**ommon **A**gricultural **P**olicy
CDR	**C**rude **D**eath **R**ate
CES	**C**onstant intertemporal **E**lasticity of **S**ubstitution
CPI	**C**onsumer **P**rice **I**ndex
CRRA	**C**onstant **R**elative **R**isk **A**version
CRS	**C**onstant **R**eturns to **S**cale
CV	**C**oefficient of **V**ariation
DWP	**D**ays **w**ithout **P**recipitation
DD	**D**ifference in **D**ifferences
EU	**E**uropean **U**nion
ETP	Potential **E**vapo**T**rans**P**iration
FADN	**F**arm **A**ccountancy **D**ata **N**etwork
FAO	**F**ood and **A**griculture **O**rganization
FAOSTAT	**F**ood and **A**griculture **O**rganization **Stat**istical Database
FD	**F**irst **D**ifferences/d
FE	**F**ixed **E**ffects
GDD	**G**rowing **D**egree **D**ays
KDD	**K**illing **D**egree **D**ays
KLEMS	**K**apital, **L**abor, **E**nergy, **M**aterial inputs and **S**ervices
KPSS	**K**wiatkowski **P**hillips **S**chmidt **S**hin
IPCC	**I**ntergovernmental **P**anel on **C**limate **C**hange
IV	**I**nstrumental **V**ariable
LFA	**L**ess **F**avoured **A**rea
LIA	**L**ittle **I**ce **A**ge
LOP	**L**aw of **O**ne **P**rice
MA	**M**oving **A**verage
MC	**M**arginal **C**ost
MRS	**M**arginal **R**ate of **S**ubstitution

NW	**N**orth-**W**est
PFA	**P**roduction **F**unction **A**pproach
Prec	**Prec**ipitation
RESET	**Re**gression **S**pecification **E**rror **T**est
RMSE	**R**oot **M**ean **S**quared **E**rror
SA	**S**upplementary **A**ppendix
SCC	**S**patial **C**orrelation **C**onsistent
SD	**S**tandard **D**eviation
SE	**S**tandard **E**rror
SRT	**S**olar **R**adiation, **T**emperature normalized
SSD	**S**um of **S**quared **D**eviations
Temp	**Temp**erature
TC	**T**otal **C**ost
TFP	**T**otal **F**actor **P**roductivity
V	**V**olatility
VAT	**V**alue **A**dded **T**ax
VAR	**V**ector **A**uto**r**egressive
VPD	**V**apor **P**ressure **D**eficit
WMO	**W**orld **M**eteorological **O**rganization

For Lisa

Chapter 1

Introduction

Agriculture was the first step towards a more complex economy with different activities. Only since humankind has been able to generate a surplus of food, a more elaborate division of labor than in hunter-gatherer societies is possible (Weisdorf, 2005). The transition to agriculture is therefore considered as revolutionary – the Neolithic Revolution (ibid.). For most of history, agriculture was the most significant economic activity measured in its share of the labor force and its contribution to output (Mellor, 2018; Weisdorf, 2005). Although agriculture has lost its importance in aggregate economic activity in most industrialized countries by now, it still provides humans with most of the food, an essential consumption good for every human being.[1]

This thesis comprises three empirical essays on agriculture in Germany. Together with my coauthors I analyze how people and nature shape agriculture in present days and shaped it in the past. A major output of agriculture since the Neolithic Revolution are cereals. This is one of the few goods that have been produced by people over such a long time span in so many different growth regimes.[2] Here the literature distinguishes the Malthusian growth regime, the post-Malthusian and the modern growth regime in which we live today (Galor and Weil, 2000). The study of grain yields and prices allows an intertemporal comparison to understand the relative importance of nature and humans in economic activity and agriculture in particular.

In what follows, I introduce the three essays: I show how they are connected and provide additional background information. Section 1.1 discusses the paper on current wheat yield variability in Germany. Paper 2 opens the intertemporal comparative perspective and analyzes the German grain market more than 300 years ago (Section 1.2). Next to the focus on grain that is common to all three studies, the first two papers are related through the theme weather shocks and climate change.[3] The third paper focuses on an exclusively man-made shock to the German economy: the impact of the Revolutionary and Napoleonic Wars on the grain market in Germany (Section 1.3). Paper 3 follows paper 2 sequentially in time and is based on data for the period 1780–1830.

[1] Fishery is an important alternative provider of food, in particular of animal protein (FAO, 2018, 2–3).

[2] The other goods I am aware of are animal based food products, clothes and basic alcoholic drinks such as beer (Pfister, 2017, supplementary appendix S2, p. 1).

[3] In paper 1 and 2, we refer to weather as the inter-annual variation of meteorological variables such as temperature (Dell et al., 2014). In paper 2, we refer to climate as the long-run temporal average 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)."

Paper 3 is also related to paper 1 through the factor institutions.[4] At the turn to the 19th century, the Napoleonic wars impacted on the institutions governing grain trade; these days agricultural policy provides the institutional framework that influences cereal production and trade. Finally, Section 1.4 provides a summary of this introduction and explains the outline of the thesis.

1.1 Paper 1: How do inputs and weather drive wheat yield volatility? The example of Germany

Paper 1 (coauthored by Silke Hüttel and Christoph Gornott) asks how inputs and weather drive wheat yield variability in current Germany.[5] Against the background of global warming as reported by the Intergovernmental Panel on Climate Change (IPCC), increasing cereal yield fluctuations are of concern.[6] First, farm incomes might become more variable and risky. Second, Germany is a major wheat producer and wheat an important provider of calories. Thus, to sustain food security it is important to understand how weather drives aggregate yield variability. While the study cannot relate global warming and current wheat yield fluctuations due to the limited period we study, we show that a range of weather variables impact on aggregate wheat yields. I now review the term food security and then discuss how wheat yield volatility can be driven by humans.

Food security is a concept of the Food and Agriculture Organization (FAO) of the United Nations. "Food security exists when all people, at all times, have physical and economic access to sufficient, safe and nutritious food that meets their dietary needs and food preferences for an active and healthy life (World Food Summit, 1996) cited by FAO (2006, 1)." Food security has four dimensions: food availability, food access, utilization, and stability (FAO, 2006, 1). The most relevant dimension for this thesis is stability: "To be food secure, a population, household or individual must have access to adequate food at all times. They should not risk losing access to food as a consequence of sudden shocks (e.g. an economic or climatic crisis) or cyclical events (e.g. seasonal food insecurity). The concept of stability can therefore refer to both the availability and access dimensions of food security (FAO, 2006, 1)." The differentiation between availability and

[4]Institutions are "[...] the formal and informal rules shaping social, economic and political behaviour [...] (Alston, 2018, 9484)." See also Kopsidis and Bromley (2016, 163–4).

[5]From an agronomic perspective, yields are the net outcome of different yield components (crop density, kernels per ear, kernel weight), which may be influenced by weather in different ways (Chmielewski and Köhn, 2000). Economically, yields are a partial productivity measure, that is, output per unit of the production factor land (Coelli et al., 2005, 3). "Total factor productivity (TFP) is the portion of output not explained by the amount of inputs used in production (Comin, 2018, 13720)."

[6]According to the IPCC report, the global warming trend from 1880 to 2012 is 0.85°C (IPCC, 2014, 2–3). While the focus of paper 1 is how nature (weather) impacts on the cereal production of humans, one could argue against this background that humans also impact on nature, more specifically on climate via greenhouse gas emissions. Following the IPCC report, humans have contributed to global warming by emitting greenhouse gases such as CO_2 with a likelihood of 95–100% (IPCC, 2014, p. 2, note 1, p. 4). The increase of greenhouse gases like CO_2 leads to warming, because they absorb particular wavelengths (infrared) of the energy radiated by the earth that would otherwise not be absorbed (NASA, 2009; Schönwiese, 2013, ch. 4.2, particularly pp. 119–21; ch. 12.3, particularly p. 353).

access goes back to the work of Sen (1981) (FAO, 2006, 1). The key point of Sen's analysis is that people face hunger, because they do not have enough food (because they are not entitled[7] to get access to food) while there may be indeed enough food available (Sen, 1981, 1). The concept of food security is admittedly broad (Wheeler and von Braun, 2013, 509). Pinstrup-Andersen (2009, 6–7) discusses problems in quantifying food security, e.g., of whether only calories matter or also micronutrients, and the sensitivity of the estimates to the used data source (consumption surveys or income data and food prices). His focus is mostly on the availability, access and usage dimension, not on the stability dimension. Based on the FAO definition of food security, the IPCC report devotes a complete chapter to food security (Porter et al., 2014).

In addition to weather, also humans are a potential factor that might explain changing patterns of yield variability for two reasons. First, we expect rational, that is, profit maximizing farmers to adapt to output and input price variations and adjust variable inputs such as fertilizer. Second, agriculture has been a heavily regulated sector since the foundation of European Union (EU) but has seen major deregulations of the EU's Common Agricultural Policy (CAP) since the 1990s (see SA2.2.6 for details). Inputs adjustment and changes of the CAP can affect the yield levels that producers realize and accordingly also yield variability.

We compile an unusually rich data set from different sources. We include wheat yields, major inputs and weather variables aggregated according to important phenological stages.[8] That is, our empirical framework acknowledges the agronomic insights that the development of plants does not follow the human calendar and that weather variables have potentially different impacts in different phenological stages.

In a first step we estimate a panel data model based on a production function and in a second step, we isolate weather driven fitted values from the regression model. These fitted values are used to calculate weather-induced yield volatility at the federal-state level. The method of this second step is developed within the paper and builds on the work by Osborne and Wheeler (2013).

The main result of this paper is that both inputs and weather matter in explaining the observed increase in wheat yield volatility. Furthermore, our work reveals the spatial and temporal heterogeneity in yield volatility: We can trace back increasing yield volatility to weather variables in only some states.[9]

The high aggregation level in our study indicates that the weather induced wheat yield volatility we derive is a measure of weather induced systemic risk in wheat production at the federal state level. Systemic risk, in this context the spatially correlated risk in agricultural production, is a problem for crop yield insurance solutions, simply because it would lead to large indemnity

[7]Entitlements are ownership relations (Sen, 1981, 1–2).

[8]Phenological stages are growth stages of plants (Meier, 2001).

[9]All empirical work in this paper and the rest of the thesis has been carried out using the open-source statistical software R (R Development Core Team, 2016). To allow replication of this research, code and data for the published paper (Albers et al., 2017) are available on the publisher's website. On the role of replication in empirical economics see Dewald et al. (1986), Hamermesh (2007), and Koenker and Zeileis (2009).

payments to many producers at the same time (see, e.g., Xu et al., 2010).[10] While systemic risk is a problem at the regional and potentially also at the national level for the insurance of weather risk in Germany today, in the past, the spatial correlation of crop losses had comparably more severe consequences than shortfalls in income. According to Ehmer (2005, 902, 907) hunger crises in pre-modern societies were linked to mortality, although mostly not directly but indirectly through a higher susceptibility of individuals to disease.[11]

1.2 Paper 2: Market integration or climate change? Germany, 1650–1790

Paper 2 (coauthored with Ulrich Pfister and Martin Uebele) zooms back in time to the pre-industrial and Malthusian era. That is, we analyze Germany in the period from roughly 1650 to 1790 in a completely different macroeconomic setting. In this period, agriculture was the dominant economic activity in Germany. For example, Pfister (2011, 5) estimates that approximately 79% of workers were working in agriculture in 1650 and still 64% in 1800.[12] By contrast, only 1.4% of the labor force worked in Germany's agriculture in 2018 (own calculation based on data from Statistisches Bundesamt, 2019b).[13] With regard to consumption, a much larger share of income was spent on food and drink items in pre-industrial times (82.6%) with bread from grain being the single most important item (34%) (Pfister, 2017, S2, p. 1). By contrast, the German federal statistical office includes food and drink items these days with a share of 9.7% in the consumer basket (fixed weight of 2015); bread – even together with other cereal based items – has a share of only 1.5% (Statistisches Bundesamt, 2019a). While the figures are not directly comparable,[14] the tendency towards a much lower income share for food and grain today is obvious.

This empirical pattern is a realization of Engel's law. "[...] Engel's law [...] states that as a household's income increases, the fraction that it spends on food (agricultural products) declines (Acemoglu, 2009, 698)." Engel's law can generate important macroeconomic effects in an economy such as structural transformation, that is, the outflow of labor from agriculture to other sectors (Kongsamut et al., 2001), or a positive effect of agricultural productivity on the growth rate of a

[10] The advantage of our approach is that we account for weather, input and macro-level shocks resulting from policy at the same time. However, note that systemic risk can be analyzed using more flexible statistical tools such as copulas. "Using a copula model allows the selection of marginal distributions separately from the modeling of their interdependence (Gaupp et al., 2017, 2213–4)." Based on a copula model, these authors show that wheat yields of important wheat producers are independent at the global level.

[11] A large literature deals with (the economics of) famines in different historical and economic settings, e.g., Ó Gráda (2007), Ravallion (1997), and Sen (1981).

[12] These estimates are based on information on urban population and non-agricultural rural workers (Pfister, 2011, 5). This period is part of the proto-statistical era, that is, statistical offices did not exist.

[13] The share of agriculture in GDP also declined over time (this is a general pattern in the development process, see Mellor, 2018, 216). To my knowledge, the earliest available estimate is available for ca. 1850. In 1850/54, the share of agriculture in output was 45.2% (Pfister, 2011, 8, based on Hoffmann, 1965). The share of agricultural labor in the total labor force was 55.6% in 1849/52 (ibid.). The productivity differential of agriculture to non-agriculture which Pfister discusses is featured in recent cross-country data as well, see Gollin et al. (2014).

[14] Note that expenditures on housing are not included in the consumer basket by Pfister, because these data are not available at all.

closed economy (Matsuyama, 1992). Engel's law is often incorporated in theoretical models via non-homothetic preferences, that is, the share of income spent on a particular good depends on the income level. Appendix A discusses non-homothetic preferences in more detail.

The paper's aim is to understand to what extent the German lands comprised a common market. To this end, we create a new grain price data set from the 15th to the 19th century with a focus on the period 1650–1790.[15] For the pre-industrial era, price data are to date the only viable way to learn about trade, because systematically recorded trade flows (e.g., Wolf, 2009) are not available.[16]

We collect as many grain prices as possible while keeping a common standard of data quality. The conversion of these price data to a common unit is far from trivial, because there were neither a common currency nor common norms for volumes in the German lands. We discuss all necessary steps and the background on the monetary regime that prevailed in pre-industrial Germany in SA3.1. In brief, there existed only minted coins, no paper money, and we use the final metal content to convert prices to grams of silver per litre.[17] A further challenge is to ensure a comparable time base for all price data, because prices are sometimes available for a particular time in the year (so-called *Martini* prices) or are averaged not for the calendar year but for so-called crop years. Figure 1.1 summarizes the workflow of data set construction and statistical analyses. We discuss further details in the paper and the corresponding SA.

[15] While it would be interesting to study yields as well, we do not know much about cereal yields in pre-industrial Germany. Exceptions include Nuremberg on the basis of tithe (1339—1670) from Bauernfeind and Woitek (1999, 465) and Saxony, 1792–1830 (Pfister and Kopsidis, 2015; Kopsidis et al., 2014). Slicher van Bath (1963b) provides yield ratios for some German estates in the pre-industrial period; these few micro data are not representative at the aggregate level (cf. Waldinger, 2015).

[16] In competitive markets, the product price must equal the marginal costs (MC) of production (Simon and Blume, 1994, 62–4). The MC function can be derived as follows. First, one derives the *conditional* factor demand functions. The conditional factor demand is a function of output and the prices of the production factors (Simon and Blume, 1994, 562). This involves using the *inverse* factor demand functions. The latter are the first-order conditions from the profit maximization problem of the firm (Simon and Blume, 1994, 558). Second, the conditional factor demand functions are used to uncover the (total) cost (TC) function (Simon and Blume, 1994, 562–3). The latter's first derivative with regard to output is the MC function (Simon and Blume, 1994, 59).

[17] Note that some nominal prices are given in so-called money of account, a term that begs further explanation. Money of account (*Rechengeld*) results from decoupled basic functions of money in pre-modern currency systems: (i) unit of account and (ii) store of value/means of payment (Metz, 1990, 3, 25). E.g., people would count in *Rechengulden* (which did not necessarily had to have an equivalent minted coin) and which was set in a fixed relation to a smaller unit such as *Rechenalbus* (1 *Rechengulden* = 24 *Rechenalbus* (Metz, 1990, 44–5). Following Metz, the latter co-existed as a minted coin (in this case called *Albus* or *Weißpfennig*) but its silver content was variable over time (usually debasement). Thus, to assess the value of *Rechengulden* or *Rechenalbus*, either the actual silver content of *Albus* needs to be known (*Albus* as so-called link money) or exchange rates to other minted coins (usually large gold coins such as *Goldgulden*) are needed (Metz, 1990, 27, 41–5).

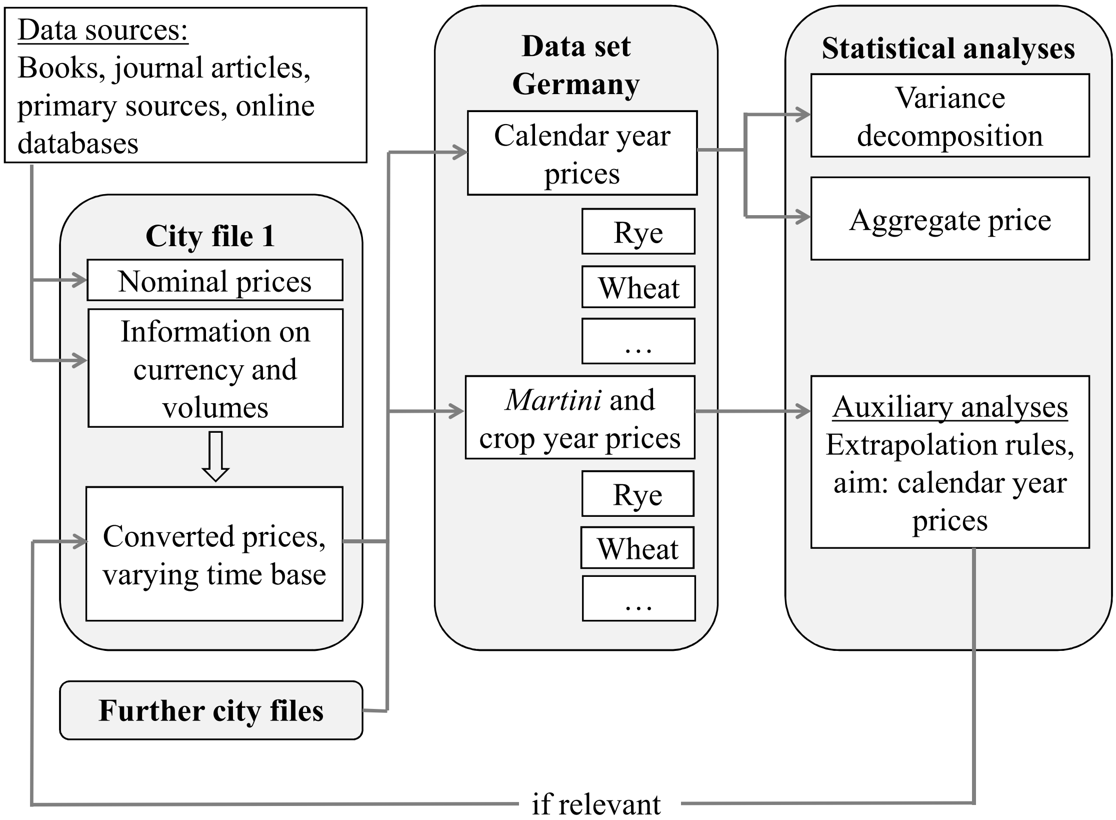

FIGURE 1.1: Workflow of data set construction and analysis. *Note*: The city files and the main data set are constructed using Microsoft Excel; data are then exported from the main data set for statistical analyses with R.

The average of all rye prices in the individual cities for a given year, that is, the aggregate rye price, can be regarded as a summary of this data set and opens a new window to the historical German economy (see, Figure S3.8 in the supplementary appendix SA3.6.4). The aggregate price signals the relevance of grain in pre-industrial times. Major peaks in the rye price are associated with subsistence crises reported in the historiography.[18] In addition, the aggregate price series is statistically significantly related to the aggregate German death rate. We provide a detailed discussion of subsistence crises in Germany and the statistical relationship to the death rate in the supplementary appendix of paper 2 (SA3.6.4 and SA3.6.5).

Market integration is relevant for two reasons. First, it allows to exploit comparative advantages and leads to a more efficient use of production factors (e.g., Federico, 2012). Even without inducing technological progress, trade between cities leads to regional specialization according to

[18]We use the term subsistence crisis when rye price peaks occur together with increasing mortality and decreasing fertility; in this sense, subsistence crises (or synonymously hunger crises) are a particular type of demographic crises (Ehmer, 2005, 900–6, see also Gailus, 2007, 712). Other reasons for demographic crises were epidemics and war (Ehmer, 2005, 901).

comparative advantage and increases TFP at the aggregate level (for the theoretical background consider a Ricardian trade model, e.g., van Marrewijk, 2012, ch. 3).[19]

Second, grain price spikes had the potential to trigger hunger crises and—in the worst case—could lead to death. Additionally, malnutrition could harm cognitive ability and reduce the productivity of humans (Baten et al., 2014). One would expect integrated markets to have a lower temporal price volatility so that price fluctuations are smoothed out in a better way (e.g., Chilosi et al., 2013, 48).

While these days often global warming motivates research on agriculture, climate changes have appeared throughout history.[20] The Little Ice Age (LIA) was a period of cooler climate in the Northern Hemisphere (Masson-Delmotte et al., 2013, 389).[21] The LIA also included the geographical area of the German lands. We critically discuss in how far one important method to measure market integration as price convergence, the coefficient of variation (CV), might signal market integration although possibly no arbitrage took place. Warming associated with the end of the LIA might have lead to increasing agricultural TFP and impacted on yields and price levels. The latter are used by the CV to detect market integration. We show that it is theoretically possible that prices of different sub-regions converged although these regions do not trade with each other. Furthermore, we show that the CV is sensitive to weather shocks in a way that might drive results towards price convergence.

When discussing the CV in detail we also support our analysis with regressions of the aggregate rye price on reconstructed weather variables. Here we transfer basic insights from the literature on yield and weather, which forms the background for paper 1, to a historical setting.

In addition to distorting the measure of spatial arbitrage, the end of the LIA might have led to less shocks to agricultural output. Less shocks might have entailed a lower temporal price volatility – without advances in market integration.

To exclude that climate change or weather shocks drive our findings in a significant way towards market integration, we use the SD of five-year-average prices, a choice that we support with a formal analysis of the CV. Additionally, our study is the first to provide results on the food crop rye. Other research has mostly focused on wheat, which was far less important in production

[19] Appendix B.1 discusses why TFP likely increased in pre-industrial Germany. The referenced Ricardian trade model assumes two goods. Examples of the relevant goods in our case could be grain and linen (for clothes). The latter is part of the consumer basket by Pfister (2017, S2, p. 1) (cf. Allen, 2001, 418, 421). Note that changes in agricultural technology were very limited in 1650–1790 so that trade is a particularly relevant channel for the expansion of output. I discuss the other potential channel, technical change in pre-industrial German agriculture, in Appendix B.2.

[20] Schönwiese (2013, ch. 11) provides an overview about climate history.

[21] On the LIA see also Glaser (2007). Kelly and Ó Gráda (2014a) challenge the established view arguing that the LIA is a statistical artifact from smoothing temperature data. We discuss this point explicitly in more detail in paper 2. Furthermore, I provide technical background information on the moving average (MA) filter in appendix C of this thesis. In short, we agree with the statistical point on the MA-filter but one can show significant changes that are consistent with the definition of climate change above. Thus, the critique by Kelly and Ó Gráda (2014a) is not sufficient to show that the LIA did not exist. Further articles on the debate are: Kelly and Ó Gráda (2013) introduce their point and also include historiographical evidence on the LIA. The climatologists' view is expressed by Büntgen and Hellmann (2013), who received a reply by Kelly and Ó Gráda (2014b).

and retail trade in Germany (details are discussed within Ch. 3.1 on p. 55). In addition, the by historical standards high spatial resolution of the data set allows to track down in more detail than earlier research where market integration appeared.

We find that rye prices converged by about 0.4% per year so that price dispersion was reduced by more than 50% during the 140 years between the Thirty Years' War and the Napoleonic Wars. Convergence appeared mostly in North-Western Germany and along major rivers.[22] We can exclude that these results are driven by weather shocks. Furthermore, the result for Northwestern Germany cannot be driven by climate change, that is, the end of the LIA (given the assumptions which are detailed in paper 2). In addition, aggregate rye price volatility decreased significantly by ca. 1 percentage point in ten years. The paper also shows formally that these two results are directly related: if spatial arbitrage increases, aggregate price volatility must decrease. However, changing shock patterns likely played a quantitatively important role in the moderation of rye price volatility as discussed in more detail within the paper.

1.3 Paper 3: War, state growth and market integration in the German lands, 1780–1830

Paper 3 (coauthored with Ulrich Pfister) builds on paper 2. We further extend the grain price data set and exploit the Revolutionary and Napoleonic Wars (1792–1815) as a natural experiment to test in how far larger states—an outcome of the wars—led to decreasing price gaps.[23] While we borrow the idea that the wars were a shock to the German economy from Acemoglu et al. (2011), our study is more closely related to the one by Keller and Shiue (2016). These authors argue that a major impact of the wars on the German economy worked through market integration that was fostered by French reforms.

One difference to their work is that we only focus on the outcome variable 'market integration' and do not estimate any further impact on output growth. More importantly, we propose a completely different channel—state growth—and do not rely on the index of French inspired reforms that Kopsidis and Bromley (2016) criticize.

Our new data set allows a different empirical strategy than the instrumental variable approach (see Wooldridge, 2016, 462–3) as in Keller and Shiue (2016). Given that we have data before the wars, we can apply the difference-in-differences (DD) estimator (Wooldridge, 2016, 410–1) to quantify the effect of state growth on market integration.[24]

[22] Waldinger (2015, 18–9) shows that the relationship of long-run temperature averages (100 or 50 years, see Waldinger, 2015, 6) and city size (which presumably works through agricultural productivity) in Europe 1600–1750 was less pronounced when cities had access to waterways. This finding indicates that trade was important in attenuating variations in climate and associated local agricultural productivity.

[23] "A natural experiment occurs when some exogenous event [...] changes the environment in which individuals, families, firms, or cities operate (Wooldridge, 2016, 410)." The event creates the control group and the treatment group, which are compared by the researcher (ibid.).

[24] Angrist and Pischke (2009, ch. 5) provide a more detailed introduction to the DD estimator.

While paper 1 and paper 3 both rely on regression analysis, in their detail they rely on different empirical strategies. Paper 1 mainly excludes endogeneity concerns arising from potentially omitted variables by exploiting a data set that covers all important variable categories. Paper 1 also uses an (adapted version of the) approach known as from general to simple modeling (Greene, 2012, 178–80). According to this approach, a larger model, which contains smaller sub-models, is specified first and than narrowed down to a final model. In contrast, paper 3 specifies a relatively simple baseline model and then criticizes that model repeatedly by adding or dropping variables and varying the sample.

Compared to paper 2, which relies on a descriptive statistical tool (the SD) in combination with trend tests, the empirical approach in paper 3, the DD estimate, allows a causal interpretation as 'average treatment effect' (Wooldridge, 2016, 410).[25]

The key result of paper 3 is that price gaps within two regions decreased after the wars compared to the control group (and the pre-war level).[26] These regions are Bavaria and a region that was part of the short-lived 'Kingdom of Westphalen' and was later integrated into Prussia.

1.4 Summary and outline of the thesis

In short, this thesis comprises three empirical analyses which focus on the agricultural sector in Germany in different macroeconomic settings. Paper 1 is an empirical investigation into the drivers of wheat yield volatility in Germany 1995–2009. While the agricultural sector has lost its macroeconomic importance measured by its labor share, wheat is one of the most important food crops worldwide and Germany an important wheat producer in the EU. As food cannot be substituted away, the economics of its production and trade remain relevant to sustain food security, particularly against the background of global warming which will likely impact on cereal crops. The paper demonstrates that current wheat yields and their volatility are driven by weather shocks and input variations. Increasing yield volatility can be traced back to weather variability in only some federal states. The paper also shows that macro-level shocks associated with agricultural policy matter.

In Malthusian Germany 1650–1790, by contrast, rye was the most important food crop and most Germans worked in agriculture. Foodgrain trade was particularly important, because it could attenuate severe consequences of local crop failure, which in those times included death by famine. Paper 2 shows that measuring market integration based on price data is challenging given the background of weather shocks and climate change but that a market for grain in fact developed in the Northwest and along major rivers. One consequence was a moderation of rye price volatility. However, this empirical pattern is likely driven also by the changing shock patterns associated with the end of the LIA.

[25] A brief introduction to causality in econometrics is given by Wooldridge (2016, 10–4).

[26] This result is visible in Figure 4.2, which constitutes a summary of the paper's empirical strategy and key result. In Figure 4.2, the large peaks in the average price gap of 'Prussia not newly included' in 1795 and 1816 are also visible in mortality data (Pfister and Fertig, 2010, 33; see also Collet and Krämer, 2017, 109–10 on the Tambora crisis in 1816, named after the eruption of the Indonesian volcano in 1815).

The main result of paper 3 is that larger states with larger internal markets that were an outcome of the Napoleonic Wars—a man-made shock to the German economy—enjoyed substantially reduced rye price gaps. This historical shock was large compared to the regime shifts in agricultural policy, which have occurred since the 1990s and which are reflected in the empirical analysis of paper 1.

The rest of this thesis is structured as follows. The main part consists of the three papers (Chapters 3–4). Chapter 2 studies wheat yield volatility and Chapter 3 focuses on market integration in pre-industrial Germany. Chapter 4 is devoted to the impact of state growth on grain market integration. Each paper is presented together with its supplementary appendix (SA), which contains additional results and supporting information. Finally, Chapter 5 develops the overall conclusion. Additional appendices at the end of the thesis provide further background information that I refer to in this introduction and the conclusion.

Chapter 2

How do inputs and weather drive wheat yield volatility? The example of Germany

Authors: Hakon ALBERS*, Christoph GORNOTT† and Silke HÜTTEL‡

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Abstract: Increases in cereals' production risk are commonly related to increases in weather risk. We analyze weather-induced changes in wheat yield volatility as a systemic weather risk in Germany. We disentangle, however, the relative impacts of inputs and weather on regional yield volatility. For this purpose we augment a production function with phenologically aggregated weather variables. Increasing volatility can be traced back to weather changes only in some regions. On average, inputs explain 49% of the total actual wheat yield volatility, while weather explains 43%. Models with only weather variables deliver biased but reasonable approximations for climate impact research.

Keywords: Yield, wheat, variability, risk, weather, Common Agricultural Policy

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* **Corresponding Author**. Martin Luther University Halle-Wittenberg, Department of Economics, Universitätsring 3, D-06108 Halle, Germany

† Potsdam Institute for Climate Impact Research (PIK), Germany.

‡ Rheinische Friedrich-Wilhelms-Universität Bonn, Institute of Food and Resource Economics, Production Economics Group. Former address: University of Rostock, Agricultural Economics, Germany.

2.1 Introduction

Climate change and its consequences for agricultural production have been open to environmental, social and economic debate for years. This is not surprising since weather conditions considerably determine crop yield levels and their variability, which are of interest for food security reasons at the macro-level (Brown et al., 2015; Wheeler and von Braun, 2013). Yields are also interesting at the micro-level, where a low level of yearly crop yield variability reduces income risks and contributes to farm income stability, which in turn could be relevant at the macro-level in that it warrants resilient food production. Hence, it is vital to better understand what determines yield variability in the most important crop-producing regions. This may also help farmers adapt their agronomic strategy towards better-known risks, and help policy makers to prevent food-crises or improve crisis management.

Undisputedly, long-term climatic changes alter cropping conditions (Siebert and Ewert, 2012) and might already have affected crop yield variability, which is identified as a key production risk of the most economically important cereals (IPCC, 2014, 71). Extreme weather events like the European heat wave in 2003 were discussed as either indicating an increase in temperature variability or resulting from a shift of the temperature distribution (Luterbacher et al., 2004; Perkins, 2015; Schär et al., 2004). Consensus exists that in the future, extreme weather events are expected to occur with greater frequency and severity in both temperate and tropical regions (IPCC, 2014, 69–73). This will likely make crop production more vulnerable, with potentially considerable impacts on farm incomes and food security, particularly in less developed regions.

Farmers can control inputs like fertilizer for a given natural production environment like soil quality but cannot control the weather, nor can they affect developments in markets, agricultural, or environmental policy. Weather[1] is exogenous to farmers and directly affects crop yields. Additionally, indirect effects entailing input adjustments exist. For instance, weed growth, pests and diseases vary depending on weather conditions and farmers usually adjust their inputs accordingly during the production period. Weather can be interpreted as the major driver of production risk in crop production, though the question remains, how much overall production risk can actually be traced back to changing weather conditions?

In this study we consider wheat—one of the most important cash crops worldwide—where considerable upward trends in both yield levels and variability have been observed. While in 1995/96, on average, about 2.5 metric tons per hectare (tons ha^{-1}) were harvested worldwide, in 2012/13 this increased to about 3.2 tons ha^{-1} (FAOSTAT, 2015). Our investigation concentrates on Germany, which produces 17% of the European Union's (EU) wheat output. In the period 1995/96 to 2012/13, German wheat yields increased from 7.1 to 7.7 tons ha^{-1}. Although a long period of relative yield stability existed in the 20th century (Calderini and Slafer, 1998), both absolute and relative wheat yield variability have increased in Germany since the 1990s (Krause, 2008; Osborne

[1]We use the term "weather" to be consistent with the majority of papers we reviewed. The literature applies different definitions. Dell et al. (2014) refer to inter-annual weather variations as long as the aggregation period is less than one year. Another strand of literature favors using a year-to-year or inter-annual variation of "climate" (e.g., Ray et al., 2015).

and Wheeler, 2013). Particularly concerning is the upward trend in relative yield variability, that is, an increased proportion of yield at risk relative to the expected mean.

Against this background, the research questions guiding our analysis are as follows: How to explain increasing relative yield variability? Particularly, can one really conjecture that production risk measured as relative yield variability has increased only through changes in weather conditions, as the climate change discussion implies?

Several other reasons for this increase exist. First, farmers might adjust input levels because of changing input and output prices (Miao et al., 2016), while Finger (2010) discussed the importance of agricultural policy for yield analyses. Farmers in the EU have been exposed to rather radical changes in the Common Agricultural Policy (CAP) since 1992. Several reforms elevated the relative competitiveness of wheat, for instance, by removing price support, subsidies and compulsory set-asides (e.g., Gohin, 2006). Additionally, renewable energy policies have been proven to favor maize for silage (in Germany, increases of about 21% in the years 1990–2009 were reported, Statistisches Bundesamt, 2015). This might also have contributed to changes in the relative competitiveness of wheat production, which has consequences for input intensity and thus crop yield levels (Banse et al., 2008; Schulze Steinmann and Holm-Müller, 2010). Overall, these policy changes may have provided incentives for farmers to use lower quality (marginal) land for wheat production, likely with negative effects on average yield levels and increased variability. Crops planted on marginal soils with low water-holding capacity might be more sensitive to extreme temperature and precipitation changes compared to more favorable soils (Perkins, 2015, 248–249). Moreover, yield can be interpreted as land productivity and may have increased due to scale and specialization effects (e.g., Yang et al., 1992, Kaufmann and Snell, 1997). Ongoing consolidation processes in the EU's agricultural sector (i.e., increased farm sizes) might enhance average yields per hectare despite the growing trend of planting marginal land with wheat.

While numerous studies consider how weather interacts with crop yield levels and their variance based on regression models (e.g., Chen et al., 2004), the relation between weather and relative yield variability of non-experimental yields has been analyzed by few researchers, for instance, Lobell (2007) or Ray et al. (2015). These authors, however, do not acknowledge any input adjustments that influence yield stability. To the best of our knowledge, thus far, the sources of yield volatility have not been disentangled into the major drivers of weather and inputs. Within this study we aim to close this gap and illustrate this idea using a case study for wheat yields in Germany.

While Iglesias and Quiroga (2007) assess the impact of weather variables on crop yields using time series regressions, we apply a panel data approach. We exploit the advantages of the panel structure to quantify whether and how weather- and input-induced risk has changed overall or only in some parts of Germany over time. Within our approach, we augment the contribution from Osborne and Wheeler (2013) and show that both inputs and weather matter for explaining yields and their relative variability. Our research contributes to the discussion of whether inputs need to be modeled when assessing climate change impacts on cereal yields. Further, understanding how weather drives observed relative yield variability today might be helpful for future adaptation challenges.

Our empirical analysis involves two major steps. First, we develop an empirical model of relative yield variability consistent with a production function approach. We consider major inputs, test for suitable functional forms and enhance this production function by a rich set of weather variables addressing phenological development. Second, we decompose the fitted values of this regression model to disentangle weather-induced compared to input- or policy-induced relative yield variability referring to the approach by You et al. (2009). To improve our understanding of whether to control for input adjustments while relating weather and yields, we present an alternative model that leaves out major inputs. Hypothesizing that the latter may suffer from omitted variables bias, our results show no considerable qualitative differences, though they do exhibit quantitative differences.

In what follows, we first unfold the conceptual framework and present related literature. After introducing the data, the presented framework leads us to our empirical strategy for disentangling crop yield volatility drivers. Following that, we report and discuss our results, and finally conclude.

2.2 Conceptual framework and related literature

Numerous studies deal with the impact of weather on yield levels by using either process-based crop simulation models (Müller and Robertson, 2014) or regression techniques.[2] The latter approach finds its roots in Oury (1965) and has two major strands. First, many studies exist that simply relate yield and weather within a regression model (e.g., Butler and Huybers, 2015; we refer here to the literature overview Tables S2.3–S2.5 in the supplementary appendix [SA]). In the second strand, weather impacts are analyzed within a production function framework including inputs. These models treat weather exogenously; however, a need to adjust inputs to changing weather might exist. For instance, the precipitation level will affect fertilizer intensity. Temperature instead affects length of the growing season and as such contributes to yield levels but rarely induces short-run adjustments to the input mix. While the first group of models takes this tacitly as a motive for leaving out inputs, the second strand of literature can also be criticized. While accounting for adjustments in the input mix in the short-run, production functions often fail to capture long-term adaptations to changes in climate such as altering crop rotation or alternative land-uses (e.g., Mendelsohn et al., 1994 or Deschênes and Greenstone, 2007).

When hypothesizing yield to be a function of inputs and weather, neglecting one group in the estimation of the impact of the other could result in biased parameter estimates as discussed by Kaufmann and Snell (1997), Reidsma et al. (2007, 417) or more recently by Miao et al. (2016, 201). In light of this debate, rather surprisingly only few recent studies include inputs or acknowledge other economic variables while analyzing weather impacts on yields (e.g., among others Schlenker and Lobell, 2010, Lobell et al., 2011, Blanc, 2012 or Ward et al., 2014). In this context, scale effects

[2]Literature reviews can be found in Dell et al. (2014), Schlenker and Roberts (2009), Tannura et al. (2008) and Ward et al. (2014).

with regard to land have also been shown to influence yield levels (e.g., Chen et al., 2004). Hence, we rely on a production function approach including major inputs.

Disentangling the impacts that weather and inputs have on crop yield levels and their volatility, however, remains a challenge (You et al., 2009, 1013). Technically, a variety of approaches exist that quantify weather effects in a production function framework. We identify three crucial choices: the *selection* of weather variables, *aggregation* levels of weather data and the *functional form* describing the input-output and weather-yield relationships (further discussion of these choices in the SA).

Using aggregated data allows us to isolate the systemic component of weather risk at the federal state level simply because idiosyncratic shocks evident at the farm level are "averaged out" at higher aggregation levels (Marra and Schurle, 1994, 69; Woodard and Garcia, 2008). On the other hand, using aggregated data includes the disadvantage of a loss of information. Climate impact research typically works at lower levels though it focuses on identifying location-specific impacts under climate change.[3] In addition, we acknowledge that statistically more advanced and flexible ways to model systemic risk in yields or weather exist, for instance copulas (e.g., Gaupp et al., 2017; Xu et al., 2010). Our approach, however, targets at disentangling how, in addition to weather, inputs, policy, and macroeconomic shocks specifically drive wheat yields. As such, we connect insights from risk and productivity analysis, agronomic, and climate impact research.

2.3 Data

In what follows we describe the variables for the production function, yields and inputs, followed by weather and phenological stages (all details in the SA).

2.3.1 Production function for wheat

We analyze 12 German federal states[4] for the years 1995—2009. To specify the production function at the regional levels, we use accounting data from the European Farm Accountancy Data Network (FADN) provided by the European Commission (European Commission, 2015a). These data contain representative farms from a stratified, rotating sample (Barkaszi et al., 2009). We refer to published results aggregated at the federal state level and select specialized crop farms referring to the EU's classification (i.e., specialist field crops according to the TF-8 grouping).[5] Our sample represents, on average, 4344 farms per state.

We specify the production function with one output (wheat yield) and eight inputs: capital, labor, wheat acreage, energy, material, services and seed expenses. In material inputs, we summarize fertilizer and plant protection. We use total livestock units per hectare as a proxy for manure.

[3]Input data would be available at the farm level but information about the farm location would be available only at the federal state level due to data privacy reasons.

[4]We exclude the federal state *Saarland* and the cities of *Berlin*, *Hamburg* and *Bremen* due to their small geographical size and minor importance in wheat production.

[5]We preferred the TF-8 data given the higher representativeness; see SA for a comparison of TF-8 and TF-14 grouped data (Figure S2.1).

Except for land, labor and livestock, all inputs are deflated using national price indices provided by the German statistical agency (Statistisches Bundesamt, 2014b) and normalized by the total utilized agricultural area per farm excluding fallow and set-aside land. We include an additional control variable for the share of spring wheat for each federal state and year from German official agricultural statistics (BMEL, 2015).

Land planted with wheat is considered to account for positive specialization and scale effects or negative yield effects resulting from marginal land (Kaufmann and Snell, 1997; Yang et al., 1992). On average, sample farms plant about 65 ha wheat. For historical reasons considerable differences in the farming structure (e.g., size, organizational and ownership structure, technology) between Eastern and Western Germany prevail (Table 2.1). To account for these we use a dummy variable indicating the five Eastern federal states.

Our data cover three major reforms of the EU's CAP that are known to provide incentives for planting crops (background information in the SA). To capture policy and other macroeconomic effects such as the price boom in 2008, we take time dummy variables into account.

TABLE 2.1: Summary Statistics

	Mean	sd	min	max
Wheat yield [100kg ha^{-1}]	69.81	10.66	33.28	92.46
Per farm variables				
Land wheat [ha]	65.42	60.72	8.80	251.73
Land wheat East [ha]	126.80	47.06	46.45	251.73
Land wheat West [ha]	21.59	10.79	8.80	66.63
Total land [ha]	228.37	196.76	36.89	766.25
Total land East [ha]	448.30	94.72	196.90	766.25
Total land West [ha]	71.29	21.49	36.89	135.60
Capital [EUR ha^{-1}]	473.71	172.46	217.82	968.97
Labor [hours ha^{-1}]	51.31	23.06	22.45	142.63
Energy [EUR ha^{-1}]	171.25	41.75	106.36	270.85
Material inputs [EUR ha^{-1}]	308.96	63.03	163.92	517.41
Seeds [EUR ha^{-1}]	93.54	31.17	49.92	196.32
Manure [livestock units ha^{-1}]	0.37	0.20	0.06	0.91
Regional weather variables and controls				
Pot. evapotranspiration stage 1 [mm]	133.32	15.21	103.10	180.60
Prec. stage 1 [mm]	413.40	76.31	219.60	576.60
Prec. stage 1 East [mm]	413.00	78.67	227.20	562.30
Prec. stage 1 West [mm]	413.70	74.96	219.60	576.60
GDD stage 2 [°C]	324.74	33.50	212.96	422.26
Solar radiation stage 2 [kJ cm^{-2}]	60.22	6.49	37.71	77.05
Prec. stage 2 [mm]	74.47	20.80	26.00	117.20
KDD stage 3 [°C]	11.08	7.14	0.64	27.15
Prec. stage 4 [mm]	95.99	29.36	42.80	177.90
Prec. stage 4 East [mm]	97.28	30.91	42.80	176.30
Prec. stage 4 West [mm]	95.06	28.31	50.40	177.90
Share spring wheat [%]	1.97	1.49	0.30	8.59

Note: Data sources are Deutscher Wetterdienst (2012; 2014), FADN (European Commission, 2015a); share spring wheat: BMEL (2015).

2.3.2 Weather and phenological stages

We merge the annual FADN data with daily meteorological observations from 1218 weather stations and phenological data for winter wheat from 5671 stations scattered across Germany provided by the German Meteorological Service (Deutscher Wetterdienst, 2012; 2014).

For all federal states we distinguish four macro phenological periods (Table 2.2) and aggregate all weather variables accordingly (similar e.g., Butler and Huybers, 2015; details in SA). Temperature and solar radiation are mainly responsible for potential crop growth; however, day temperatures above the optimal level temperatures might induce heat stress and decrease wheat yields (Rötter and van de Geijn, 1999). To capture these effects, the average temperature is split into temperatures below and above an optimal temperature of 20°C, above which growing conditions are likely not optimal (Rötter and van de Geijn, 1999). Accordingly, days with temperatures below the optimum but above the minimum of 4°C are denoted as growing degree days (GDD; expected positive effect on wheat growth). Temperatures above 20°C, on the other, lead to heat stress and are summarized for each phenological period as killing degree days (KDD; expected negative impact on crop yields; Roberts et al., 2013, 237).

TABLE 2.2: Four Defined Phenological Periods

Period	Corresponding stages
1	Sowing – stem elongation (minus 1 day) [30]
2	Stem elongation – heading (minus 1 day) [51]
3	Heading – early milk ripening (minus 1 day) [73]
4	Early milk ripening – harvested product [99]

Note: stages following Meier (2001); decimal code in [].

Crop water supply is determined by supply in the form of precipitation and atmospheric demand in the form of evapotranspiration. To appropriately account for water supply, we consider potential evapotranspiration according to Turc-Ivanov (ETP_{TI}) following Conradt et al. (2013) (see SA for details and an alternative measure according to Haude).

Marginal effects of additional water supply depend on actual levels and may switch signs. That is, precipitation might have a positive impact on plant growth if actual water supply is below a plant's optimum. On the other, precipitation might hamper growth of plants in case of water supply being greater than the plant's optimum. Since our weather variables are aggregated over the phenological periods, dry spells are not considered by the sole precipitation amount. Thus, in addition, we consider days without precipitation (DWP) to capture the distribution of precipitation (see SA for details).

Particularly in the eastern parts of Germany, low precipitation amounts might have a higher marginal effect than in the western parts, given the lower soil quality resulting from large shares of sand and low water-holding capacity. We account for this by an interaction of precipitation with the dummy variable $East_i$. Additionally, we consider solar radiation (SR in J cm^{-2}) and temperature normalized radiation (Gornott and Wechsung, 2016).

2.4 Econometric strategy

In what follows, we explain the two steps guiding our analysis.

2.4.1 Empirical model wheat yield variability in Germany 1995–2009

From an economic theorist's perspective, it seems important to consider variable and quasi-fixed inputs. However, economic theory has little to say about functional form, relationships among inputs, and the inter-relation with weather variables (Coelli et al., 2005). Hence, our model building and selection approach is based on an empirical procedure suggested by Greene (2012, 178–80) to select variables with the objective of finding a suitable model that is robust against misspecification (further inspiration from Roberts et al., 2013). Given the rather short number of years available to us, we have to scrutinize a rich set of variables and focus on relevant information. For instance, the set of theoretically optimal weather variables at all phenological periods and the full potentially available set of inputs amounts to 18 and 8, respectively. This is not counting quadratic terms, interactions, alternative weather variables, and variations of functional form (log versus level).

First, we target at identifying an appropriate functional form for the production function relating output, wheat yield y_{it} of state i at period t, and inputs, denoted by x_{jit} where j indexes the inputs capital, labor, land used for wheat, energy, material, seed and manure. Second, the appropriate function relating yield and weather for each agronomic stage must be specified.

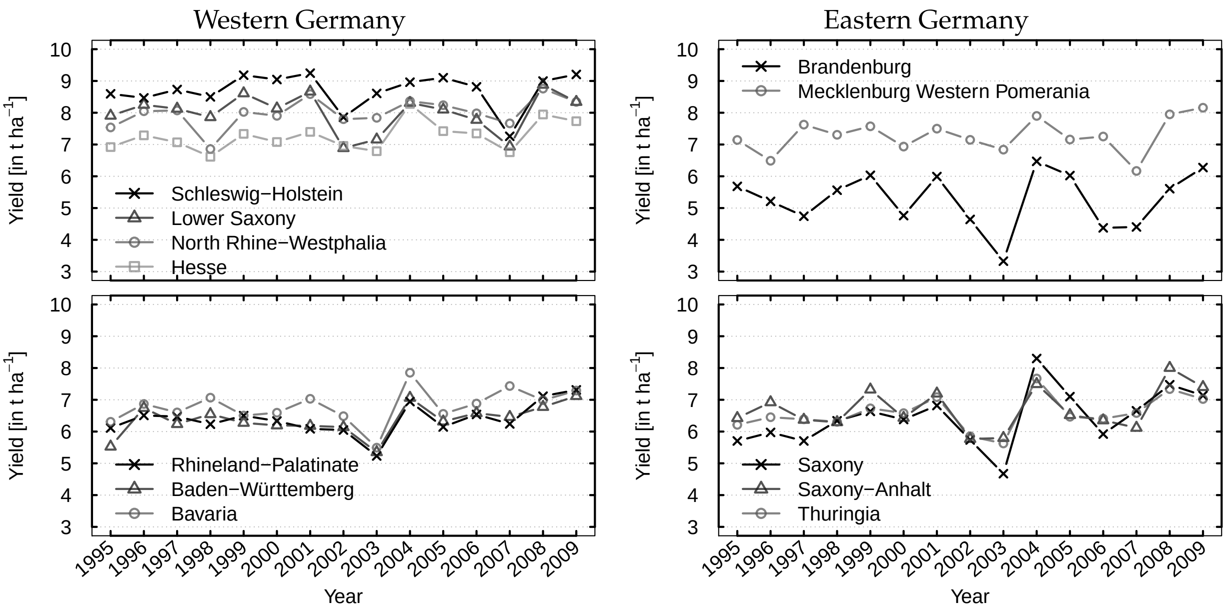

FIGURE 2.1: Yields by federal state, 1995–2009.

We consider wheat yield in log-differences ($\Delta log(y_{it})$) rather than absolute yields to approximate growth rates for two reasons. First, the log-ratio of wheat yields represents a relative change over time and is thus already a measure of relative variability. Second, analyzing first differences has advantages from a statistical point of view. Given the positive trends in the data (Figure 2.1 for visual inspection), we account in a more flexible way for trends by means of the first differences compared to assuming a deterministic trend (Brown, 2013). Moreover, potential unit root problems (the Im, Pesaran and Shin test does not reject the null of a unit root at any conventional level) are usually resolved by first differencing (Chen et al., 2004). In addition, first-differencing eliminates unobserved heterogeneity effects likely present in panel data and reduces problems of serial correlation if data are persistent (Wooldridge, 2009).

This data transformation has the disadvantage of not directly quantifying the effect of, e.g., precipitation or technological change on yield and variance at the same time as proposed by Chen et al. (2004), based on the Just and Pope production function approach (see also Isik and Devadoss, 2006). However, log-differences still allow us to capture the effects of common yield shifts by the constant term, though not (directly) the impact of technological change. The effect measured by the constant accounts for several issues, including technological change, but also yield changes caused by the CO_2-fertilization effect (Attavanich and McCarl, 2014; see also Long et al., 2006, who discuss the potential size of the fertilization effect).

Accordingly, the logged wheat yield ratios are modelled as a function of three components: first, the production function with inputs, $f(x_{jit})$, where j indexes the inputs, and second, the weather function, $g(x_{kit})$, where k indexes the number of weather variables (each aggregated to four phenological sub-periods and counted as different variables).

The third component includes the number of other controls x_{sit} indexed by s and to isolate annual effects induced by policy changes, market, price and other shocks common to all federal states we include yearly dummy variables x_{st}. These dummy variables capture other shocks, including stochastic technological changes at the national level, that depart from the linear federal state-specific trends and thus are not removed by first differencing. The annual dummy variables are also of importance from an econometric point of view since common cross-sectional dependence might prevail if such shocks are not addressed.

Our empirical strategy relies on a rich data set and by exploiting the panel data structure we eliminate potential time-constant sources of confounding by first-differencing. Simultaneous changes in the input-mix based upon expected yield variability (e.g., weather-induced) common to all federal states are captured by time dummy variables. Expectations as long as these result in land-adjustments are captured by the variable land. Also, marginal effects of inputs are estimated conditional on observed weather changes. Thus, if observed weather affects yield variability, this is accounted for while estimating these parameters.

Still, expectations about forthcoming yield changes specific to one or several (but not all) federal states might induce in-season-adjustments of material inputs such as fertilizer. Since these issues are usually unobservable, some inputs in our models might still be confounded with the error term. However, given that about 25% of the total agronomic management costs remain after seeding (KTBL, 2006), and could be adjusted based on very early yield forecasts during the growing season, we argue that the severity of such a simultaneity bias in our context falls within an acceptable range. Additionally, we address this problem by an instrumental variables estimation

in the robustness checks.
The base function is given by:

$$\Delta log(y_{it}) = \Delta f(x_{jit}) + \Delta g(x_{kit}) + \Delta h(x_{st}, x_{sit}) + \Delta \epsilon_{it}, \tag{2.1}$$

herein $\Delta\epsilon_{it}$ denotes the respective error term and symbol Δ indicates the first-differencing operator. Throughout the specification search, we work with yield in log-differences as dependent variable and all explanatory variables in first differences.

We basically test four models that share the same dependent variable but differ in the functional form for the production function $f(x_{jit})$ and the weather component $g(x_{kit})$, as well as the included variables and interactions. Here we borrow from the idea of from general to simple modeling (Greene, 2012) but with the following rules to keep the number of parameters at a reasonable level (see Figure 2.2 for an overview).

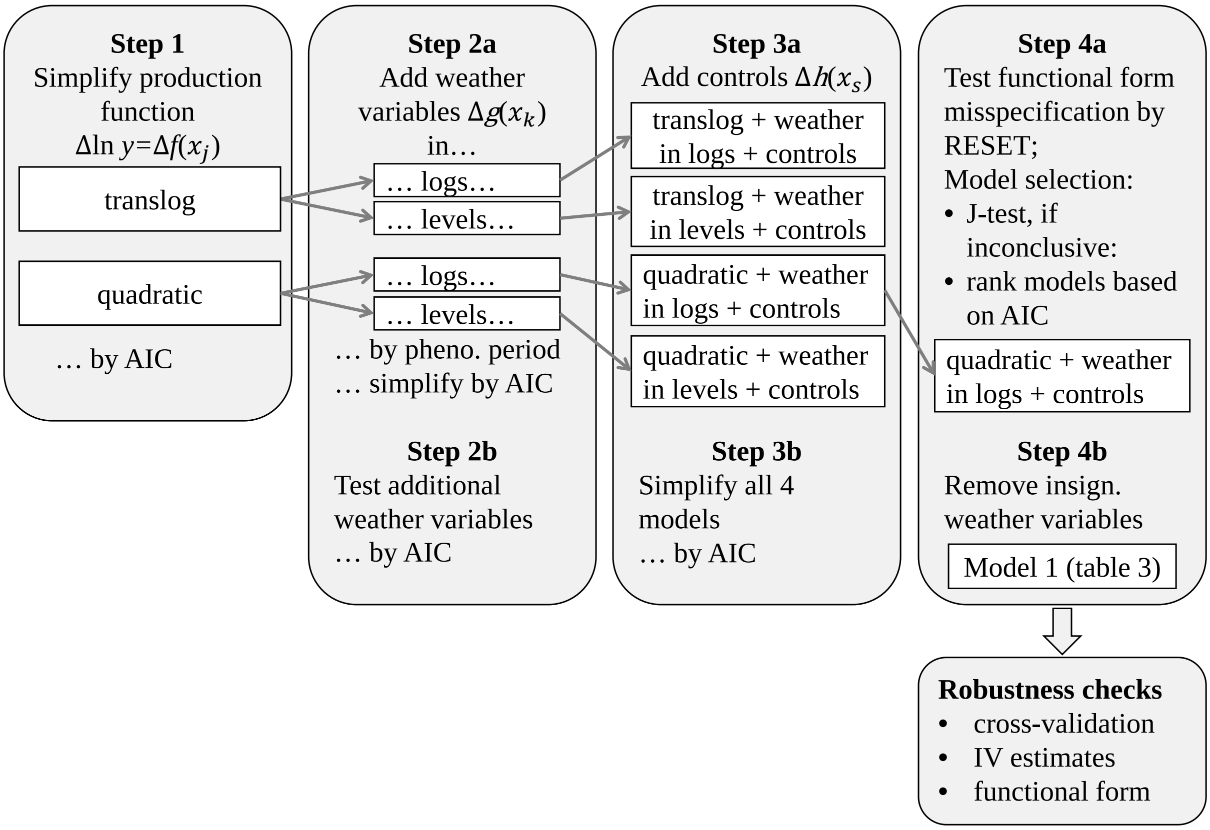

FIGURE 2.2: Workflow of model building and selection. *Note*: AIC: Akaike Information Criterion; IV: instrumental variable; J-test: Davidson and MacKinnon J-test; pheno.: phenological; RESET: regression specification error test; stat. insig.: statistically insignificant (at 10% level).

First, both considered production functions $f(x_{jit})$, that is, the transcendental logarithmic (translog) and quadratic including the full set of interactions, are simplified backwards and forwards to minimize the Akaike Information Criterion (AIC). We force linear terms into the model, while weather is left to the error term in step 1 (Figure 2.2).

Second, both simplified functions are then merged with a theoretically optimal set of weather variables: GDD, precipitation, solar radiation, potential evapotranspiration according to Haude for all four stages, and KDD in stages 3 and 4. We consider two versions of the weather function $g(x_{kit})$: logs and levels (step 2a, figure 2.2). Accordingly, four possible combinations have to be tested: translog with weather either in logs or levels, and quadratic with weather either in logs or levels. The weather variables enter each of these models by phenological period. To select the relevant weather variables at the respective phenological stage, we rely on a backwards and forwards procedure to minimize the AIC (Roberts et al., 2013). Linear terms of weather variables were forced to be part of the model if quadratic terms were considered relevant. The remaining weather variables DWP, minimum and maximum temperatures, and potential evapotranspiration (Turc-Ivanov), are tested by phenological period and kept only if the AIC can be improved (step 2b).

Third, the resulting four models are enlarged as follows. Linear terms are added again if only squared terms and if only interactions are part of the model. Then, additional control variables as well as a full set of year dummy variables are tested (step 3a). Here we consider the share of spring wheat to account for possible lower yields if the share of spring wheat is high. Additionally, we account for the Oder river flood in 2002 and the European heatwave in 2003. Since extreme values might affect the point estimates, particularly those of the weather variables, we interact those federal states most affected by the flood (see SA) with the year dummy 2002 (x_6, eq. (2.4)). Since Brandenburg suffered most from the heatwave (Figure 2.1), we interact this state with 2003 (x_7, eq. (2.4)). Again, we simplify the four enlarged models applying the back- and forwards procedure to minimize the AIC (step 3b).

Fourth, since we can rule out any remaining misspecification of functional form (RESET passed), we further carry out Davidson and MacKinnon J-tests to choose between these non-nested models. Since these results were inconclusive, we opted for the model with the lowest AIC (step 4a).[6] After all, the log-level specification, a commonly used approach in applied econometrics, e.g., Wooldridge (2009, 45–6), as well as in climate impact research (e.g., Lobell et al., 2011), seems to best fit to the input data. That is, the yields and the weather are modeled in logarithmic form (e.g., You et al., 2009), while inputs are modeled in levels. At the end of this procedure, we finally remove four more weather variables, which were statistically insignificant (robust standard errors; step 4b). Results were robust to inclusion, though we placed greater weight on econometric efficiency here (see SA, Table S2.8).

The final model, labeled as *model 1*, is given by eq. (2.1) with $f(.)$, $g(.)$ and $h(.)$ defined as follows:

[6] Table S2.6 in the SA shows the ranking of the four models that differ by their functional form of the right-hand side of the production function and the weather variables using the respective value of the AIC.

$$
\begin{aligned}
f(x_{jit}) = \sum_{j=1}^{7} \beta_j x_{jit} + \frac{1}{2} \sum_{j=1}^{5} \beta_{jj} (x_{jit})^2 + \beta_{12}(x_{1it} x_{2it}) \\
+ \beta_{13}(x_{1it} x_{3it}) + \beta_{24}(x_{2it} x_{4it}) + \beta_{35}(x_{3it} x_{5it}).
\end{aligned} \tag{2.2}
$$

Aside from services, all inputs x_{jit} remain in the final model specification. Capital (x_{1it}) appears in interaction with labor (x_{2it}) and seeds (x_{3it}), while labor with energy (x_{4it}), and seeds with manure (x_{5it}). Symbols x_{6it} and x_{7it} denote material inputs and wheat land, respectively. Seven weather variables in logarithmic terms enter the model in the spirit of a translog production function:

$$
\begin{aligned}
g(x_{kit}) = \sum_{k=1}^{7} \beta_k log(x_{kit}) + \beta_{11} \frac{1}{2} log(x_{1it})^2 \\
+ \sum_{k=2}^{3} \left[\beta_{k2} log(x_{kit}) East_i + \beta_{k3} \frac{1}{2} log(x_{kit})^2 East_i \right]
\end{aligned} \tag{2.3}
$$

with x_{1it}: solar radiation stage 2, x_{2it}: precipitation stage 1, x_{3it}: precipitation stage 4, both interacted with a dummy variable for the Eastern German federal states ($East_i$), x_{4it}: potential evapotranspiration (ETP_{TI}) stage 1, x_{5it}: GDD stage 2, x_{6it}: precipitation stage 2, x_{7it}: KDD stage 3.

$$
h(x_{st}, x_{sit}) = \sum_{s=1}^{13} \beta_s x_{st} + \beta_{61}(x_6 flood_i) + \beta_{72}(x_7 Brandenburg_i) + \beta_{14} x_{14it} \tag{2.4}
$$

herein x_1 to x_{13} denote annual dummy variables for 1997–2009 and x_{14it} the share of spring wheat on total wheat land. The dummy variable $flood_i$ is set to 1 for states that suffered from the flood in the year 2002; $Brandenburg_i$ denotes a dummy variable for this federal state.

Given the interaction terms, we consider all variables except yields and dummy variables in mean-centered form, that is, each observation is normalized by its corresponding sample mean. Using panel data allows us to eliminate unobserved heterogeneity effects by first-differencing, and thus all models are estimated by Ordinary Least Squares (Greene, 2012). To address the potential endogeneity problem while estimating the effect of material inputs, we apply an instrumental variable (IV) estimation approach as a robustness check, where we use the second lagged differences of material inputs as instruments (see Table S2.10 in the supplementary appendix). The IV estimates reveal the same qualitative results (signs preserved) but a modestly higher estimate for material inputs; however, IV estimates are only consistent and do not epitomize an unbiased point of reference (Wooldridge, 2009, 510). Hence, we proceed upon the OLS estimation approach on first-differenced data, but provide the IV estimates in the supplementary appendix.[7]

To discuss the potential of an omitted variables bias, we present another model leaving out input variables similar to Miao et al. (2016, 201). To define *model 2*, within eq. (2.1) all inputs (eq. (2.2)) are removed. We assume that the specification search would have led to choosing the same weather variables.

[7]Dynamic modeling approaches fail in this context due to the limited number of observations.

2.4.2 Investigating the effect of inputs and weather on yield volatility

Model 1 passes the RESET procedure and hence we conjecture that model 1 is linearly separable in parameters (Table S2.7, SA). This is a pre-condition for an unbiased and reliable decomposition of the wheat yield variability as carried out in the second step of this empirical analysis.

Generally, two approaches to measure crop yield variability exist: absolute or relative. Chen et al. (2004), for instance, refer to an absolute measure. These authors rely in their rigorous analysis on a Just and Pope type production function approach to quantify weather effects on mean yield and its variance. In these type of models, however, the dependent variable must be stationary without first differencing, where our yield data require first differencing to ensure stationarity. Furthermore, absolute measures rely on absolute changes in yields, which might lead to seemingly increased risk if positive time trends prevail (Finger, 2010). Thus, we focus on relative risk measures to ensure the comparability of weather-induced wheat yield variability by region and over years. Relative variability, specifically volatility, can be measured for instance by using the standard deviation of time differences of logged yields (log-returns), or by a coefficient of variation (CV) for a time series of yield levels (Finger, 2010, 177; Ray et al., 2015, 2).

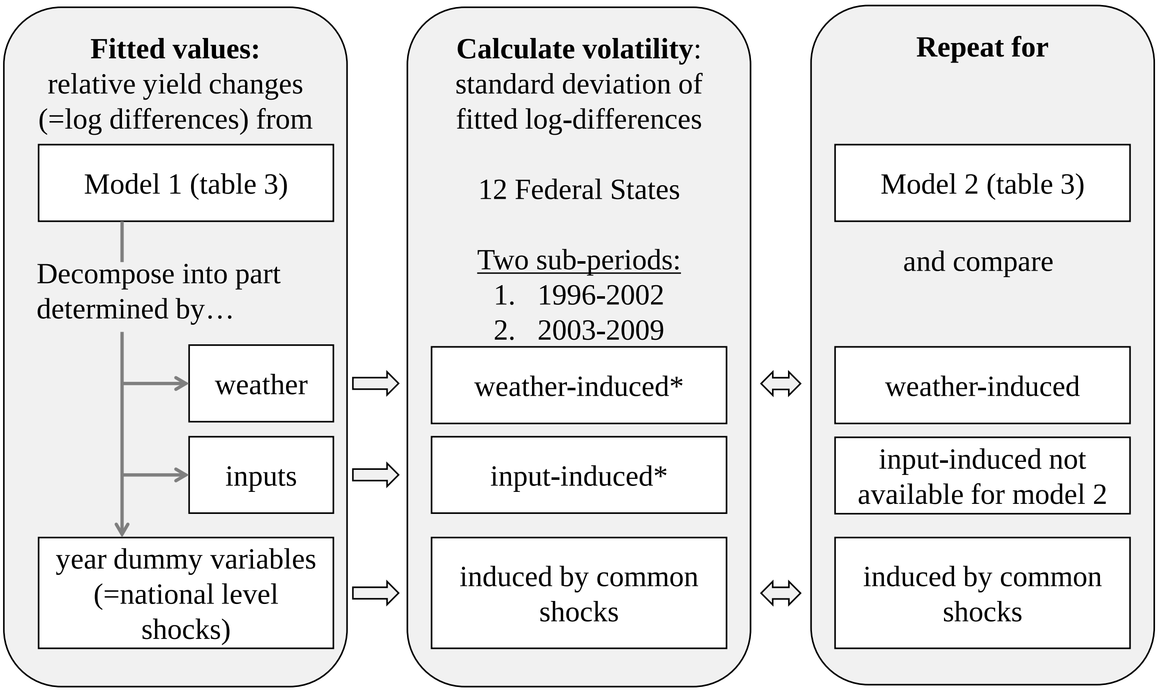

FIGURE 2.3: Workflow of volatility decomposition. *Note*: *: decomposition illustrated in Figure 2.4. Full results in SA, Table S2.11.

In this second step, we extract and analyze weather-determined yield variability, similar to Osborne and Wheeler (2013). We augment these authors' approach by adding inputs, a wider range of weather variables, and analyzing a regional panel data set inspired by You et al. (2009, 1012).

In contrast to Osborne and Wheeler (2013, 4, 7) and Ray et al. (2015, 2), we separate the weather explained portion of the yield from the input-determined part (see Figure 2.3). In addition, our volatility measure directly refers to the weather-induced variation of yields, while the measure "*climate explained yield variability*" of Ray et al. (2015, 2) is in absolute terms not independent of the yield dimension.

Weather caused relative yield changes $\widehat{y_{it}^{w}}$ are defined as the fitted first differences of log yields in the regression model (eq. (2.1)) resulting from weather variations following eq. (2.3). That is, inputs are evaluated at their means (zero since in mean-centered form) and other controls such as year dummy variables are not included: $\widehat{y_{it}^{w}} = \widehat{\Delta log(y_{it})} = \widehat{\Delta g(x_{kit})}$. Based on the fitted series we extract weather-induced volatility $\widehat{v_{i}^{w}}$ for each state $i = 1, ..., n$ using the standard deviation over $\widehat{y_{it}^{w}}$:

$$\widehat{v_{i}^{w}} = \sqrt{\frac{1}{T-1} \sum_{t=1}^{T} \left(\widehat{y_{it}^{w}} - \overline{\widehat{y_{i}^{w}}}\right)^{2}} \tag{2.5}$$

with year $t = 1, ..., T$, $T = 7$ and mean $\overline{\widehat{y_{i}^{w}}} = \frac{1}{T} \sum_{t=1}^{T} \widehat{y_{it}^{w}}$.

To capture changes in weather-induced risk over time, we divide the sample into two equal-length sub-periods: 1996—2002 and 2003—2009. To compare weather-induced volatility changes and those induced by inputs, we adopt this approach for input-determined yield changes accordingly. For additional comparisons, we calculate volatilities based on fitted values allowing both—inputs and weather—to fluctuate while excluding controls and year dummy variables.

2.5 Results and discussion

First, we report the results of the production function estimation and the robustness checks; second, we discuss the estimated volatilities as determined by weather and inputs. In Table 2.3 we present the estimates for two models.[8] While model 1 refers to the fully specified model, model 2 is dedicated to the omitted-variables-bias discussion. All inputs are in levels. Hence, the coefficients can be interpreted as semi-elasticities while the weather variables enter in logs; these estimates can be interpreted as elasticities. A Davidson and MacKinnon J-test and AIC values show that model 1 is superior to a log-log model (see SA, Table S2.9, model 5).

2.5.1 Production function inputs

The considerable number of statistically significant interaction terms underpins our choice of a flexible functional form. *Energy* and *material* inputs reveal positive linear and negative quadratic effects, though energy only does so in a significant quadratic term. That is, material inputs positively affect yield changes with decreasing marginal productivity. Starting from the sample mean, a 10% increase (roughly 30 Euros per hectare) leads to positive yield changes of about 1.7%. Be-

[8]The p-values are based on spatial correlation consistent (SCC) standard errors, which are also robust to (cross-) serial correlation (Driscoll and Kraay, 1998; Millo, 2014). Note that the time dimension of our data is relatively small ($T = 14$) and as such at the lower bound (Driscoll and Kraay, 1998, 556).

TABLE 2.3: Effects of Inputs and Weather on Relative Wheat Yield Variability in Germany, 1996–2009

	(1) Final specification	(2) Drop inputs
Intercept	0.072 (0.0116)***	0.083 (0.0127)***
Inputs		
Capital	0.043 (0.0399)	
Labor	−0.195 (0.1743)	
Energy	0.123 (0.0765)	
Material inputs	0.091 (0.0305)***	
Seeds	−0.142 (0.0495)***	
Manure	−5.540 (6.9399)	
Land wheat	−0.050 (0.0418)	
Capital squared	0.001 (0.0002)***	
Labor squared	0.016 (0.0095)*	
Energy squared	−0.011 (0.0020)***	
Material inputs squared	−0.001 (0.0002)***	
Seeds squared	0.004 (0.0011)***	
Capital * labor	−0.004 (0.0015)***	
Capital * seeds	−0.001 (0.0001)***	
Labor * energy	0.011 (0.0036)***	
Seeds * manure	0.288 (0.1611)*	
Weather		
Prec. stage 1	0.050 (0.0159)***	0.022 (0.0194)
Prec. stage 1 * East	−0.167 (0.0439)***	−0.139 (0.0783)*
Prec. stage 1 squared * East	−0.913 (0.1412)***	−0.970 (0.3244)***
Pot. evapotranspiration stage 1	0.225 (0.0512)***	0.250 (0.0757)***
Growing degree days stage 2	0.114 (0.0313)***	0.125 (0.0727)*
Solar radiation stage 2	−0.023 (0.0423)	0.034 (0.0743)
Solar radiation stage 2 squared	0.442 (0.1561)***	0.133 (0.1236)
Prec. stage 2	−0.019 (0.0067)***	−0.015 (0.0091)*
Killing degree days stage 3	−0.018 (0.0027)***	−0.017 (0.0040)***
Prec. stage 4	−0.039 (0.0164)**	−0.044 (0.0195)**
Prec. stage 4 * East	0.084 (0.0519)	0.095 (0.0408)**
Prec. stage 4 squared * East	0.262 (0.1615)	0.261 (0.1064)**
Controls		
Share land spring wheat	−1.020 (0.4498)**	−0.840 (0.6020)
Year 2003 * Brandenburg	−0.388 (0.0229)***	−0.386 (0.0145)***
Flood 2002	−0.073 (0.0134)***	−0.103 (0.0152)***
Year 1997	0.005 (0.0189)	−0.058 (0.0071)***
Year 1998	−0.051 (0.0283)*	−0.149 (0.0192)***
Year 1999	−0.073 (0.0494)	−0.180 (0.0438)***
Year 2000	−0.199 (0.0485)***	−0.334 (0.0466)***
Year 2001	−0.217 (0.0584)***	−0.340 (0.0555)***
Year 2002	−0.311 (0.0677)***	−0.476 (0.0670)***
Year 2003	−0.466 (0.0860)***	−0.673 (0.0892)***
Year 2004	−0.351 (0.0860)***	−0.533 (0.0973)***
Year 2005	−0.511 (0.1075)***	−0.755 (0.1220)***
Year 2006	−0.602 (0.1175)***	−0.856 (0.1294)***
Year 2007	−0.722 (0.1196)***	−0.946 (0.1338)***
Year 2008	−0.638 (0.1349)***	−0.889 (0.1512)***
Year 2009	−0.688 (0.1482)***	−0.994 (0.1681)***
R^2	0.83	0.74
Adj. R^2	0.77	0.69

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: first differences of logged wheat yield. Num. obs.: 168. Weather in logs, inputs in levels. Coefficients/standard errors for inputs multiplied by 100. Explanatory variables first differenced; weather/inputs mean centered. Spatial and serial correlation robust standard errors in () (Driscoll and Kraay, 1998). Pot.: potential; prec.: precipitation.

cause of the non-linear relationship, following a 30% increase (roughly 90 Euros), yields would already decrease by 1.06%.

Seeds show a negative linear and positive quadratic coefficient. This implies that, starting from the sample mean, a reduction of seeds would lead to soaring yields, whereas increases would first lead to decreases and then to increases again. This effect might be traced back to the variable definition in monetary terms. Reducing seeds would lead to increasing yields, though low seed densities require nearly perfect water supply conditions. Agronomic relations might explain the negative range. Late sowing requires a higher rate of seed input per hectare to ensure the full development of a plant population. On the other hand, too dense populations can reduce yields.

Turning to *capital* and *labor*, two quasi-fixed inputs, we find positive linear and quadratic effects of capital changes on yield growth rates but negative interaction effects with labor (and seeds). For labor, we find negative linear and positive quadratic terms. That is, any deviation from the sample mean in labor would cause increases in the yield growth rate evaluated at the sample mean. These estimates should, however, be interpreted in light of the observed trend in the data to reduce overall capital and labor input per hectare from 1999 on. To disentangle the capital-labor relationship, we follow the idea of simple slopes in two-way interactions and use capital as a moderator. We find for *low levels of capital* (mean minus standard deviation) that a reduction of labor input negatively affects yield changes, whereas additional labor units at low capital levels contribute positively to the growth rate of wheat yields. For *high capital levels* (mean plus standard deviation), the substitution effect is less obvious: deviating from the sample mean level of labor causes positive impacts, whereas the positive effect of reducing labor input is more pronounced. This implies that in high capital production systems, capital productivity might be improved through labor reduction.

The effect for *land planted with wheat* remains insignificant. This might be traced back to two opposing effects. First, due to specialization and scale effects, we would expect yield to increase in land planted with wheat. Second, more marginal land might be used for cropping wheat in the course of time, incentivized for instance by rising wheat prices (Haile et al., 2016). In addition, such land might be more sensitive to changes in weather conditions, and thus we would expect a negative effect of a higher share of marginal land on wheat yields. While the first effect might be particularly relevant for Western Germany, the second effect might be more relevant for the eastern part.

All time dummy variables are significant from 2000 onwards; all are negative. That is, compared to changes between 1995/96, the wheat growth rate decreases for these years. Changes in the relative competitiveness of wheat due to de-coupled direct payments within the CAP (starting from 2000 onwards) might provide an explanation for this finding. These coefficients capture common yield shifts, also possibly due to weather or other common macroeconomic shocks (e.g., prices, technological change), the effects of which cannot be isolated. For example, our modeling approach does not allow us to identify the direct impact of technological change, nor to disentangle the effect of CO_2 on yield variability as shown by Attavanich and McCarl (2014). The time dummy variables for the flood and the European heatwave reveal significant negative impacts on the wheat growth rates for the respective federal states.

2.5.2 Weather

In the first early development stage of the plant (sowing/end of tillering), we find significant positive effects of precipitation levels and potential evapotranspiration (ETP_{TI}): a 10% increase in ETP_{TI} would lead to a 2.3% increase in yields. A sufficient water supply improves biomass production, determining the yield potential of the plant (Chmielewski and Köhn, 2000). Also, Roberts et al. (2013) find positive effects of a vapor pressure deficit, a main component of the evapotranspiration measure. For the Eastern German federal states, which are known for soil conditions with low water holding capacity, we find negative coefficients for the linear and quadratic terms of precipitation in stage 1. Given the predominating soil condition in Eastern Germany, nutrient leaching might be problematic at higher precipitation levels in the first phenological stage. This might hamper yield potentials. For the fourth stage (early milk ripening/harvesting), we find negative impacts of precipitation, that is, as expected, in late-ripening crops, additional water may lead to yield losses. However, for Eastern Germany, we find positive impacts of precipitation; marginally significant with $p = 0.106$ (linear term) and $p = 0.108$ (squared term). Since the fourth stage includes early milk ripening, in which water scarcity is very likely in the Eastern federal states compared to other regions in Germany, additional water supply could foster increased yield quantities (Fricke and Riedel, 2015).

For the second stage (stem elongation until heading), we find negative effects of precipitation and positive effects for GDD. That is, temperatures below the optimal level positively influence wheat growth and thus yields. Solar radiation has a non-linear positive effect that is attributable to increased photosynthesis (Roberts et al., 2013). An increasing water supply in this developmental stage, however, rather hinders growth as indicated by the negative coefficient. This result points to water supply being close to optimum on average. In the third stage (heading/early milk ripening), we find considerable impacts of temperature: KDD affect yield negatively, which is in line with existing research (Roberts et al., 2013). These effects, however, remain small: starting from the sample mean, an increase in KDD by one standard deviation (approx. 65%) would lead to yield losses of 1.2%.

Despite taking regional KDD into consideration, we find an additional significant effect of the heatwave for Brandenburg. This might be traced back to Brandenburg's natural conditions, particularly sandy soils with low water-holding capacity (Wessolek and Asseng, 2006) and uncaptured soil-specific heatwave dynamics (Perkins, 2015). Together with different effects of precipitation for Eastern and Western Germany in two phenological stages, our results reveal the importance of the spatial-temporal distribution of water supply and its dependence on soil conditions.

To summarize, all weather effects can be grounded on agronomic-theoretical explanations and are in line with previous findings. Our variable selection and phenological data aggregation reflect the complexity of yield formation. Model 1 may be criticized regarding the inclusion of the extreme weather years. Thus, we performed a leave-one-out cross-validation confirming the robustness of the model (e.g., Blanc, 2012; details in SA).

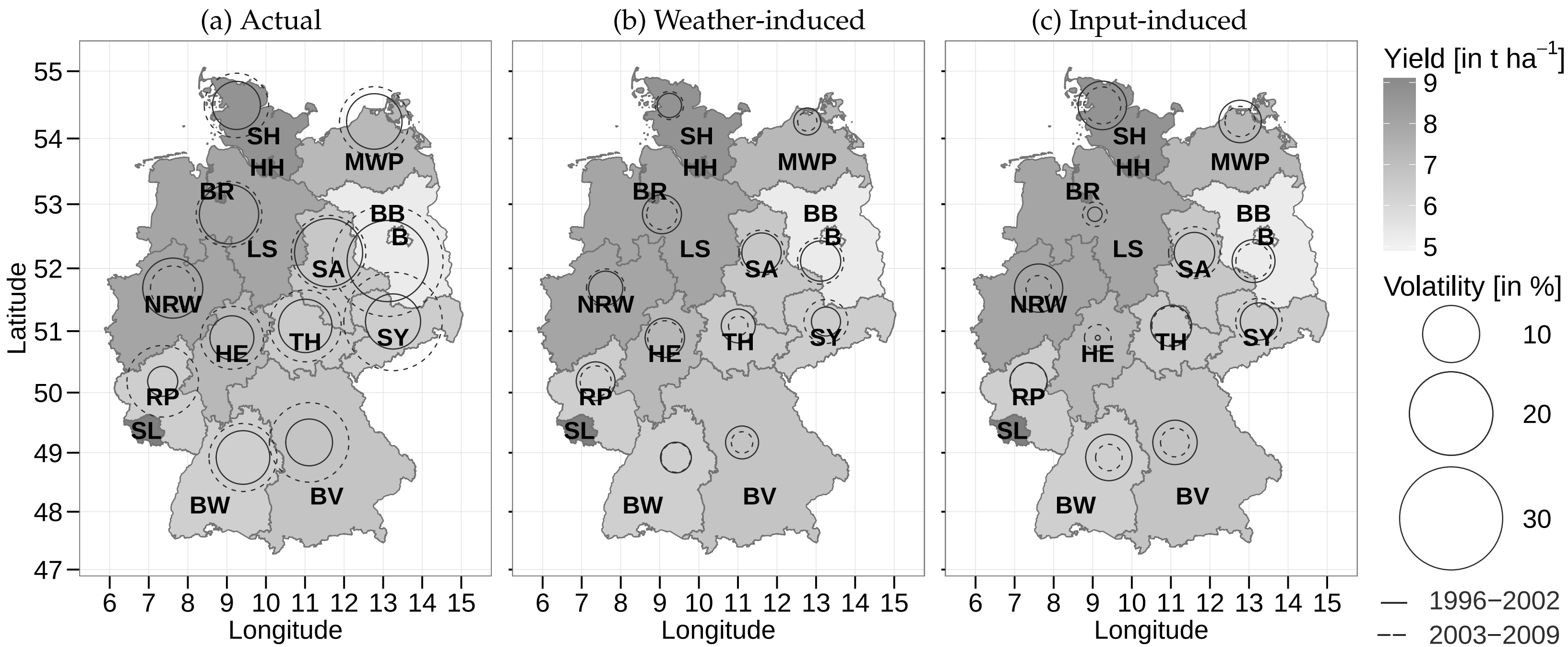

FIGURE 2.4: Actual, weather- and input-induced wheat yield volatility for two sub-periods.

Note: Bubbles indicate volatilities of different magnitudes. B: Berlin, BB: Brandenburg, BR: Bremen, BV: Bavaria, BW: Baden-Wuerttemberg, HE: Hesse, HH: Hamburg, LS: Lower Saxony, MWP: Mecklenburg-West Pomerania, NRW: North Rhine-Westphalia, RP: Rhineland-Palatinate, SA: Saxony-Anhalt, SH: Schleswig-Holstein, SL: Saarland, SY: Saxony, TH: Thuringia.

2.5.3 Decomposing wheat yield volatility

To answer the question of how volatility differs across regions and over time, as well as to disentangle its drivers, we decompose the standard deviation of the wheat growth rates (see Figure 2.3). We illustrate these measures in Figure 2.4 (based on Table S2.11, SA). Actual, weather- and input-induced volatilities are plotted for two sub-periods: 1996—2002 and 2003—2009 (grey-solid and grey-dashed circles). Averaging over regions and time, inputs explain ca. 49% of the total actual wheat yield volatility, while weather explains 43% (evaluated at the sample means, based on values in Table S2.11, SA). Comparing actual volatilities for the sub-periods over time, wheat yield volatility increases except for one state (North Rhine-Westphalia, Figure 2.4-a). Riskier areas regarding weather and inputs are found in the eastern part of Germany.

We use regional aggregated yield data at the federal state level. Spatially uncorrelated risks, that is, idiosyncratic shocks, "*self-diversify*" at this higher aggregation level compared to firm-level data, while more systemic variation remains (Woodard and Garcia, 2008, 37; Marra and Schurle, 1994).[9] Hence, weather-induced volatility at the state level can be interpreted as a measure of systemic weather risk in agricultural production (cf. Xu et al., 2010, 267–268). As illustrated in Figure 2.4-b, weather-caused volatility differs slightly by region with higher volatilities in the eastern part. Comparing these volatilities with those caused by input adjustments (Figure 2.4-c), we find for the entire eastern region, as well as some western regions, higher input-induced volatilities compared to the volatilities traced back to weather changes (e.g., Bavaria). Over time, we observe increases in actual volatility, on average. However, this can only be traced back to joint increases in weather and inputs in some regions (e.g., Saxony), while in other regions weather- and input-induced volatility changes reveal opposite signs. For instance in Brandenburg, weather-induced yield volatility considerably increases but input-induced volatility decreases. Still, the overall increase of actual volatility cannot be fully traced back to weather and input adjustments.

We find a higher share of explained actual volatility, for the first compared to the second period. To illustrate an extreme, in Hesse from 1996–2002, about 93% of the actual volatility are explained (6.11 *volatility inputs and weather* / 6.55 *actual volatility*, values according to Table S2.11, SA); from 2003–2009 this amounts to 35%, however. Averaging over all regions, we find 83% of the actual volatility explained by inputs and weather from 1996–2002, while 44% in the period 2003–2009. These findings can be explained in part by the use of time dummy variables in the regression model, which are isolated in the volatility measures for inputs and weather but are particularly important in the second period. Hence, we likely underestimate the weather effect because common weather shocks are captured by time dummy variables. As a robustness check, we investigated whether the different geographical sizes of the states affect our results (aggregation bias); this is not the case (see SA for details).

[9]Note that the illustration of the aggregation argument by Woodard and Garcia (2008, 37–38) does not acknowledge that a part of the weather risk might be included among the idiosyncratic risks that self-diversify. Precipitation and related variables that are functions of the latter are expected to vary more across space than temperature.

Weather-driven volatility at the state level seems to be rather low given that we would expect higher changes caused by varying weather conditions. In other words, we conjecture that systemic risk cannot only be traced back to weather as measured in our model (regional temperature, solar radiation, precipitation and evapotranspiration). Common shocks at the macro (i.e. national) level are relevant. The latter include weather extremes but also policy and price changes affecting many farms as well as consequential input-adjustments. The significant year dummy variables from 2000 onwards capture exactly such macroeconomic and policy changes (Table 2.3). In this period several reforms of the CAP affected farms' production (intensity) decisions, for instance, the decoupling of direct payments from the crop being planted starting in Germany in 2000, which was discussed in 2003, reinforced in 2005 and verified in 2008. The price boom for agricultural commodities in 2007/08 also occurred during this period. The national level volatility based on year dummy variables reflects the increasing importance of common shocks: 5% in 1996–2002, 11.2% in 2003–2009 (see Figure 2.3 and Table S2.11, SA).

How would the results, particularly the weather-induced volatility, look like if input adjustments were neglected? Similar to the full model 1, the estimates for the reduced form model 2 reveal increases in weather-induced volatility for some regions, while for others decreasing measures for the second period prevail (not shown in Figure 2.4; results in Table S2.11, SA). The unexplained share of volatility and the national level volatility are also higher in the second period. Averaging over regions and time, weather explains practically the same fraction of the total actual volatility in model 2 (both models: 43.5%). Comparing the regional volatility estimates within period 1996–2002, some are overestimated in model 2 (e.g., Saxony-Anhalt and Brandenburg) while some are underestimated (e.g., Hesse). In period 2003–2009, the majority of the measures of the reduced form model overestimate the weather component, though in some regions only by a minor rate. For only one state (Hesse) the two models differ qualitatively: while the full model detects decreases in weather-induced volatility in period 2003–2009, model 2 finds small increases. Thus, one could draw misleading conclusions regarding the weather-induced volatility in both sign and size while neglecting inputs. However, and most importantly, the unexplained part is higher compared to the full model. As a consequence, too much emphasis would be placed on the interpretation of the common shocks (significant for all years in model 2, considerably higher estimates; this results in higher national level volatility). As such, the systemic macro risk would be overestimated. At the same time, input-adjustments as possible consequences of price and policy shocks, which are simply rational adaptation by farmers, would not be discussed at all.

2.6 Concluding remarks

Wheat is a major commodity that plays a crucial role for food security. The recently observed increases in relative wheat yield variability for Germany—an important wheat producer in the EU—begs the question: Can these increases be traced back to weather changes? Or is it "simply" the result of farms' adaptations to changing institutional and macroeconomic conditions leading to adjustments in their input-mix? To answer these questions, we analyze relative wheat yield variability consistent with production economics and agronomic climate impact research. We use a rich set of regional accountancy data and weather variables at the respective phenological stages from 1996 to 2009. Obtained wheat-yield volatilities are decomposed into weather- and input-driven categories.

In line with production economics and agronomic research, we find that both inputs and weather impact relative yield changes. Common shocks at the national level play a significant role from 2000 onwards, a period characterized by fundamental changes in the EU's CAP and price booms for agricultural commodities. Decomposing wheat yield volatility reveals regionally heterogeneous weather-induced instabilities. Splitting the sample into two sub-periods, we find increases in actual volatility over time, where macro-level shocks including weather extremes contribute. These increases, however, can only in some regions be traced back to joint increases of the weather-induced component and the part caused by adjustments in the input-mix. A number of regions even show decreases in weather-caused volatility over time.

This study is relevant for several reasons. First, future climate impact analyses, which inform policy makers, could utilize this case study as a proof of concept. We could show that omitting inputs would rarely alter our results in a qualitative manner, though would do so quantitatively. Weather impacts and common shocks would be overestimated in the case of leaving out input choices, and adjustments in the input mix would not be discussed at all. We thus contribute to the debate of whether inputs should be a part of climate impact research, where purely statistical approaches are still prominent (Liu et al., 2016; Miao et al., 2016). To conclude, independent of the model type, relating yield and weather offers reasonable results and valid approximations.

Second, these insights support approaches such as the European Commission's MARS[10] project, which is relevant for policy makers for crisis intervention. Considering yield vulnerability by phenological stage at the regional level could improve the seasonal forecasting of potential crop shortages attributable to weather. Better knowledge about yield vulnerability might also help farmers adjust their agronomic management to better cope with downside risk (Chipanshi et al., 2015).

Third, wheat-yield vulnerability by phenological stage at the regional level might also be of interest for insurance design and modeling weather risk (Conradt et al., 2015; Odening and Shen, 2014). Since our approach decomposes the influences of weather- and input-related impacts on wheat yields and averages out idiosyncratic shocks, it might help insurers to improve the determination of insurance claims. For insurers, it might be relevant to only indemnify weather-related yield losses. In addition, a more cost-effective assessment of common weather-related yield losses might enable insurance companies to better cope with systemic risk. This would benefit both the insurer and the insured (cf. Finger, 2013). Insurance-based solutions have recently gained attention because of their potential to contribute to stabilizing farm incomes and thus food security, particularly in regions where smallholder farming prevails (Surminski et al., 2016).

Since many of the European CAP reforms aim to reduce the impact that disbursed subsidies have on input-intensity and to protect the environment at the same time (Levers et al., 2016), our

[10]Monitoring Agricultural Resources: `https://ec.europa.eu/jrc/en/mars`; accessed 16.02.2017.

results may further offer insights into how wheat yield variability is related to the interplay between weather, input-use and agricultural policy. These may help investigate how to reduce distorting policy impacts on input-intensity, while taking into account that these choices also relate to farmers' risk mitigation strategies. This in turn is a pre-condition for ensuring secure, resilient and sustainable food production. The recently-established risk management toolkit under pillar two measures of the EU's CAP might offer a reasonable starting point, though it has seen heterogeneous acceptance among member states thus far. Additionally, as shown by Gaupp et al. (2017), wheat yields are independent at the global level. This is a pre-condition to stabilize food supply by international trade, which could be another option for policy makers (e.g., Brown et al., 2017).

The work presented here displays some shortcomings. First, our data do not allow us to account for land use changes. For instance, farmers might reallocate land for highly subsidized renewable energy plants closer to the farm to save transportation costs. As a consequence, wheat might be reallocated to more distant plots, possibly with lower soil quality. The land used for wheat would not change overall, but yields would be more sensitive to weather impacts. Additionally, yield variability might be subject to technological progress (Chen et al., 2004), which could not explicitly be modelled with our data set. Using improved technology to increase agricultural productivity could be one way to globally ensure the stable supply of sufficient food quantities (Pardey et al., 2016). As shown by Emerick et al. (2016), against the backdrop of increasing weather risk, particularly new seed varieties with a reduced downside risk have the potential to crowd-in inputs such as fertilizer to increase yields (in addition to the positive agronomic effect on yield). Given that farmland expansion is already at its limit and in some regions only possible at the costs of biodiversity (e.g., Foley et al., 2011), such technical change could contribute to closing yield gaps with less negative environmental impact that the pure intensification of crop production by increasing fertilizers or irrigation would likely have.

Finally, from a producer's perspective, economic risk matters as well. Output price variation has increased in the recent decade and proven to reduce production intensity (Haile et al., 2016). Future research analyzing weather impacts on agricultural production should thus consider farm-level input adaptations to changes in weather- and price-risk as well as policy changes and macroeconomic developments.

Chapter SA2

Supplementary appendix: How do inputs and weather drive wheat yield volatility? The example of Germany

Authors: Hakon ALBERS*, Christoph GORNOTT† and Silke HÜTTEL‡

SA2.1 Literature Review

First, we discuss how weather can be included in production functions. Second, overview tables introduce the reviewed papers in a condensed way.

SA2.1.1 Three choices

Three choices are important for including weather into the production function framework: the *selection* of weather variables, *aggregation* levels of weather data and the *functional form* describing the input-output and weather-yield relationships.

Among the ***weather variables***, the majority employs temperature and precipitation, where radiation and evapotranspiration are also found, though sparsely. Variables capturing soil moisture are rarely applied since this requires spatially detailed data usually not available (e.g., Bakker et al., 2005). More recent papers emphasize vapor pressure deficit (VPD) as an important yield-determining variable (e.g., Lobell et al., 2014; Roberts et al., 2013). However, similar to other applied evapotranspiration variables, VPD does not include the water-holding capacity of soil and might thus indicate dry conditions in cases where water supply is sufficient. In addition, atmospheric CO_2 concentration has been proven to be important for yield levels, but in the literature we reviewed its small variation over time and space prevented researchers from quantifying its impacts in econometric models (e.g., Finger and Schmid, 2008, 26) with only few exceptions (e.g., Blanc, 2012).

As agronomic knowledge suggests, to ***aggregate data*** at different phenological stages improves evidence of weather effects on yield levels/variability, although data requirements are high (Dixon et al., 1994). For instance, Butler and Huybers (2015) or Ortiz-Bobea and Just (2013)

***Corresponding Author**. Martin Luther University Halle-Wittenberg, Department of Economics, Universitätsring 3, D-06108 Halle, Germany

†Potsdam Institute for Climate Impact Research (PIK), Germany.

‡Rheinische Friedrich-Wilhelms-Universität Bonn, Institute of Food and Resource Economics, Production Economics Group. Former address: University of Rostock, Agricultural Economics, Germany.

incorporate phenological stages while assessing impacts of weather conditions on maize yields to analyze adaptation possibilities of farms. While these studies emphasize temporal aggregation, spatial aggregation has been analyzed in greater detail by Garcia et al. (1987).

Finding the appropriate ***functional form*** describing the input-output relationship is an ongoing research topic, particularly in production economics literature (e.g., Coelli et al., 2005, Griffin et al., 1987). Indeed, an appropriate functional form is crucial since biases due to functional form misspecification are in the same dimension as omitted variable biases. Typically applied functional forms include Cobb-Douglas, linear, or quadratic with linear-additive weather effects. Other forms, particularly those accounting for non-linear weather effects, are rarely taken into account; exceptions are Odening et al. (2007), Schlenker and Roberts (2009) and Lobell et al. (2014).

SA2.1.2 Literature overview tables

In order to organize the vast literature on non-experimental cereal yield data (or output, excluding rice) in a convenient way, we produced overview tables which allow to compare several articles at once. A meta-table (S2.1) summarizes information from the overview tables (S2.3–S2.5). The order within the tables is chronological.

TABLE S2.1: Meta analysis of literature review

Characteristic	Absolute number	Share [%]
Total number of studies	37	100
Species maize	24	65
Species wheat	19	51
Variabe class evapotranspiration	8	22
Variabe class soil moisture	8	22
Aggregation over time: phenological	4	11
Total number of studies since 2010	16	100
Using production function approach	3	19
Using economic variables	5	31

Note: Source: own representation. Not all categories are exclusive. E.g., some studies analyse both maize and wheat, which explains why “maize” and “wheat” sum up to more than 100 percent.

Particularly, the row “PFA” (production function approach) addresses whether inputs are included. The criterion for being classified as “PFA: yes” was: The empirical model includes at least one variable measuring or approximating inputs, and can be classified as one category of the KLEMS approach (capital, labor, energy, material inputs and services; Coelli et al., 2005, 141) or as land. Furthermore, for each article, the tables provide characteristics such as time period, geographical focus, information on data aggregation and detrending, estimation technique, and included weather variables. In specific, several variables are used to quantify the process of evapotranspiration. These variables are collapsed into the group of variables called “e.” Similarly, variables approximating the water-holding capacity of soil are grouped into the class “sm.”

TABLE S2.2: Abbreviations literature overview tables

	meaning		meaning		meaning
agg	aggregation	im	implicit	qu	quadratic
AI	Ångström-Index	it	interaction terms	r	variables which measure
avg	average	irr	irrigated		or proxy radiation
cbnr	considered but not relevant	IV	instrumental variable	RCM	random coefficents
cc	cloud cover	lfa	less-favoured area		model
clr	crop land relevant	loc reg	local regression	RDI	Rainfall Deficit Index
cor	correlation	lsdv	least squares	re	random effects
Cobb-D.	Cobb-Douglas		dummy variable	reg	regional
CR	cumulative rainfall index	MARS	multivariate adaptive	rel	relative
cs	cross section		regression splines	RESET	regression specification
CV	coefficient of variation	max	maximum		error test
d	dummy variables	min	minimum	RSI	Rainfall Sum Index
DD	degree days	ml	maximum likelihood	Sept	September
DDD	damaging degree days	MPSMD	maximum potential	sh	sunshine hours
dev	deviation		soil moisture deficit	sm	variables which measure
DTR	diurnal temperature range	NAO	North Atlantic Oscillation		or proxy soil moisture
dyn	dynamic	NA	information not available	sp	spring
e(pot)	(potential) evapo-	nat	national	sqrt	squareroot
	transpiration	nls	non-linear least squares	sr	solar radiation
ex	explicit	Nov	November	SSA	Sub-Saharan Africa
fc	field capacity	NUTS	nomenclature of units	SWAP	soil water available
fd	first differencing		for territorial statistics *		to plants capacity
ff	functional form	ols	ordinary least squares	t	variables which measure
FGLS	feasible generalised	org	organic farming		temperature or related
	least squares	p	precipitation		variables (GDD, GSL)
GDD	growing degree days	PDSI	Palmer Drought	tfe	time fixed effects
gr	global radiation		Severity Index	transc	transcendental
gs	growing season	PFA	article follows a prod-	ts	time series
	(fixed months)		uction function approach	v	variable(s)
GSL	growing season length	pheno	phenological growth	VPD	vapor pressure deficit
GWR	geographically weighted		stages acknowledged	w	winter
	regression	pl	piecewise linear	ws	weather station
HDD	extreme heat degree days	pu	political units	wtd	within transformed
hm	hydrometeorological	prov	provinces		data
ife	individual fixed effects	psp	pre-season precipitation		

Note: Source: own representation. * European Commission (2007, 4–6).

TABLE S2.3: Literature review – main facts

Authors	Region(s) / Period(s) / Species	Dependent v. / PFA; land / econ.; inst. v.
Gornott/Wechsung (2016)	Germany: counties / 1991–2010 / silage maize, wheat	fd log(yield) / yes / yes (nat); land
Miao et al. (2016)	US: 1,144 counties / 1977–2007 / maize	yield / no / yes; d
Butler/Huybers (2015)	U.S.: counties of 17 states / 1981–2012 / maize	yield / no / no; no
Ray et al. (2015)	World: 13,500 pu / 1979–2008 / maize, wheat	yield / no / no; no
Cai et al. (2014)	U.S.: 958 counties / 2002–6 / maize	(1) yield; (2) 5-year-yield / no / no; tfe
Lobell et al. (2014)	U.S.: farms of 9 states / 1995–2012 / maize	yield / no / no; no
Ward et al. (2014)	SSA: grid (2,653 cells) / 1997–2003 / cereals	yield / yes; no / yes; no
Brown (2013)	Scotland / 1963–2005 / barley, oats, wheat	yield / no / no; no
Ortiz-Bobea/Just (2013)	U.S.: counties of 8 states / 1985–2005 / maize	yield / no / no; no
Osborne/Wheeler (2013)	World: 14 countries / 1961–2009 / maize, wheat	yield % changes / no / no; no
Roberts et al. (2013)	Illinois / 1950–2010 / maize	yield / no / no; no
Blanc (2012)	SSA: 37 countries / 1961-2002 / maize, millet	fd log(yield) / yes / yes; d, land
Heimfarth et al. (2012)	Germany: 22 farms / 1997–2009 / w wheat	yield / no / no; no
Lobell et al. (2011)	World: countries / 1980–2008 / maize, wheat	log(yield) / no / no; no
Li et al. (2010)	China: (1) 27 prov; (2): grid / 1978–2000 / sp, w wheat	yield / no / no; no
Schlenker/Lobell (2010)	SSA: countries / 1961–2002 / maize, sorghum, millet	log(yield) / no / yes; no
Schlenker/Roberts (2009)	U.S.: counties of 30 states / 1950–2005 / maize	log(yield) / no / no; no
You et al. (2009)	China: 22 prov / 1979–2000 / soft wheat	yield / yes; no / yes; d
Lobell et al. (2008)	12 regions / 1961–2002 / maize, sorghum, wheat	fd of yield; fd of output / no / no; no
Deschênes/Greenstone (2007)	U.S.: counties / 1987, 92, 97, 2002 / maize	yield / no / yes; no
Hussain/Mudasser (2007)	Pakistan: 2 regions / 1975/6–99/2000 / w wheat	log(yield) / no / no; no
Lobell (2007)	10 countries / 1961–2002 / maize, wheat	fd log(yield); cbnr: yield / no / no; no
Odening et al. (2007)	Northeast Germany / 1993–2005 / w wheat	yield / no / no; no
Reidsma et al. (2007)	51,843 farms of EU-15 / 1990, 1995, 2000 / barley, maize, wheat	yield / yes / yes; lfa, land
Isik/Devadoss (2006)	Idaho: 4 crop districts / 1939–2001 / sp barley and wheat	yield / no / no; no
Schlenker/Roberts (2006)	U.S.: 1839 counties of 24 states / 1950–2004 / maize	log(yield) / no / no; no
Bakker et al. (2005)	several EU states: NUTS-2 or 3 / 1970s–2000 / soft wheat	avg-yield; trend yield / no / yes; no
Lobell et al. (2005)	Mexico: 2 regions / 1988–2002 / durum, w wheat	fd of yield / no / no; no
Chen et al. (2004)	U.S.: 48 states / 1973–97 / maize, sorghum, w wheat	yield; log(yield) / yes / yes; land
Lobell/Asner (2003)	U.S.: 618 counties / 1982–98 / maize	trend yield / no / no; no
Carter/Zhang (1998)	China: 249 counties (5 hm-regions) / 1980, 85, 87–90 / cereals	yield / yes; no / yes; no
Kaufmann/Snell (1997)	U.S.: 78 counties of 8 states / 1969, 74, 78, 82, 87 / maize	yield / yes / yes; loan rates, land
Nicholls (1997)	Australia / 1952–92 / wheat (unspecified)	yield / no / no; no
Dixon et al. (1994)	Illinois: 9 regions / 1953–1990 / maize	yield / yes; no / yes; no
Kaylen et al. (1992)	U.S.: 10 regions / 1949–88 / maize	yield / no / yes; no
Yang et al. (1992)	North Dakota / 1929–88 / wheat: sp, durum	yield / yes / yes; land
Hansen (1991)	U.S.: 3,057 firms of 10 states / 1988–9 / maize	yield / yes; no / yes; no

Note: Source: own representation. Abbreviations according to table S2.2.

TABLE S2.4: Literature review – econometric details

Authors	Data structure / Detrending / Estimator	FF / RESET / Validation
Gornott/Wechsung (2016)	ts, panel / fd / ols; ml; RCM	Cobb-D / yes / yes
Miao et al. (2016)	panel / im (qu) / IV; ife	linear, qu, square root / no / no
Butler/Huybers (2015)	panel / im (linear) / ols	linear / no / no
Ray et al. (2015)	ts / ex (linear – cubic) / ols	qu; it / no / yes
Cai et al. (2014)	(1) panel; (2) cs / im (linear) or tfe / loc reg (GWR)	linear / no / yes
Lobell et al. (2014)	panel; seperate cs / im; ex / MARS	pl / no / yes
Ward et al. (2014)	cs / not relevant / (Non-) spatial Heckit; ols	qu / no / no
Brown (2013)	ts / fd / cor	linear / no / no
Ortiz-Bobea/Just (2013)	panel / im (qu) / lsdv (ife) or ols on wtd	linear, qu / no / no
Osborne/Wheeler (2013)	ts / percentage changes, fd / ols	linear / no / no
Roberts et al. (2013)	ts / ex (linear) / ols	qu; it / no / yes
Blanc (2012)	panel / fd / lsdv (ife and tfe)	qu, log; it, / no / yes
Heimfarth et al. (2012)	separate ts / ex (linear) / ols (cor)	linear / no / no
Lobell et al. (2011)	panel / im (country, qu) / lsdv (ife)	qu / no / no
Li et al. (2010)	separate ts / ex (linear) and fd / ols (cor)	linear / no / no
Schlenker/Lobell (2010)	panel / im (qu) / lsdv (ife) or ols on wtd	linear, qu, pl / no / no
Schlenker/Roberts (2009)	panel, cs, ts / im (qu) / lsdv (ife); cbnr: (tfe)	step function, pl, polynomial, qu / no / yes
You et al. (2009)	panel / im (linear, cbnr: qu) / ols (reg d)	Cobb-D / yes / no
Lobell et al. (2008)	ts / fd all v / ols	linear / no / no
Deschênes/Greenstone (2007)	panel / im (tfe) / lsdv (ife); state by year tfe	qu; it / no / no
Hussain/Mudasser (2007)	panel / im (linear) / lsdv (ife)	transc [cbnr: Generalised Cobb-D, qu] / no / no
Lobell (2007)	ts / fd [cbnr: spline trend] / ols	linear / no / no
Odening et al. (2007)	ts / NA / ols	Leontief [cbnr: logarithmic, qu] / no / no
Reidsma et al. (2007)	cs, multilevel / not relevant / ml, RCM	linear, qu / no / no
Isik/Devadoss (2006)	panel / im (linear) / ml	qu / no / no
Schlenker/Roberts (2006)	panel / im (linear) / lsdv (ife) or ols on wtd	polynomial, qu / no / no
Bakker et al. (2005)	cs / not relevant / ols	linear / no / yes
Lobell et al. (2005)	separate ts / fd all v / ols	linear / no / no
Chen et al. (2004)	panel / im (linear); ex (stochastic) / ml, (re)	linear, Cobb-D / no / no
Lobell/Asner (2003)	cs / NA / ols; cor	linear / no / no
Carter/Zhang (1998)	5 panels / NA / FGLS (including reg d)	Cobb-D / no / no
Kaufmann/Snell (1997)	pooled cs / im (linear) / ols	logarithmic, linear, qu / no / no
Nicholls (1997)	ts / fd / ols	linear / no / no
Dixon et al. (1994)	panel / NA / FGLS	linear / yes / yes
Kaylen et al. (1992)	ts / stochastic / state space model	linear; weather v qu / no / no
Yang et al. (1992)	ts / im (cubic) / ols	linear, qu / no / no
Hansen (1991)	cs / not relevant / ml (tobit model)	qu, log; it [cbnr.: Cobb-D] / no / no

Note: Source: own representation. Abbreviations according to table S2.2.

TABLE S2.5: Literature review – details weather

Authors	Weather: p, t, r / e / sm / other	agg. time / space (finer scale than region applied)
Gornott/Wechsung (2016)	p, t, r: t normalized sr / e: epot / no / no	gs: vegetative, reproductive / no
Miao et al. (2016)	p, t: G(H)DD, dev / no / no / p: pre gs	gs (partly): monthly / grid
Butler/Huybers (2015)	p, r: cbnr: sh; t: max, min, G(K)DD / no / no / cbnr: freezing days	pheno.: 4 stages / no
Ray et al. (2015)	p, t / no / no / no	gs (monthly avg) / harvested area grid
Cai et al. (2014)	p, t / no / no / no	gs for U.S. (avg of monthly avg) / no
Lobell et al. (2014)	p, t: min, max / e: VPD / sm: cbnr: PDSI / no	dyn gs (sowing), 30-day-avg / grid
Ward et al. (2014)	p: avg, CV; t: avg, DTR / no / no / no	gs: monthly avg / no
Brown (2013)	p, t: DD, r: sh / e: MPSMD / sm: as e / NAO	monthly, annual, w, summer / clr grid
Ortiz-Bobea/Just (2013)	p, t: G(D)DD / no / no / no	pheno.: 3 stages / clr grid
Osborne/Wheeler (2013)	p, t / no / no / no	dyn gs (crop calendar) / clr grid
Roberts et al. (2013)	p, t: G(H)DD / e: VPD / no / no	dyn gs (t) / clr grid
Blanc (2012)	p: SPI, flood, t: t, drought / e: Hargreaves / no / CO_2	calendar year / clr grid
Heimfarth et al. (2012)	p: CR / no / no / no	gs (partly): daily avg / ws max 200 km from farms
Lobell et al. (2011)	p, t: avg, min, max / no / no / no	dyn gs (crop calendar), avg of monthly avg / clr grid
Li et al. (2010)	separately: p; t / no / no / no	gs for China / no
Schlenker/Lobell (2010)	p, t: avg, G(D)DD / no / no / no	gs: see Lobell et al. (2008); / clr grid
Schlenker/Roberts (2009)	p, t (includes GDD) / no / no / no	gs: one for U.S.: / clr grid
You et al. (2009)	p, t, r: cc / no / no / no	gs on provincal level monthly avg / grid
Lobell et al. (2008)	p, t / no / no / no	gs: avg of monthly avg / clr grid
Deschênes/Greenstone (2007)	p, t: GDD / no / sm: moisture capacity / v soil characteristics	dyn gs (t) / grid
Hussain/Mudasser (2007)	p, t: GSL / no / no / no	gs: (fixed months, t determined) / no
Lobell (2007)	p, t: avg, DTR / no / no / no	gs: avg of monthly avg / clr grid
Odening et al. (2007)	p: RDI, RSI / no / no / no	gs (partly): daily avg / no
Reidsma et al. (2007)	p, t / no / no / no	gs: avg of monthly avg / grid
Isik/Devadoss (2006)	p, t / no / no / no	gs (partly): monthly avg, annual / no
Schlenker/Roberts (2006)	p, t / no / no / no	dyn gs (t) / clr grid
Bakker et al. (2005)	p, t, r: gr / e: epot / sm: soil depth, SWAP / no	calendar year / sm: grid
Lobell et al. (2005)	t: min, max, r: sr / no / no / no	gs (partly): daily avg / no
Chen et al. (2004)	p, t / no / no / no	gs (partly) / no
Lobell/Asner (2003)	separately: trend of p (t), r: sr / no / no / no	gs (partly) for U.S. / no
Carter/Zhang (1998)	/ e: AI / no / dev from sample mean	monthly / no
Kaufmann/Snell (1997)	p: avg, max stage 6; t: min, daily max / no / no / length stage 1, 2	pheno.: 8 stages / no
Nicholls (1997)	p, t: min, max / no / no / cbnr: CO_2	calendar year / no
Dixon et al. (1994)	p, t, r: sr / no / sm, / no	pheno.: 4 stages / no
Kaylen et al. (1992)	p, t: monthly avg / no / sm: pre-gs p / no	gs: monthly and divisional avg / area and clr weighted
Yang et al. (1992)	t, / no / sm: avg of p and psp / no	gs (partly) / no
Hansen (1991)	p, t (all monthly, 30-year-avg) / e: it p t / sm: cbnr: fc / erosion	gs (partly) / 83 weather districts

Note: Source: own representation. Abbreviations according to table S2.2.

SA2.2 Data

We provide further details on the different FADN data (TF-8 vs. TF-14), yield data and data sources for the variable *share spring wheat*, details on input and weather data (used deflators, data aggregation). Furthermore, we include additional formal weather variable definitions, explain how we merge different data sets and provide background information on the CAP reforms.

SA2.2.1 TF-8 versus TF-14

Alternatively, we considered to analyze farm type 13 "Specialist cereals, oilseed and protein crops (COP) (European Commission, 2010, 49)" (TF-14 grouping). One could argue that the latter type of farm would provide more precise data on inputs applied in wheat production in Germany. Higher aggregated data as from the TF-8 type 1 "Specialist field crops" are influenced to a higher degree by the use of inputs in production of other crops than winter wheat. However, we opted for the much higher representativeness of the TF-8 data. TF-8 is considerably closer to actual national output (see Figure S2.1; data from Statistisches Bundesamt, 2014a; other data sources as in paper. Represented output is calculated by multiplying average output of wheat SE110N with number of represented farms SYS02.). The importance of wheat within these groups does not differ much though. For 1995–2009, the mean shares of wheat in utilized land (without fallows and set-aside) are 0.3079 (TF8) and 0.3373 (TF14). A remaining limitation which applies to both groupings is that the data do not allow to determine exactly the amount of inputs applied in wheat production.

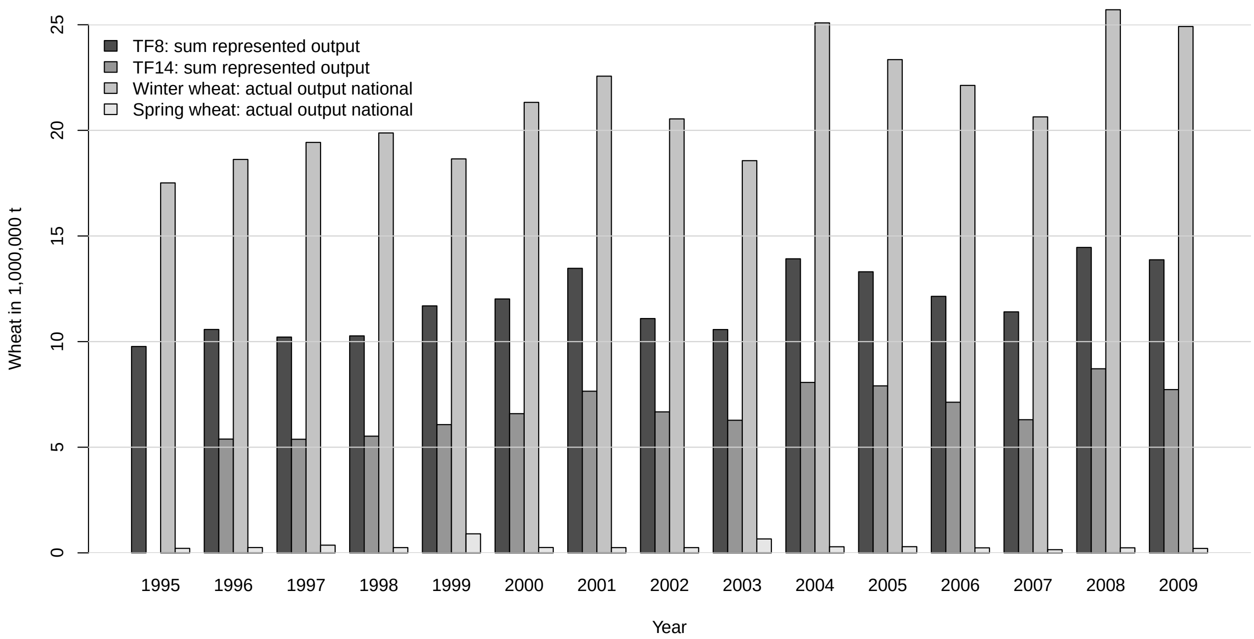

FIGURE S2.1: Comparison of output represented by TF-8 and TF-14 grouping and actual national level output

SA2.2.2 Yields

We cannot rule out that wheat yields include production of spring wheat and spelt (European Commission, 2015b). However, the contained share of spring wheat in total hectares of wheat can be regarded as low if the considered farms do not depart from the federal state trends. Spelt is of even lower acreage than spring wheat.

It can be argued that spring wheat might react differently to weather and application of inputs and that this might affect the results. For example, Chmielewski and Köhn (2000, 260) find different effects of weather on winter rye compared to spring cereals. Agricultural statistics show that the share of spring wheat in total wheat hectares (winter and spring wheat) was ca. 2% for 1995–2009, with a minimum of 0.3% and a maximum of 8.6%. We calculated these shares for each federal state and year using data from the German agricultural statistics (BMEL, 2015, various years). We included this variable as robustness check in the empirical model in order to account for regional differences and changes in spring wheat cultivation. The exact data sources are: BMEL (1995; 1996; 1997; 1998; 1999), BMVEL (2000; 2001; 2002), BMELV (2005; 2008; 2009).

Data on acreage of spelt are neither available from German agricultural statistics nor from Statistisches Bundesamt nor from the federal state divisions of the latter, Statistische Ämter des Bundes und der Länder. A German non-scientific journal specializing in agriculture reports 22,833 hectares spelt harvested in Germany in 2004; most of the spelt is harvested in Bavaria and Baden-Wuerttemberg (Agrarzeitung online, 2004). For comparison: winter wheat amounted to 3,046,000 hectares and spring wheat to 48,300 hectares (sum of acreage for German federal states in 2004).

We refrained from using specified winter wheat yield data which are available from German agricultural statistics instead of the FADN yields. This would result is a loss of information of which yields match which inputs and other farm characteristics. In addition, we would not have been able to determine the threshold of altitude as discussed below.

SA2.2.3 Details on input and weather data

We apply indices (base year 2010) without value-added tax (VAT) for the accounting year July until June (European Commission, 2010, 58; Statistisches Bundesamt, 2014b; 2014c, 4, 16). VAT is not considered because all data are recorded without VAT (definition variable SE395, European Commission, 2015b). Used price indices differ by input to be as precise as possible (Coelli et al., 2005, 155).

Capital: We approximate capital service flow, which would be the optimal measure according to Coelli et al. (2005, 144–151) by aggregating two variables. (1) Depreciation based on replacement value (SE360). Used deflator: equipment (contains machinery, tractors, buildings) and services for agricultural investment (series: *Waren und Dienstleistungen landwirtschaftlicher Investitionen*). (2) Machinery and building current costs (SE340); used deflator: price index based on price indices for maintenance of both machinery and buildings (series: *Instandhaltung von Maschinen und Material*; *Instandhaltung von Bauten*); applied weights (ca. 0.75 and 0.25, respectively) deduced from original weights in the complete index (Statistisches Bundesamt, 2014c, p. 10, table 3).

Labor: total labor input in hours (SE011).
Land: for wheat in hectares (SE110D).
Energy: SE345, contains "motor fuels and lubricants, electricity, heating fuels (European Commission, 2015b);" used deflator: price index for energy and lubricants (series: *Energie und Schmierstoffe zusammen*).
Material inputs: aggregated from 2 variables. (1) Fertilizer (SE295), used deflator: price index all fertilizers (series: *Düngemittel zusammen*); (2) Crop protection (SE300), used deflator: price index of all types of crop protection (series: *Pflanzenschutzmittel zusammen*).
Seeds: SE285, used deflator: price index of seeds and plants (series: *Saat- und Pflanzgut*).
Services: approximated with contract work (SE350): "Costs linked to work carried out by contractors and to the hire of machinery (European Commission, 2015b)." There is no ready to use deflator; we applied equally weighted price indices of maintenance of machinery and fuel (series: *Instandhaltung von Maschinen und Material*, *Treibstoffe zusammen*).
Manure: approximated by total livestock units (SE080).
Potential evapotranspiration: according to Turc-Ivanov (ETP_{TI}) calculated following Conradt et al. (2013, pp. 2950–2951, equation 1). ETP_{TI} considers average temperature, solar radiation, relative humidity (in %) and treats temperatures below and above 5°C differently. This measure has been superior over the evapotranspiration measure according to Haude (Haude, 1955).
Flood 2002: The Oder river flood affected several federal states: Bavaria, Brandenburg, Lower Saxony, Mecklenburg-Western Pomerania, Saxony, Saxony-Anhalt, Schleswig-Holstein, and Thuringia.
Other crop specific costs (SE305) are not considered, because they are a mixture of different costs which seem hardly relevant for this analysis. These include "[...] soil analysis, purchase of standing crops, renting crop land for a period of less than one year, purchase of crop products (grapes, etc.), costs incurred in the market preparation, storage, marketing of crops, etc. (European Commission, 2015b)."

SA2.2.4 Weather: merging data sets and additional variable definitions

The daily weather variables (e.g., precipitation) are available for each weather station separately. Complex variables that are not observed, such as potential evapotranspiration, are calculated for each weather station using the observed weather variables (see following formulas). We then aggregate these daily weather variables (station level) to daily federal state averages.

The phenological data (not the same stations) are aggregated to federal state averages, that is, phenological stages differ by federal state and year. We then determine the days that are part of one of the four phenological periods as described in table 2.2 in the paper. Finally, the daily weather variables (federal state level) are summed up for each of the phenological stages (federal state level).

Additional variable definitions

Days without precipitation (DWP):

$$DWP_p = \sum_{d=1}^{D} dwp_d = \begin{cases} 1, \text{ if } PREC_d = 0 \\ 0, \text{ if } PREC_d > 0 \end{cases} \tag{S2.1}$$

with $PREC_d$ denoting the daily precipitation level and subscript d denoting a day within a phenological period p as defined in the paper (table 1).

Growing degree days (GDD):

$$GDD_p = \sum_{d=1}^{D} gdd_d = \begin{cases} Temp_{\text{opt}} - Temp_{\text{min}} & \text{if } Temp_{\text{avg},d} > Temp_{\text{opt}} \\ Temp_{\text{avg},d} - Temp_{\text{min}} & \text{if } Temp_{\text{min}} < Temp_{\text{avg},d} \leq Temp_{\text{opt}} \\ 0 & \text{if } Temp_{\text{avg},d} \leq Temp_{\text{min}} \end{cases} \tag{S2.2}$$

with $Temp_{\text{opt}} = 20°\text{C}$ and $Temp_{\text{min}} = 4°\text{C}$. All temperatures refer to the daily *average* temperature ($Temp_{\text{avg},d}$).

Killing degree days (KDD):

$$KDD_p = \sum_{d=1}^{D} kdd_d = \begin{cases} Temp_{\text{avg},d} - Temp_{\text{opt}} & \text{if } Temp_{\text{avg},d} > Temp_{\text{opt}} \\ 0 & \text{if } Temp_{\text{avg},d} \leq Temp_{\text{opt}} \end{cases} \tag{S2.3}$$

Potential evapotranspiration according to Haude (ETP_H) was calculated by the product of vapor pressure deficit (VPD) and an empirical correction factor, the Haude factor f_H (Haude, 1955). For f_H, we applied the factors for wheat used by Schrödter (1985). VPD_d was calculated using the Magnus formula (Sonntag, 1990) with the maximum ($Temp_{\text{max},d}$) and the minimum temperature ($Temp_{\text{min},d}$) instead of the dew point temperature (Castellvi et al., 1996; 1997).

$$ETP_{H_p} = \sum_{d=1}^{D} f_{H_d} \underbrace{6.11\left(e^{\left(\frac{17.269\ Temp_{\text{max},d}}{237.3+Temp_{\text{max},d}}\right)} - e^{\left(\frac{17.269\ Temp_{\text{min},d}}{237.3+Temp_{\text{min},d}}\right)}\right)}_{=VPD_d} \tag{S2.4}$$

Temperature normalized solar radiation: Similar as Gornott and Wechsung (2016, 92):

$$SRT_p = \sum_{d=1}^{D} \frac{SR_d}{Temp_{\text{avg},d} + 20}. \tag{S2.5}$$

SA2.2.5 Altitude of farms, weather stations and phenological observational units

Merging different data sets must acknowledge the altitudes of farms, weather stations, and phenological observational units, because climate depends on altitude. Unfortunately, the publicly available FADN data do not provide information on altitudes at which the farms operate in Germany. Matching such information with data on the altitude of weather stations would be superior. Because this is not feasible, we do not consider weather stations above 600 m above sea level. The reason is that weather stations above this threshold are not relevant for the type of farm for which we analyze FADN data. According to IEEP (2006, 9), classification as Less Favoured Area (LFA) (Mountain Areas) in Germany applies to altitudes above 600m (with additional slope condition; see also European Commission, 2008, 4). According to an analysis of FADN data for the years 2004/05, farms specializing in field crops (TF-8 group 1, European Commission, 2010, 52) are not represented in areas classified as LFA Mountain Areas in Germany (European Commission, 2008, 61). Likewise we do not consider phenological observational units above 600 m.

SA2.2.6 Details on CAP reforms

First, the Agenda 2000 reform, which was implemented in the same year, reduced product-specific price support and introduced area-based compensation payments. The payment level still depended on the planted crop. Biased incentives through still-coupled per-hectare payments may have led farmers to grow wheat as a major grain on marginal land. The mid-term review of the Agenda 2000 in 2003 (Fischler-reform; active in 2005) led to payments being fully de-coupled from production in Germany. This may have also fostered incentives to plant wheat on marginal land. The Health Check of the CAP reforms in 2008 led to advanced reduction of the price support, and compulsory set-aside was abolished.

SA2.3 Regression model

The complete analysis was carried out using the software R (R Development Core Team, 2016). In the following, we provide information on model building and selection, robustness checks, and specification tests. In addition, we present additional plots and the cross-validation investigating the robustness of the regression model.

SA2.3.1 Model building, selection, specification tests and robustness

The following table S2.6 shows the tested functional forms for the right-hand side of the model. The model with the right-hand side of the production function in quadratic form (that is, the inputs are in levels but the dependent variable is in logs) and weather in logs achieves the lowest Akaike Information Criterion (AIC).

TABLE S2.6: Results specification simplified models ordered by AIC

Production function	Specification weather	p-value RESET	AIC	R^2	k
Quadratic dv log	log	0.420	-393.970	0.835	49
Quadratic dv log	level	0.732	-389.550	0.839	53
Translog	level	0.300	-388.340	0.843	56
Translog	log	0.231	-371.600	0.846	66

Note: dv: dependent variable; k: number of parameters; RESET: regression specification error test.

Table S2.7 shows results of specification tests of the two models in table 3 of the paper.

TABLE S2.7: Specification Tests

Test	p-values model 1	p-values model 2
Breusch-Pagan	0.301	0.769
Wooldridge's FD test h0=FE	0.002	0.896
Wooldridge's FD test h0=FD	<0.001	<0.001
RESET	0.49	0.834

Note: Croissant and Millo (2008, 28–9) provide details on Wooldridge's FD test.

The following three tables present several robustness checks. Table S2.8 shows which statistically insignificant variables were dropped after model building and selection (see step 4b, Figure 2.2 in paper).

TABLE S2.8: Robustness Check: Dropping Statistically Insignificant Weather Variables

	(1) Final specification	(3) from stepwise model search	(4) Drop stat. insign. weather var.
Intercept	0.0719 (0.0116)***	0.0883 (0.0078)***	0.0746 (0.0123)***
Inputs			
Capital	0.0433 (0.0399)	0.0413 (0.0406)	0.0451 (0.0406)
Labor	−0.1953 (0.1743)	−0.1881 (0.1902)	−0.1906 (0.1772)
Energy	0.1227 (0.0765)	0.1462 (0.0832)*	0.1247 (0.0769)
Material inputs	0.0911 (0.0305)***	0.0957 (0.0313)***	0.0879 (0.0308)***
Seeds	−0.1419 (0.0495)***	−0.1396 (0.0342)***	−0.1421 (0.0484)***
Manure	−5.5401 (6.9399)	−10.0438 (8.3737)	−6.5750 (6.8994)
Land wheat	−0.0495 (0.0418)	−0.0403 (0.0389)	−0.0475 (0.0434)
Capital squared	0.0007 (0.0002)***	0.0008 (0.0002)***	0.0007 (0.0002)***
Labor squared	0.0165 (0.0095)*	0.0179 (0.0108)*	0.0168 (0.0095)*
Energy squared	−0.0114 (0.0020)***	−0.0112 (0.0024)***	−0.0118 (0.0020)***
Material inputs squared	−0.0011 (0.0002)***	−0.0013 (0.0002)***	−0.0012 (0.0002)***
Seeds squared	0.0041 (0.0011)***	0.0040 (0.0008)***	0.0038 (0.0010)***
Capital * labor	−0.0044 (0.0015)***	−0.0045 (0.0015)***	−0.0044 (0.0015)***
Capital * seeds	−0.0008 (0.0001)***	−0.0009 (0.0002)***	−0.0008 (0.0001)***
Labor * energy	0.0111 (0.0036)***	0.0105 (0.0038)***	0.0113 (0.0037)***
Seeds * manure	0.2884 (0.1611)*	0.3210 (0.1607)**	0.2810 (0.1669)*
Weather			
Prec. stage 1	0.0499 (0.0159)***	0.0733 (0.0227)***	0.0545 (0.0163)***
Prec. stage 1 * East	−0.1675 (0.0439)***	−0.1542 (0.0477)***	−0.1600 (0.0437)***
Prec. stage 1 squared * East	−0.9129 (0.1412)***	−0.7915 (0.1949)***	−0.8648 (0.1333)***
Pot. evapotranspiration stage 1	0.2251 (0.0512)***	0.3017 (0.0430)***	0.2413 (0.0551)***
Growing degree days stage 2	0.1136 (0.0313)***	0.1881 (0.0425)***	0.1571 (0.0428)***
Solar radiation stage 2	−0.0232 (0.0423)	−0.0030 (0.0350)	−0.0260 (0.0420)
Solar radiation stage 2 squared	0.4425 (0.1561)***	0.6280 (0.1943)***	0.4576 (0.1495)***
Prec. stage 2	−0.0190 (0.0067)***	−0.0314 (0.0177)*	−0.0350 (0.0147)**
Killing degree days stage 3	−0.0182 (0.0027)***	−0.0171 (0.0030)***	−0.0183 (0.0028)***
Prec. stage 4	−0.0394 (0.0164)**	−0.0403 (0.0107)***	−0.0319 (0.0137)**
Prec. stage 4 * East	0.0843 (0.0519)	0.0903 (0.0484)*	0.0840 (0.0510)
Prec. stage 4 squared * East	0.2616 (0.1615)	0.2941 (0.1409)**	0.2946 (0.1587)*
Additional weather variables			
Prec. stage 2 squared * East		0.1682 (0.1116)	
Max. Temp. stage 2		−0.1540 (0.0876)*	−0.1247 (0.0941)
Temp. normalized radiation stage 4		0.1122 (0.1363)	
Pot. evapotranspiration stage 4		−0.0388 (0.0965)	
R^2	0.8290	0.8355	0.8303
Adj. R^2	0.7679	0.7691	0.7677

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: first differences of logged wheat yield. Model 1 as in table 3 in paper. Controls and year dummy variables included but not shown. Weather in logs, inputs in levels. Coefficients/standard errors for inputs multiplied by 100. Explanatory variables first differenced; weather/inputs mean centered. SCC standard errors in () (Driscoll and Kraay 1998). Pot.: potential; prec.: precipitation; temp.: temperature.

TABLE S2.9: Robustness Check: Logged Inputs

	(1) Final specification	(5) log inputs
Intercept	0.0719 (0.0116)***	0.0744 (0.0123)***
Inputs		
Capital	0.0433 (0.0399)	0.3298 (0.1280)**
Labor	−0.1953 (0.1743)	−0.0797 (0.0631)
Energy	0.1227 (0.0765)	0.1216 (0.0744)
Material inputs	0.0911 (0.0305)***	0.1476 (0.0788)*
Seeds	−0.1419 (0.0495)***	−0.1275 (0.0430)***
Manure	−5.5401 (6.9399)	−0.0245 (0.0268)
Land wheat	−0.0495 (0.0418)	−0.0105 (0.0538)
Capital squared	0.0007 (0.0002)***	1.7877 (0.7489)**
Labor squared	0.0165 (0.0095)*	0.0359 (0.4125)
Energy squared	−0.0114 (0.0020)***	−3.9231 (0.7876)***
Material inputs squared	−0.0011 (0.0002)***	−0.7344 (0.1895)***
Seeds squared	0.0041 (0.0011)***	0.3143 (0.1695)*
Capital * labor	−0.0044 (0.0015)***	−0.9657 (0.5539)*
Capital * seeds	−0.0008 (0.0001)***	−0.3515 (0.0972)***
Labor * energy	0.0111 (0.0036)***	1.4662 (0.4706)***
Seeds * manure	0.2884 (0.1611)*	0.0963 (0.0845)
Weather		
Prec. stage 1	0.0499 (0.0159)***	0.0416 (0.0141)***
Prec. stage 1 * East	−0.1675 (0.0439)***	−0.1620 (0.0414)***
Prec. stage 1 squared * East	−0.9129 (0.1412)***	−0.7536 (0.1300)***
Pot. evapotranspiration stage 1	0.2251 (0.0512)***	0.2105 (0.0580)***
Growing degree days stage 2	0.1136 (0.0313)***	0.0923 (0.0423)**
Solar radiation stage 2	−0.0232 (0.0423)	−0.0152 (0.0539)
Solar radiation stage 2 squared	0.4425 (0.1561)***	0.1543 (0.1307)
Prec. stage 2	−0.0190 (0.0067)***	−0.0253 (0.0070)***
Killing degree days stage 3	−0.0182 (0.0027)***	−0.0155 (0.0032)***
Prec. stage 4	−0.0394 (0.0164)**	−0.0415 (0.0175)**
Prec. stage 4 * East	0.0843 (0.0519)	0.0716 (0.0525)
Prec. stage 4 squared * East	0.2616 (0.1615)	0.2484 (0.1835)
R^2	0.8290	0.8138
Adj. R^2	0.7679	0.7472

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: first differences of logged wheat yield. Model 1 as in table 3 in paper. Controls and year dummy variables included but not shown. Weather in logs, inputs in levels (model 1) or in logs (model 5). Coefficients/standard errors for model 1 multiplied by 100. Explanatory variables first differenced; weather/inputs mean centered. SCC standard errors in () (Driscoll and Kraay, 1998). Pot.: potential; prec.: precipitation.

Table S2.9 provides a version of the final model with logged inputs (model 5). Model 5 passes specification tests such as the regression specification error test (RESET). However, model fit criteria for model 5 are slightly weaker compared to model 1 (R^2 is 0.83 and AIC: -396 for model 1; R^2 is 0.81 and AIC: -381 for model 5).

Moreover, the Davidson-MacKinnon-J-test (Wooldridge, 2009, 305) rejects the null hypothesis that model 1 cannot add explanatory power to model 5 at the 1% level, but the reverse is not true (based on robust covariance matrix). That is, the J-test favors model 1 over model 5.

Finally, table S2.10 shows instrumental variable estimates. For illustration purposes, we rely on a simpler model where the dependent variable is still the difference of the logged yield but only the differenced material inputs enter on the right-hand side (and the dummy variable for the extreme observation in Brandenburg 2003). We use the second lagged difference of material inputs (linear and squared term) as instruments. These are exogenous to yields in year t.

First-stage regressions show a statistically significant correlation (based on robust standard errors) of the difference of material inputs (linear and squared terms) and the IVs (both R^2 are roughly 0.2, irrespective of whether the Brandenburg dummy variable is included). Hence, the used instruments could be treated as valid.

The results are shown in the column labeled model 6. While the results do not qualitatively change, that is, the signs of material inputs are preserved within the IV results, we find a modest quantitative difference between the marginal effects of material inputs in model 1 and the IV estimate. The marginal effect in the IV estimate is about 1 percentage point larger at a 30 EUR increase of material inputs (2.8% compared to 1.7% in model 1), but also has a smaller positive range (due to the more prominent quadratic term). This difference—in our judgment—does not query the results of model 1. While the linear term is not significant alone at conventional levels ($p = 0.15$; robust standard errors), both linear and squared term are jointly significant (F-test; $p < 0.001$; robust standard errors). For further robustness checks, we have also included capital, where the interpretability of the results suffered from considerable increases in the standard errors, which is a known problem of IV regression (Wooldridge, 2009, 511).

TABLE S2.10: Robustness Check: Instrumental Variable Estimates

	(1) Final specification	(6) IV: material inputs	(7) IV: material inputs + capital	(8) IV: logged material inputs
Intercept	0.0719 (0.0116)***	0.0094 (0.0209)	0.0861 (0.0945)	0.0074 (0.0195)
Capital	0.0433 (0.0399)		0.4865 (0.5843)	
Material inputs	0.0911 (0.0305)***	0.2192 (0.1519)	0.2265 (0.1829)	0.3630 (0.3123)
Material inputs squared	−0.0011 (0.0002)***	−0.0042 (0.0019)**	−0.0046 (0.0022)**	−3.7625 (0.8254)***
Year 2003 * Brandenburg	−0.3877 (0.0229)***	−0.5639 (0.0340)***	−0.5679 (0.0461)***	−0.6301 (0.0759)***
R^2	0.8290	0.2593	0.1409	0.2649

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: first differences of logged wheat yield. Model 1 as in table 3 in paper but other parameters not shown. Inputs in levels (except model 8). Coefficients/standard errors for models with inputs in levels multiplied by 100. Explanatory variables first differenced; weather/inputs mean centered. Used instruments: own second lagged difference. SCC (Driscoll and Kraay, 1998) standard errors in ().

SA2.3.2 Actual and fitted values

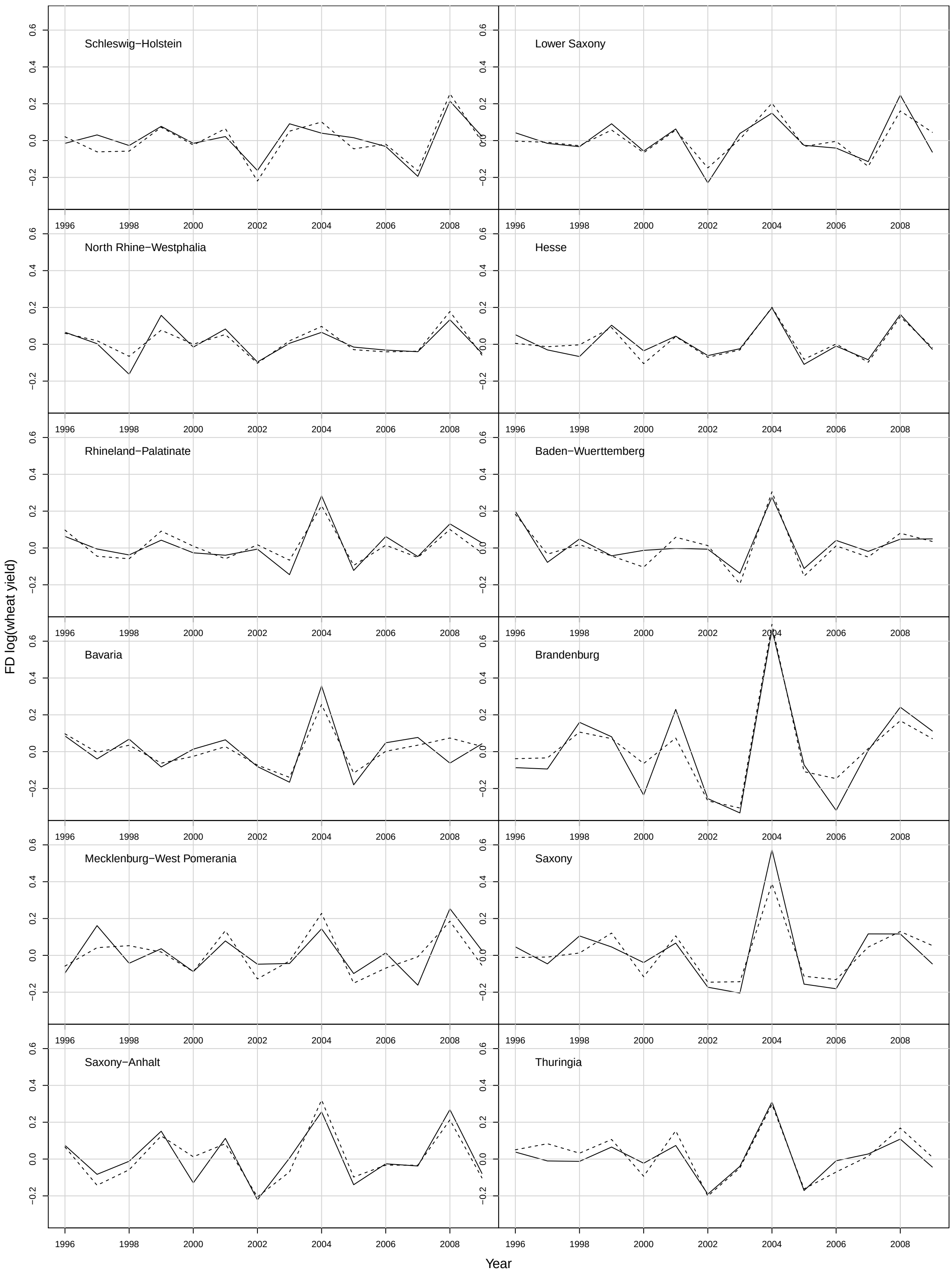

FIGURE S2.2: Actual (solid line) and fitted (dashed) yield changes based on model 1

SA2.3.3 Cross validation

The average root mean squared error for the predicted years is 0.11. The Pearson correlation between actual and the out-of-sample predicted values is 0.51 ($R^2 = 0.26$). Notably, the out-of-sample prediction errors (RMSE, mean absolute error) are higher for yield changes following the heatwave year. That is, the model's predictive power is considerably weaker for 2004 and the Federal State of Brandenburg, which considerably suffered from the heatwave (Figure S2.3). Within the prediction, we do not take time shocks into account and thus argue that the interaction of weather, inputs, and yield changes seems to be well-captured by the production function model.

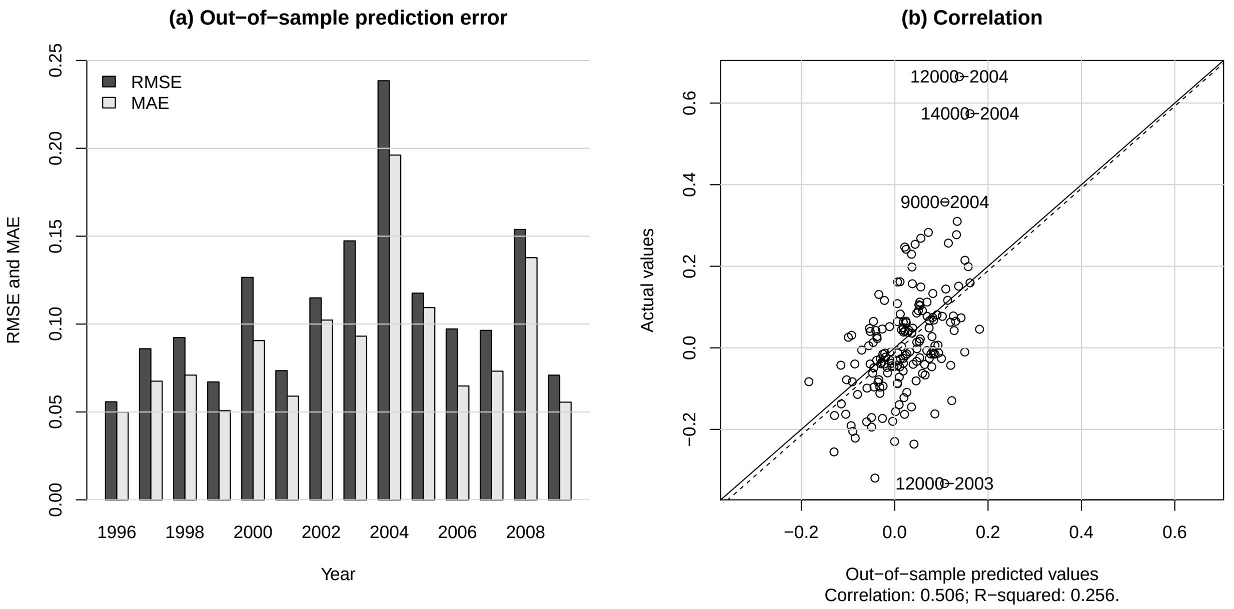

FIGURE S2.3: Cross-validation. (a) Prediction errors; (b) Actual and out-of-sample predicted values. Solid: perfect fit line; dashed: from linear regression

SA2.4 Yield volatility

The section contains the result table underlying the maps depicting the volatilities in the main paper and information on the robustness check regarding a possible aggregation bias.

SA2.4.1 Detailed results

TABLE S2.11: Actual, Weather- and Input-induced Yield Volatility [in %]: (1) 1996–2002 and (2) 2003–2009

	actual		weather		Volatility model 1 inputs		Volatility model 1 inputs + weather		Volatility model 2 weather	
Period	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
Western Germany										
Schleswig-Holstein (SH)	7.52	12.38	3.23	3.72	7.64	5.12	6.61	7.90	2.88	4.26
Lower Saxony (LS)	10.70	12.84	5.50	4.04	2.40	3.26	5.88	5.21	5.15	5.05
North Rhine-Westphalia (NRW)	10.94	6.66	4.62	5.38	7.57	3.34	6.39	7.65	4.23	6.49
Hesse (HE)	6.55	11.85	5.57	4.63	2.14	3.50	6.11	4.20	4.41	4.53
Rhineland-Palatinate (RP)	3.97	14.99	5.46	4.11	5.17	5.05	7.53	3.61	4.93	4.87
Baden-Wuerttemberg (BW)	8.99	13.66	4.08	3.90	7.10	3.49	7.52	5.56	3.69	3.74
Bavaria (BV)	7.20	18.24	4.42	2.93	6.64	3.80	3.77	5.51	3.75	2.45
Eastern Germany										
Brandenburg (BB)	18.97	34.58	5.74	7.08	6.31	4.81	9.19	8.56	7.34	8.23
Mecklenburg-West Pomerania (MWP)	9.54	14.23	3.57	2.73	6.22	4.00	6.50	4.93	3.09	2.59
Saxony (SY)	9.43	27.42	3.83	6.55	5.16	7.10	7.30	10.95	4.34	7.07
Saxony-Anhalt (SA)	13.67	16.17	5.57	6.82	5.80	8.48	10.34	7.38	6.59	7.47
Thuringia (TH)	8.92	15.09	4.58	2.75	5.90	5.41	9.71	5.65	4.15	3.72

Note: National level volatility based on yearly dummy variables: *model 1*: (1) 4.99, (2) 11.17; *model 2*: (1) 6.1, (2) 13.03.

SA2.4.2 Aggregation bias

Admittedly, spatial dependence in weather-related yield variability might not adhere to the administrative borders of federal states. In addition, the level of geographical aggregation differs by state; however, to acknowledge input changes as determinants of yield changes, a common level of aggregation for which all data are available must be applied. While we cannot circumvent different geographical aggregation, we must acknowledge that this might drive the results towards more heterogeneous volatilities across regions because in smaller geographical units the share of idiosyncratic factors in the variance might be higher due to the lower aggregation level. To investigate a possible aggregation bias, we run an auxiliary regression of weather-volatilities on hectares of wheat at the state level in each period. In the presence of aggregation bias, volatilities would decrease in the number of hectares planted with wheat. The estimated coefficient remains insignificant and we thus rule out any aggregation bias.

Chapter 3

Market Integration or Climate Change? Germany, 1650–1790

Authors: Hakon ALBERS*, Ulrich PFISTER† and Martin UEBELE‡

Note: This chapter is a revised version of the EHES Working Paper No. 135: Albers et al. (2018).

Abstract: In Malthusian economies, larger markets for grain held the potential for Smithian growth and 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, particularly against the background of the Little Ice Age. We show analytically how spatial arbitrage can be measured robust to shocks and that arbitrage must reduce aggregate volatility. A novel data set allows to track in detail the spatial pattern and chronology of grain market integration in the German lands.

Keywords: Market integration, price volatility, agriculture, weather, climate, coefficient of variation.

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3.1 Introduction

Market integration improves allocative efficiency of production factors via market access (Federico, 2012; Kelly, 1997). In addition, better integration of markets for agricultural products is

***Corresponding Author**. Martin Luther University Halle-Wittenberg, Department of Economics, Universitätsring 3, D-06108 Halle, Germany

†University of Münster, Institute of Economic and Social History, Germany.

‡Düsseldorf, Germany. Former address: University of Groningen, Netherlands.

likely to have improved the allocation of grain for consumption through more stable prices, because integrated markets are expected to have a lower price volatility (Chilosi et al., 2013; Jacks et al., 2011). This is particularly relevant in a Malthusian growth regime. Even if people survived food crises, the latter had serious consequences. Baten et al. (2014) show that grain price shocks harmed cognitive ability in England from 1780 to 1850, which had further negative effects on subsequent income generation.

Several studies have found price convergence for parts of pre-industrial Europe and have interpreted this as evidence for the onset of market integration well before the transport revolution of the 19th century (Bateman, 2011; Chilosi et al., 2013; Federico, 2011).[1] However, one potential alternative explanation for price convergence and/or declining price volatility are changes in the frequency and/or severity of weather shocks resulting from climate change.[2] A large part of the market integration literature analyzes grain prices, which can be highly influenced by weather conditions. 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 related to the waning of the Little Ice Age (LIA) since the early 18th 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. 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 explore this methodological problem by analyzing pre-industrial Germany, one of the most populous economies of 18th century Europe (Malanima, 2010) and create a novel grain price data set that sheds light on the dark spots of pre-railway integration. While Chilosi et al. (2013) reveal geographical clusters of integration within Europe, our data allow a more detailed analysis for Germany. Our stable sample of rye prices with less than 5% missing observations comprises 14 German cities in 1651 until 1790; our unbalanced sample includes 33 cities. The period from 1651 until 1790 defines the 140 years between two major historical events: the end of the Thirty Years' War and the War of the First Coalition following the French Revolution. We focus on rye, because it was the most important cereal used for bread, the main source of calories in Germany, which is also evident in cropping shares: 38% rye, 23% oats, 17% barley and 7% wheat (van Zanden, 1999, 368).[3] Additionally, data from the urban grain market in Cologne show that about 71.8 % of

[1]Apart from the law of one price, there are other theoretical approaches to study market integration based on land values (Donaldson and Hornbeck, 2016) or trade flows (Wolf, 2009). However, these data are usually not available for the pre-industrial era.

[2]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)."

[3]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); about 8% of consumed calories were provided by potatoes (Pfister, 2017, S2, p. 3).

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 gain a broader picture, we also included the cash crop wheat, barley and oats.

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.

We find cross-sectional price convergence at the national level and then use variance decomposition (Federico, 2011) to understand where this decrease in price gaps stems from. Prices converged by 0.4% per year (1651–1790) from a cross-sectional dispersion relative to the mean price of 31% (1651–75) to 14% (1766–90; based on the standard deviation of real five-year average prices). Our data set allows to reveal that price convergence mostly appeared in North-Western Germany and along major rivers (Elbe and Rhine). The importance of waterways for integration of bulk commodities in general has been emphasized by Chilosi et al. (2013) and for Germany in particular by Wolf (2009). Temporal price volatility declined from 25% (1651–75) to 14% (1766–90).

Our results are the first for the food crop rye and show much stronger market integration than previous studies that were based on the cash crop wheat (e.g., Bateman, 2011). We partly attribute this to increased data quality, which we achieve through a careful conversion of currencies. 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 show formally that the decline in aggregate price volatility must be at least partly a result of spatial arbitrage, which is reflected in price convergence.

We present our data in Section 3.2. We then explore how the CV is affected by weather shocks and climate change in Section 3.3 and develop a method to identify price convergence that is not contaminated by weather shocks or climate change. Section 3.4 presents our empirical results on price convergence at the national level and in North-Western Germany. In Section 3.5, we show that market integration deepened mainly along major rivers. One consequence of increased spatial arbitrage was a reduction in foodgrain volatility as shown in Section 3.6 but climate change likely contributed to this phenomenon. Section 3.7 concludes.

3.2 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. We include cities in our sample that were part of both the Holy Roman Empire in 1792 and

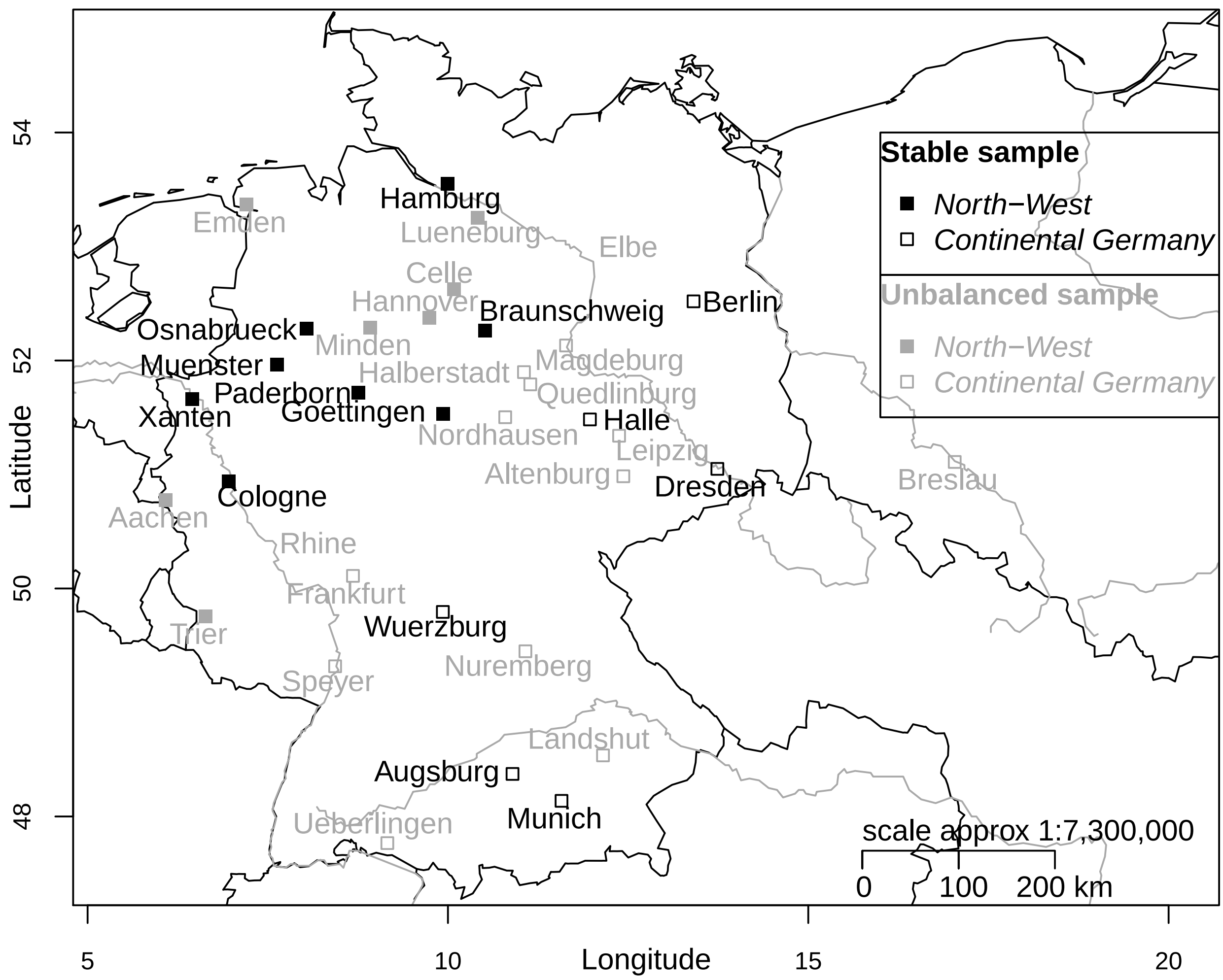

FIGURE 3.1: Cities in the sample. Stable sample 1651–1790: ≤5% missing observations per individual series. Cities are allocated to North-West if continentality≤ 17.0°C. Data on continentality from Müller-Westermeier et al. (2001, map 7), Deutscher Wetterdienst (2018) and WMO (2017a).

the German Empire of 1871 (for details see Fertig et al., 2018, 8). In what follows, we focus on a stable sample of calendar year rye prices for 14 cities in 1651–1790 (Figure 3.1, black) measured in grams of silver per liter (g Ag / l). The stable sample includes price series with ≤5% missing observations. Next to the stable sample, Figure 3.1 shows 19 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 3.3.3. 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 ≤ 17.0°C, a city is allocated to the North-Western sub-sample with a rather oceanic climate, otherwise to 'continental Germany' (see supplementary appendix SA3.7 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. We increased data quality by converting all prices to a common time base and using new series of silver content. Further information on data preparation and sources for each individual city and cereal are relegated to SA3.1–SA3.4. For the analysis of aggregate volatility, we calculate the aggregate real rye price as the arithmetic mean of the nominal city prices in our sample and deflate it with the national CPI by Pfister (2017; see SA3.6 for details). 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 the food crop rye.

We carried out Augmented Dickey Fuller (ADF) and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) tests in various specifications for all 14 prices series of the stable sample (all details reported in SA3.5.1). At the level of individual cities, the majority of real rye price series can be regarded as level or trend stationary.

3.3 Anatomy of the coefficient of variation

Stationarity of most grain prices in our sample for most of the time precludes the application of cointegration analysis to study market integration (see, e.g., Dobado-Gonzàlez et al., 2012, Jacks, 2004, Persson, 1999, 96–100 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). Section 3.3.1 analyzes shocks in the CV. Sections 3.3.2 and 3.3.3 discuss how to measure price convergence robust to weather shocks and climate change, respectively.

3.3.1 Shocks in the cross-sectional coefficient of variation

In Sections 3.3.1 and 3.3.1, we assume that a shock alters the price level in absolute terms. We analyze several cases of shocks depending on how these affect cross-sectional units (all: symmetric or only one: asymmetric) and on the presence of arbitrage (perfect or none). The representation for the absolute shock is simpler and hence, we start with this variant. In Section 3.3.1, we relax this assumption and discuss proportional shocks. This is necessary, because we cannot rule out either type of shock on empirical or theoretical grounds. Section 3.3.1 summarizes the main results.

Symmetric absolute shock

The CV is calculated as follows:

$$CV_t = \frac{\sqrt{\frac{1}{N-1}\sum_{i=1}^{N}(p_{it}-\bar{p}_t)^2}}{\bar{p}_t} \tag{3.1}$$

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:

$$\bar{p}_t^z = \frac{(p_{1t} + s_t) + ... + (p_{Nt} + s_t)}{N} = \frac{p_{1t} + ... + p_{Nt}}{N} + s_t = \bar{p}_t + s_t, \tag{3.2}$$

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:

$$\sum_{i=1}^{N}(p_{it} - \bar{p}_t^z)^2 = (p_{1t} + s_t - [\bar{p}_t + s_t])^2 + ... + (p_{Nt} + s_t - [\bar{p}_t + s_t])^2. \tag{3.3}$$

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

$$CV_t = \frac{\sqrt{\frac{1}{N-1}\sum_{i=1}^{N}(p_{it} - \bar{p}_t)^2}}{\bar{p}_t + s_t} \tag{3.4}$$

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.

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[4] 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}.

$$\bar{p}_t^z = \frac{(p_{1t} + s_{1t}) + ... + p_{Nt}}{N} = \frac{p_{1t} + ... + p_{Nt}}{N} + \frac{s_{1t}}{N} = \bar{p}_t + \frac{s_{1t}}{N}. \tag{3.5}$$

This shock affects the sum of squared deviations as follows:

$$\sum_{i=1}^{N}(p_{it} - \bar{p}_t^z)^2 = (p_{1t} + \frac{s_{1t}}{N} - [\bar{p}_t + \frac{s_{1t}}{N}])^2 + ... + (p_{Nt} + \frac{s_{1t}}{N} - [\bar{p}_t + \frac{s_{1t}}{N}])^2. \tag{3.6}$$

The local shock cancels. However, the CV is still affected and decreases as well (if $s_{1t} > 0$), but to a lesser extent than in the case of a symmetric shock ($\frac{s_{1t}}{N}$ in the denominator instead of 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. By contrast, the SD turns out as a measure of price dispersion that is robust to both symmetric and asymmetric absolute shocks with perfect arbitrage.

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

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:

$$\begin{aligned}\textstyle\sum_{i=1}^{N}(p_{it}-\bar{p}_t^z)^2 &= (p_{1t}+s_{1t}-[\bar{p}_t+\frac{s_{1t}}{N}])^2+...+(p_{Nt}-[\bar{p}_t+\frac{s_{1t}}{N}])^2 \\ &= (p_{1t}+(1-\frac{1}{N})s_{1t}-\bar{p}_t)^2+...+(p_{Nt}-[\bar{p}_t+\frac{s_{1t}}{N}])^2.\end{aligned} \tag{3.7}$$

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.

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, the denominator can amplify a decrease in the numerator, because now a smaller numerator is divided by a larger denominator. 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.

Whereas both CV and SD react to asymmetric absolute shocks without arbitrage, it turns out that the CV might amplify or dampen local shocks depending on the relative position of the shocked city's price with regard to the cross-sectional mean price.

Proportional shock

We summarize the main results of this sensitivity check and relegate all details to SA3.16, 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. Thus, the CV corrects for symmetric monetary inflation while the SD does not (Rönnbäck, 2009, 102).

Summary

Taken together, our formal analysis of the CV demonstrates that the CV is neither robust to spatially symmetric nor asymmetric shocks except for one case, namely, a *symmetric proportional shock with no arbitrage*. By contrast, the measurement of price dispersion using the SD reacts in only three out of seven cases to weather shocks (Table 3.1, cases (3), (5) and (7)). Furthermore, the analysis of proportional shocks shows that the SD should be calculated on the basis of prices that are corrected for inflation.

While earlier discussions of the CV (Federico, 2012, 479–80) assume that at least information about prices flows between markets, our framework allows to relax this assumption ('no arbitrage'). This is important in our view, because for the pre-industrial era, we cannot be sure about the existence of the exchange of information and trade.

TABLE 3.1: 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			
(1) Symmetric	↓↓	(.)	(.)
(2) Asymmetric, perfect arbitrage	↓	(.)	(.)
(3) Asymmetric, no arbitrage	↓↑	↓↑	(.)*
Proportional price changes due to weather shock			
(4) Symmetric, perfect arbitrage	↓↓	(.)	(.)
(5) Symmetric, no arbitrage	(.)	↑↑	(.)*
(6) Asymmetric, perfect arbitrage	↓↓	(.)	(.)
(7) 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. *: given the assumption that local shocks are nullified by averaging. Source: own representation.

3.3.2 Measuring price convergence robust to weather shocks

We exploit the lower susceptibility of the SD to shocks and can thereby better avoid misinterpretations of price convergence resulting from shocks as market integration. To deal with the remaining three cases in which weather shocks can lead to changes of the SD (Table 3.1, cases (3), (5) and (7)), we average the local 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. In addition, from an economic perspective, using five-year-averages brings prices closer to equilibrium than relying on annual averages.[5]

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

Note that 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.

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

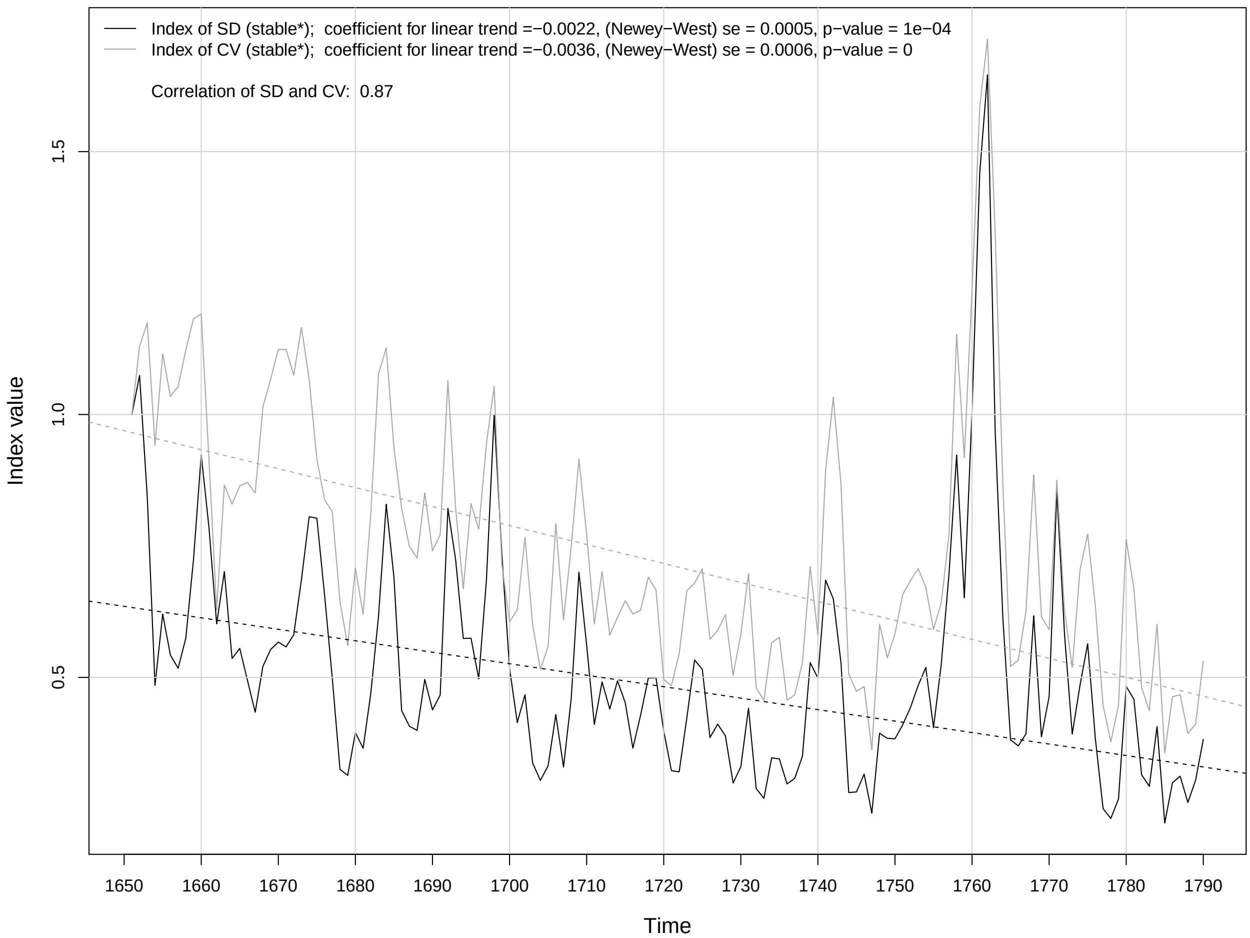

FIGURE 3.2: Inter-urban price dispersion: cross-sectional standard deviation and coefficient of variation of real rye prices Germany, stable sample (≤5% missing observations per individual series). Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Cross-sectional nominal price series are deflated with the national CPI from Pfister (2017).

The higher sensitivity of the CV to shocks leads to a larger trend estimate than the SD in our sample (Figure 3.2). To render SD and CV comparable, each series is calculated for exactly the same data (stable sample and deflated) and indexed on the year 1651. While the difference in the trend estimates is not very large (CV:−0.5%, SD: −0.45%, based on logged dependent variables), the series are remarkably different prior to 1720 given that the only difference is the division by the sample mean, the aggregate real rye price. This mean 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 SA3.8.2).

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 S8 in SA3.8.1). 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, 2014a). Additionally, the 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, 2014a, 1374; Xoplaki et al., 2005, 2).[6]

3.3.3 Measuring price convergence robust to climate change

Climate change involves two types of changes in climate variables: 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 the cross-sectional SD of five-year-mean prices in those three cases, where the SD in susceptible to shocks (Table 3.2, cases (3), (5), (7)).

TABLE 3.2: 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			
(1) Symmetric	as absolute price changes, symmetric weather shock		
(2) Asymmetric, perfect arbitrage	as absolute price changes, asymmetric weather shock		
(3) Asymmetric, no arbitrage	↓↑	↓↑	↓↑**
Proportional price changes due to climate change			
(4) Symmetric, perfect arbitrage	as proportional price changes, symmetric weather shock		
(5) Symmetric, no arbitrage	(.)	↑↑	↑↑*
(6) Asymmetric, perfect arbitrage	as proportional price changes, asymmetric weather shock		
(7) 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 sub-regions. Climate change refers to a change of the "mean state" and/or a change of the variability around the "mean state" (= persistent change of weather shocks). Source: own representation.

Climate is usually defined over a large area. Hence, climate change can be considered as a symmetric shock to our sample. We first discuss this case and then allow for regionally different impacts of climate change.

By construction, the SD is affected by symmetric shocks in only one case: *proportional symmetric shock with no arbitrage* (Table 3.2 case (5)). We now have to differentiate the two types of climate change introduced above, a changing mean state and a changing variability. First, assume that the

[6]Kelly and Ó Gráda (2014a) challenge the existence of the LIA based on a critique of defining the LIA based on MA-filtered temperature data. While we agree that spurious cycles are a problem, this is not sufficient to deny the existence of the LIA, because the evidence we report here is consistent with the standard WMO-definition of climate change.

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 SD 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 so that 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 LIA had begun to wane around 1700. This downward trend would appear even if we use five-year mean 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 SD and bias our results towards market integration.

We can approach both dimensions of climate change (change in mean state of climate; change in variability) by evaluating whether we observe any downward (upward) trend for the real price for the full sample (Table 3.2).

An additional issue is created by the possibility that climate change affects individual regions within a larger market area in different ways. Consequently, prices in a sub-region might 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) in two sub-regions.

More specifically, when focusing on the 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.

Additionally, price convergence in the full sample might 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, our analysis demonstrates 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. With regard to climate change, we evaluate trends in real prices to understand whether negative trends could bias our results towards price convergence.

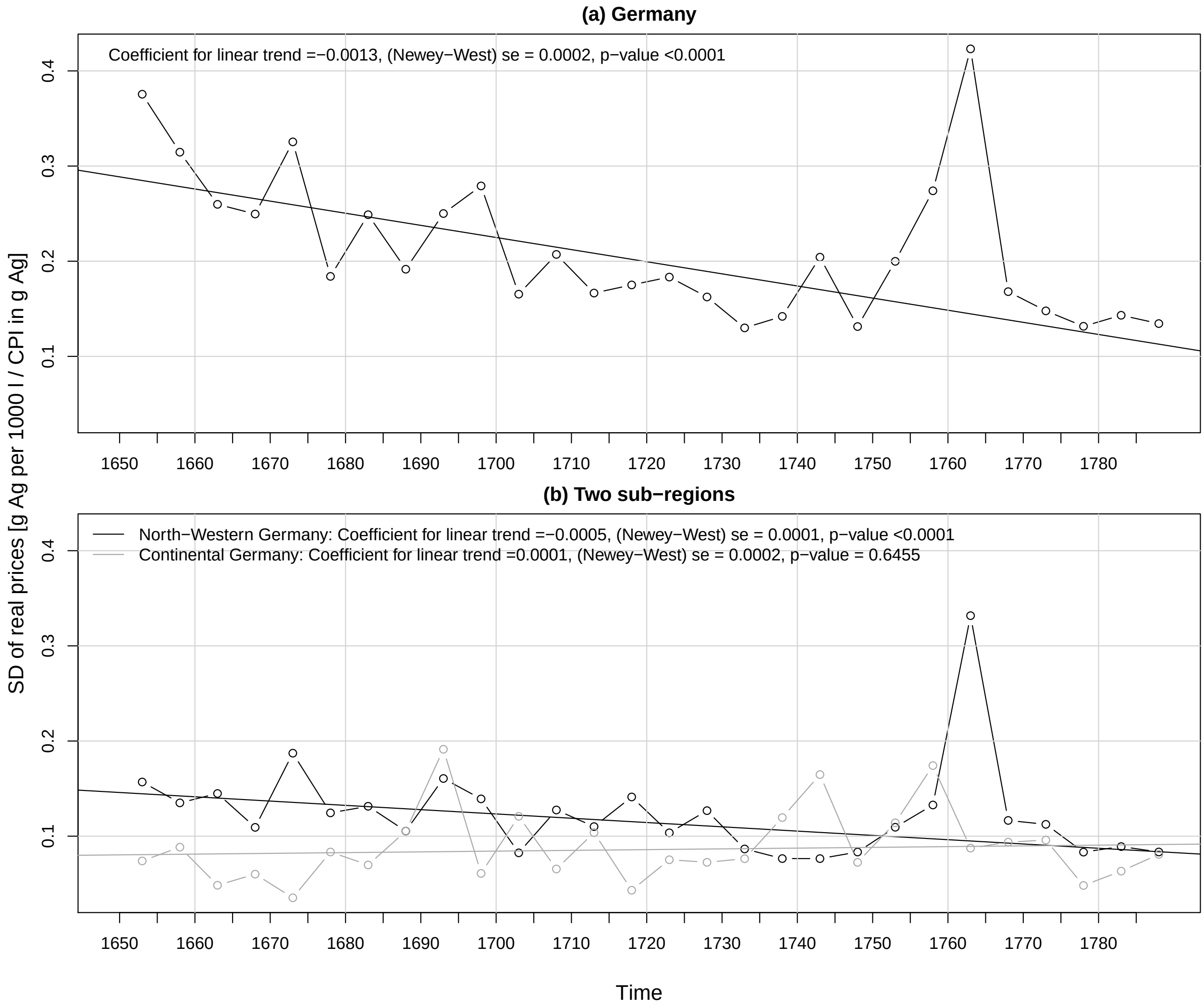

FIGURE 3.3: Inter-urban price dispersion 1651/5–1786/90. Cross-sectional standard deviation of real 5-year-mean-prices, rye (stable sample). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

3.4 Price convergence at the national level and in North-Western Germany

The cross-sectional SD based on five-year-averages of rye prices exhibits a clear downward trend (Figure 3.3, panel a). At the national level, prices converged by ca. 0.4% per year in 1651–1790 (0.4% $\approx$ parameter/average SD 1651–75 ($= -0.0013/0.3$).[7] In 1651–75, the SD was roughly 0.3 (CV$\approx$32%); this value declined to 0.15 (CV$\approx$14%) for the years 1766–90. This contrasts with earlier work by Bateman (2011, 459) who finds no strong trend of price convergence for Germany.

[7] The trend estimate with logged SD (CV) as dependent variable is 0.6% (0.64%) per year.

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 well 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.

But without further investigation, we cannot exclude effects of *asymmetric absolute or 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 3.3 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 rest of the sample (cases: asymmetric absolute or proportional changes; cf. Section 3.3.3). Although both regions might not have been integrated in one market, we would observe a reduction of the inter-urban price dispersion at the national level, because one sub-region experienced a downward trend of the real price, lowering 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, regionally specific climate change, we brake the total sample down into 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[8] remains robust within North-Western Germany but the trend weakens; convergence is not apparent in continental Germany (Figure 3.3 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; see also Persson, 1996, 711–2). 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 we are unable to adjust all individual price series for war-related inflation.

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 S3.11 in SA3.9.1). The

[8]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.

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 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 3.4 plots the between-SD, that is, the square root of the between variance of the two subregions. 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 3.3 b).

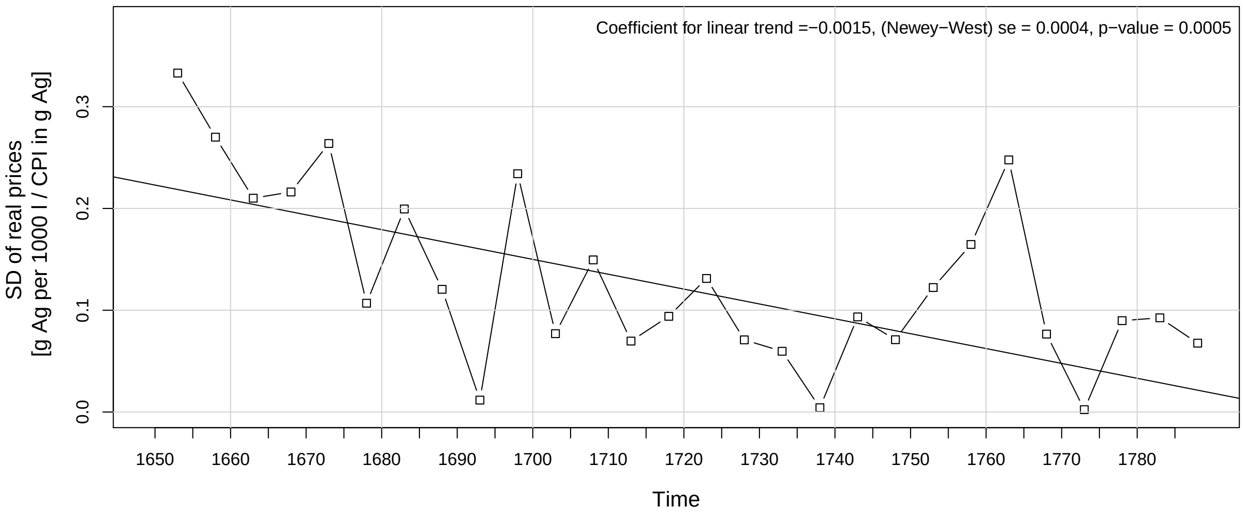

FIGURE 3.4: Inter-regional price dispersion 1651/5–1786/90. Between-region standard deviation (square root of variance between North-Western and continental Germany); real 5-year-mean-prices, rye (stable sample). Regression for linear trend includes a dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

As discussed in Section 3.3.3, a negative trend of the average price in North-Western Germany (potentially caused by warming, the end of the LIA) might lead to price convergence between North-Western and continental Germany 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 -0.0140 [trend parameter inter-regional price dispersion] $-(-0.0008)$ [trend parameter real rye price Northern Germany] $= 0$ ($p < 0.01$; Newey and West standard error).[9] Thus, the trend of the inter-regional dispersion remains substantially negative in magnitude, -0.0132, and is significantly different from zero. That is, inter-regional dispersion decreases by 1.32% per year net of price convergence due to the decreasing real price in North-Western Germany. This test is directly related to our formal analysis, because we show that, *without arbitrage*,

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

a *symmetric proportional shock* with a negative shock rate remains as a factor that decreases the SD (see SA3.16, eq. (S3.12)). A similar test reveals that the negative trend estimate for price dispersion within North-Western Germany remains significant with $p < 0.01$.

In short, our empirical result on price convergence is robust to weather shocks and climate change resulting in symmetric absolute price changes at the national level. The regional split in combination with statistical tests of trends in variances and mean prices based on the framework outlined in Section 3.3 allows to control for spatially different effects of climate change. Thus, our results on price convergence show market integration in 17th- and 18th-century Germany and cannot be explained by the end of the LIA. Markets integrated in North-Western Germany and already in the second half of the 17th century, that is, well before the transport revolution of the 19th century. Furthermore, convergence between the two regions accounts for an important fraction of the national convergence. We explore the nature of this phenomenon below after a summary of our robustness checks.

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 $n = 14$ cities from the stable sample 1651–1790. The trend estimates for the SD of five-year-prices at the national level is between -0.0014 and -0.0011 (stable sample: -0.0013) and always significant at $p < 0.01$. Similarly, the trend of the inter-regional SD also remains negative between -0.0017 and -0.0012 (stable sample: -0.0015) and statistically significant at $p < 0.01$. The trend estimate for the within SD of North-Western Germany is always negative but not significantly different from zero in 1 out of 91 cases.

In our second robustness check we carried out the tests of our analysis for an unbalanced sample of 33 cities for the period 1651–1790 (see SA3.10). The results remain similar.

Additionally, the analyses for other cereals (barley, oats, wheat), which we report and discuss in more detail in SA3.15, corroborate our earlier results.

3.5 Market integration along major rivers

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).

The map in the data section (Figure 3.1) shows two major rivers, the Elbe and the Rhine, flowing roughly from South-East to North-West. The importance of these rivers for trade costs has been emphasized by Wolf (2009, 852, 856) against the background of the economic integration of the German Empire 1885–1933. To test the hypothesis that falling trade costs on navigable rivers contributed to price convergence in large areas (cf. Chilosi et al., 2013, 47), 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 33 price series.

The first aggregate region consists of the Elbe region and North-Eastern Germany (all details in SA3.13). The former consists of cities connected by the Elbe and its tributaries. 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'. The second aggregate region contains the Rhine region and South-Western Germany. Cities in the Rhine region are located along the rivers Rhine, Main and Moselle. South-Western Germany includes cities in South-Western Germany that are not part of the Rhine region. In what follows, we first report the results found for the variances within these four sub-regions. Second, we discuss the between-region variances.

The decomposition 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 S23 and S24 in SA3.13). By contrast, the variance within South-West shows no statistically significant downward trend. The variance within North-East, where the distances between several cities are relatively small (Figure 3.1, data section) also decreases. These results are consistent with decreasing trade costs along the main river systems and at shorter distances. 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 magnitude of the trend of the between variance of the two aggregate regions 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 (-0.0013 vs. -0.0003, compare Figures S3.13 and S3.24) and insignificant ($p = 0.11$; dummy variables for the crisis of the 1690s and the Seven Years' War included). Moreover, an additional variance decomposition demonstrates that the remaining convergence between the large aggregates (Elbe plus North-East vs. Rhine plus South-West) is by and large due to price convergence within North-Western Germany, which reflects the baseline result (SA3.14). 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.

3.6 The reduction in grain price volatility and long-run sample results

In this section we first show that market integration went together with a moderation in grain price volatility. Second, we compare our baseline result with a sample for the years 1601–1790. This allows to evaluate in how far price convergence between 1651 and 1790 reflects 'only' the reconstruction of the state of market integration that prevailed before the Thirty Years' War (1618–1648). Furthermore, this extension helps to understand the role of exogenous shocks in the volatility moderation.

3.6.1 More stable foodgrain prices

Figure 3.5 shows a substantial reduction of aggregate grain price volatility (calculated as the CV over time) until the French Revolution. The coefficient of the trend counts years and indicates that in ten years volatility is reduced by roughly one percentage point.

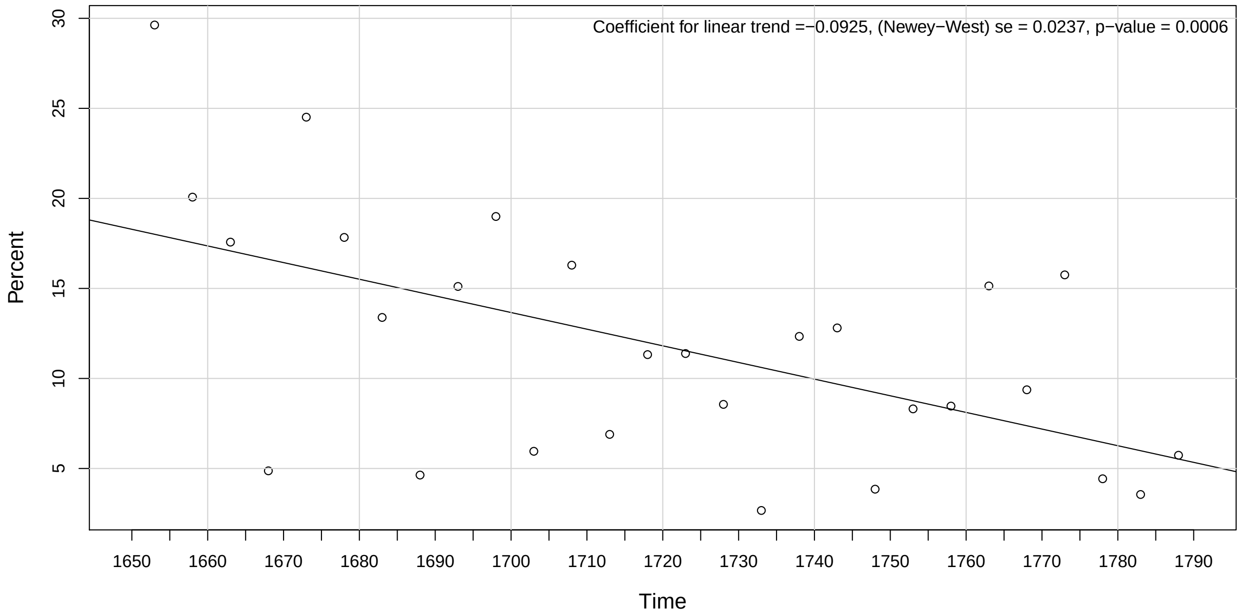

FIGURE 3.5: Volatility of the aggregate real rye price in Germany (stable sample), 1651–1790. Each circle represents a five-year-period centered at the given year (e.g., the circle for year 1653 represents the period 1651–55). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

The pattern visible in Figure 3.5 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; SA3.9.2, Table S30). The result is also robust to sample variation. For aggregate volatility, the trend estimates range between -0.1116 and -0.0772 (stable sample: -0.0925). The estimates are significant at least at the 1% level. All other cereals also show some reduction of volatility until 1790 (see SA3.15).

Note that the moderation of volatility is restricted to the 140 years until the Napoleonic Wars (1792–1815). The latter led to war finance driven inflation that might not be correctly accounted for in currency conversions and real war related shocks; both factors increased volatility (see also Jacks et al., 2011, 807). The decline of volatility prior to the Napoleonic Wars is not confined to Germany. Chilosi et al. (2013, 51, figure 4) show a reduction of wheat price volatility in 18th century Europe (see Persson, 1999, 106–13 on selected European cities; cf. Bateman, 2011, 455–6, 459 who finds only weak evidence of decreasing volatility in Europe and none for Germany).

Integrated markets are expected to have a lower price volatility (Chilosi et al., 2013, 48). But a reduction of the frequency and severity of weather shocks due to the end of the LIA could also have reduced price volatility via a reduction of the fluctuation of cereal yields. In SA3.17 we show that an increase of 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 played a role, at least a part of the volatility moderation was a result of spatial arbitrage, that is, market integration.

While this framework attributes the reduction in volatility either to market integration or shocks, the set of alternative hypotheses is potentially larger and might include changes in storage possibilities or behavior (Federico, 2012, 484). Private carry-over of grain to the next harvest was limited in pre-industrial Europe, because it was risky and hardly profitable but public market interventions using granaries with the aim of stable bread prices existed (Persson, 1996; see Bauernfeind et al., 2001 on Nuremberg). Nevertheless, there is little evidence of a move towards more market intervention; rather the opposite appeared and in the German context (Prussia) mostly after the Revolutionary Wars in the 19th century (Persson, 1999, 131).

The reduction of grain price volatility is relevant, because it increases the stability of the access to food at the aggregate level. Thus, average food security[10] improved. But a major subsistence crisis such as the one around 1770 was still possible (see Section SA3.6.4). That is, the price volatility moderation should not lead to the interpretation that the emergence of grain markets already provided perfect insurance of households against the consequences of price shocks.

3.6.2 The role of post-war reconstruction in grain market integration was limited

For a stable sample of nine cities and roughly six additional cities in an unstable sample, we are able to extend the analysis backwards to 1601 (all details in SA3.12). Over the whole period 1601–1790, price convergence and the decline of volatility were smaller than in 1651–1790 but remain quantitatively significant. The most important difference with respect to the baseline result is that the volatility moderation is attenuated much and that the estimated trend is just significant at conventional levels in the balanced sample ($p = 0.097$) and insignificant in the unbalanced sample ($p = 0.124$). Aggregate volatility is still larger for the 15 years before the Thirty Years' War 1601–1615 (14.3%) than for the years prior to the Napoleonic Wars 1776–1790 (7.6%); but the decrease of some six percentage points or ca. 47% is smaller than in our baseline result where the reduction with 15-year-reference periods is 18 percentage points or 71% (1651–65: volatility was 25.2% in the stable sample; 1776–90: 7.2%).

Furthermore, we find that the disintegration of markets at the national level in the wake of the Thirty Years' War stems entirely from the two sub-regions (North-Western and continental Germany) falling apart.

[10]See Wheeler and von Braun (2013, 509) for a definition and discussion of food security.

Rye price volatility in our data set surged from 14.3% in 1601–1615 to 26.3% in 1651–1665. In North-Western Germany, we also observe a drastic increase of volatility from 14.9 to 25.9 % (based on an aggregate price comprising only North-Western cities of the stable sample). This deserves attention given the fact that spatial dispersion *within* this region remained virtually constant.

Our analytical framework shows that aggregate volatility must increase with decreasing spatial arbitrage; but empirically arbitrage remained constant within North-Western Germany. Alternatively, volatility increases with more frequent, larger or more symmetric shocks. Against this background, the surge in price volatility in North-Western Germany during the period 1651–1675 compared to the pre-War level can be explained by the changing behavior of exogenous shocks. A potential explanation 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 SA3.8.1). Thus, a substantial amount of the decline of price volatility within North-Western Germany 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.

3.7 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. We demonstrate that the cross-sectional standard deviation of five-year averages of real grain prices is robust to 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. Five 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 end of 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 within 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 the decline of price gaps 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, the Elbe and Rhine rivers 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.

Fourth, we observe a drastic moderation of price volatility for rye and wheat. Whereas a part of the moderation was likely 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, the reduction of grain price volatility is at least partly a result of market integration.

Finally, the backwards-extension of our sample period reveals that market integration was not purely due to reconstruction after the Thirty-Years' War. From the early 18th century price dispersion was lower than at the beginning of the 17th century. The reduction in volatility is smaller but quantitatively still significant.

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 could produce 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 dismissed. Empirically, previous research has found little evidence so far for an advance of market integration in Central Europe between 1650 and 1800.

Our results show that market integration improved significantly before the railway age with two important effects. One major consequence was Smithian growth. The timing of price convergence coincides with the rise of regional export industries or proto-industries (Kaufhold, 1986) and a non-stationary German real wage since the late 17th century (Pfister, 2017, 723). Additionally, more stable foodgrain prices improved aggregate food security. Potentially, this exerted a positive impact on two further variables that are important for long-run development: human capital and mortality. The decline in grain price shocks was associated with fewer and less severe food crises, which likely attenuated negative shocks to human capital (Baten et al., 2014). Furthermore, improved market integration is one factor that possibly contributed to the mortality decline in Germany. Improvements in survival chances and grain market integration can be linked theoretically (Ravallion, 1987a, ch. 2) and appeared in several pre-industrial European economies (Persson, 1999, 33–4, 39). More research in this direction could improve our understanding of the mechanisms underlying the transition to the post-Malthusian growth regime.

Chapter SA3

Supplementary Appendix: Market Integration or Climate Change? Germany, 1650–1790

Authors: Hakon ALBERS*, Ulrich PFISTER† and Martin UEBELE‡

SA3.1 Data appendix

Section SA3.1.1 provides an overview. SA3.1.2 explains the principles followed in converting original prices into grams of silver per litre. Section SA3.1.3 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 SA3.1.4).

SA3.1.1 Overview

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).

To improve data quality, we convert data from sources containing crop year prices to a common time base (calendar year) by building on a procedure proposed by Bateman (2011, 451), 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.[1] The transformation of *Martini* prices increases the correlation of prices in our stable sample. Hence, the standardization of the time reference of grain prices is highly relevant for market integration studies (Table S3.12

****Corresponding Author**. Martin Luther University Halle-Wittenberg, Department of Economics, Universitätsring 3, D-06108 Halle, Germany

†University of Münster, Institute of Economic and Social History, Germany.

‡Düsseldorf, Germany. Former address: University of Groningen, Netherlands.

[1]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.

in SA3.3.1). We also develop regional series of silver contents for currency conversion, which is crucial for analyses of price dispersion accross space.

SA3.1.2 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.

SA3.1.3 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 S3.1). However, two forms of data require a transformation of the original series: crop year and *Martini* prices.

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 SA3.3. 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 SA3.3).

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

TABLE S3.1: 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			Altenburg	x		
Cologne	x			Breslau	x		
Dresden	x			Celle	x		
Hamburg	x			Emden	x		
Xanten	x			Landshut	x		
				Nordhausen	x		
				Nuremberg	x		
				Quedlinburg	x		
				Trier	x		
Augsburg	x	x		Leipzig		x	
Munich	x	x		Frankfurt a. M.		x	
Würzburg	x	x		Speyer		x	
				Überlingen		x	
Göttingen	x		x	Halberstadt			x
Halle	x		x	Hanover	x		x
Münster			x	Lüneburg	x		x
Osnabrück			x	Magdeburg	x		x
Paderborn			x	Minden	x		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.

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.

SA3.1.4 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 SA3.3 in Tables S3.6–S3.11 (*Martini* prices) and S3.16–S3.17 (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. (S3.4); the one for crop year prices in eq. (S3.8). The commodity-specific parameters for crop year conversions are in Table S3.18.

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 S3.2 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 S3.2: Interpolated years in data set

Augsburg			Frankfurt				Leipzig			
barely	oats	rye	barley	oats	rye	wheat	barley	oats	rye	wheat
1653	1759	1650	1722	1729	1747	1651	1660	1767	1655	1664
		1759	1735	1745	1755	1660	1670		1660	1666
			1740	1767	1780	1706	1689		1681	1675
			1743	1781	1785		1767		1687	1697
			1772	1786					1767	1701
			1781							1758
Munich			Überlingen			Wuerzburg			Speyer	
barley	rye	wheat	barley	oats	rye	oats	rye	wheat	rye	
1684	1689	1662	1731	1666	1669	1704	1685	1654	1677	
1687		1666		1684	1684	1712	1687	1690	1686	
		1682		1697	1693	1724	1779	1695	1698	
				1702	1718	1740	1791	1697	1799	
				1711	1731	1750		1792		

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.

SA3.2 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 S3.2–S3.5 in Section SA3.4 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 SA3.3.2).

Altenburg

Currency and volume conversion
Conversion to grams of silver as for Leipzig based on Pfister (2017); litre equivalent of local *Scheffel* from Anonymous (1864, 412).

Rye (1746–1863)

Calendar year prices in *Reichstaler*, *Neue Groschen* and *Pfennig* given by Anonymous (1864).

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 S3.18).

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 S3.16. 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 S3.18). 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*. Factors for Berlin in a related publication (Jacobs and Richter, 1935) are equal to those in Statistisches Reichsamt.

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.

Breslau

Currency and volume conversion
Currency conversion as for Berlin except for the period of the Prussian vellon inflation (1808–1823). During this period we do not apply adjustment factors for Berlin but those for Breslau from Jacobs and Richter (1935, 20). We use the currency information from Berlin also for the period 1741–1762, because there is no important level difference between the converted rye prices in Breslau and Berlin before and after the 1760s. In 1750, the same division of coins applied as in the rest of Prussia (Schrötter, 1908, 552).

We convert old and new *Scheffel* as for Berlin.

Rye and wheat (1741–1871)
Average of June and December prices 1740–1810 in *Reichstaler* per *Berliner Scheffel* from Königliches Statistisches Bureau (1867, 113–4); annual prices in *Guten Groschen* (until 1821) or *Silbergroschen* (from 1822) per *Berliner Scheffel* for the years 1811–1859 (except 1818, 1824) from Geheimes Staatsarchiv Berlin (b); annual prices 1818, 1824 and 1860–64 in *Silbergroschen* per *Berliner Scheffel* throughout (that is, including 1818) from Königliches Statistisches Bureau (1867); same type of data 1865–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72).

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. (S3.4)), 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 S3.6.

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. (S3.8) using the commodity-specific parameters in Table S3.18.

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 S3.6 and S3.7. 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).

Halberstadt

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.

Volume conversion applies 1 *Halberstädter Wispel* = 22.5 *Berliner Wispel*; 1 *Berliner Wispel* = 24 *Berliner Scheffel* (Naudé and Schmoller, 1901, 563–4); the latter converted as for Berlin.

Barley, rye, wheat (1639–1740; 1816–1871), oats (1816–1871)
Martini prices for the years 1638–1740 from Naudé and Schmoller (1901, 561–3). Data are in *Taler* per *Halberstädter Wispel*; from 1720 per *Berliner Wispel*. Calendar year prices extrapolated from *Martini* prices applying the general extrapolation rule (eq. (S3.4)).

Data for 1741–1815 not available.

Calendar year prices for the years 1816–1859 in *Guten Groschen* (*Silbergroschen* from 1822) per *Berliner Scheffel* from Geheimes Staatsarchiv Berlin (b); 1824 and 1860–4 from Geheimes Staatsarchiv Berlin (a); 1865–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72).

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 S3.7). 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 S3.8.

Landshut

Currency and volume conversion

Currency conversion follows Munich. *Scheffel* of Landshut prior 1813 are converted to litre using the relationship to the *Scheffel* of Munich given by Seuffert (1857, 138); for the litre equivalent of the latter see Munich. We also use a different volume for oats. We apply the new Bavarian *Scheffel* from 1813 as for Augsburg.

Barley, oats (1584–1700), rye, wheat (1584–1700; 1815–55)

Calendar year prices in *Gulden* and *Kreuzer* per *Scheffel* of Landshut from Seuffert (1857, 138, 282–5).

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 S3.18; 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 S3.8 and S3.9.

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.

Magdeburg

Currency and volume conversion

Currency conversion follows Berlin including periods of debasement.

Volume conversion assumes 1 *Magdeburger Wispel* = 24 *Magdeburger Scheffel*; 1 *Magdeburger Scheffel* = 6/7 *Berliner Scheffel*, for the latter see Berlin (Naudé and Schmoller, 1901, 516, 549).

Barley (1667–1756; 1816–71), rye, wheat (1642–1756; 1816–71), oats (1740–1756; 1816–71)
Calendar year prices for the years 1642–1739 extrapolated from *Martini* prices applying the general extrapolation rule (eq. (S3.4)). *Martini* prices in *Taler* and *Groschen* per *Magdeburger Wispel* (from 1714 *Berliner Wispel*) from Naudé and Schmoller (1901, 552, 545–8) and Naudé et al. (1910, 615). Prices for 1641–66 based on an unspecified manuscript; 1667–1739 based on a list created by the municipal authorities in 1747.

Calendar year prices 1740–1756 based on Naudé et al. (1910, 601–614). Original source is the *Wöchentliche Magdeburgische Frag- und Anzeigungsnachrichten*. Naudé et al. filled gaps in 1747 and 1748 with monthly reports written by the *Magdeburger Kammerpräsident* (President of regional authorities). For rye the transcribed annual average for 1755 calculated by Naudé et al. is implausibly high and was thus replaced with a newly calculated average based on the given monthly data.

1757–1815: gap.

Calendar year prices for 1816–71 in *Guten Groschen* (until 1821; except 1818) and *Silbergroschen* (1818 and from 1822) per *Scheffel* of Berlin from Prussian sources. Data for the years 1816–1859 are from Geheimes Staatsarchiv Berlin (b); those for 1818, 1824, 1860–64 from Königliches Statistisches Bureau (1867); and data for 1865–1871 from Königlich Preussisches Statistisches Bureau (1865; 1867–72).

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 S3.9.

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 S3.16. 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 S3.16 and S3.17. 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 S3.10). 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).

Nordhausen

Currency and volume conversion

To convert local currency to grams of silver for the period 1669–1763, we apply the average rate of grams silver per *Taler* from three geographically close cities: Göttingen (currency of Hanover), Halle (Berlin/Prussian currency) and Leipzig (Saxony). We use the basic relationship of 24 *Gute Groschen* per *Taler*.

For the years 1764–70, we follow Gerhard (2002, 270) and apply the *Konventionsfuß* (20 *Gulden* = 10 *Konventionsspeciestaler* per *Mark* of Cologne). That is, the currency conversion applies the same fine metal content as Leipzig.

Currency conversion from 1816 follows Berlin (including the period of debasement, 1808–21).

Scheffel of Nordhausen are converted using the local rate given by Witthöft (1993, 349). For the conversion of *Berliner Scheffel* see Berlin.

Barley, oats, rye and wheat (1669–1770; 1816–1871)
Data for 1669–1770 are based on monthly prices in *Guten Groschen* per *Scheffel* of Nordhausen for two qualities (good/bad) from Oberschelp (1986, 51–77). To obtain a single calendar year price for each type of grain, we first calculate calendar year averages and then average across the two qualities. We do not calculate a calendar year average for the years 1668 and 1771, because only a few months are available for these years.

1771–1815: gap.

Data for the years 1816–1871 are from various Prussian sources in *Guten Groschen* (until 1821) or *Silbergroschen* (from 1822) per *Scheffel* of Berlin. Data for the years 1816–1859 are from Geheimes Staatsarchiv Berlin (b); those for 1824, 1860–64 from Königliches Statistisches Bureau (1867); and data for 1865–1871 from Königlich Preussisches Statistisches Bureau (1865; 1867–72).

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. (S3.4)). 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 S3.10 and S3.11). 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. (S3.8) (parameters in Table S3.18), because prices from Elsas and Seuffert overlap for only one year. Calendar year prices 1821–55 from Seuffert (1857, 323); accessed through Pfister (2017).

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).

Überlingen

Currency and volume conversion
Currency follows the series in Augsburg (Pfister, 2017) except for the values in 1757–60, which were replaced by linear interpolation based on the values in 1756 and 1761. This is because nominal values suggest that quotes were made in good currency.

Volume conversion follows Göttmann (1991, 485).

Barley (1723–1811), oats (1663–1811), rye (1661–1811)
Crop year prices (August–July) in *Gulden* per *Überlinger Malter* from Göttmann (1991, 478–9). Calendar year prices were obtained by extrapolation following eq. (S3.8) using the commodity-specific parameters in Table S3.18.

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 S3.17 (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. (S3.8) using the parameters from Table S3.18.

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.

SA3.3 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.

SA3.3.1 Martini prices

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

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 SA3.1.3 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:

$$p_{it} = \alpha_0 + \beta_1 m_{it} + \beta_2 m_{it-1} + \sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S3.1}$$

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)[2], 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 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. (S3.1) 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 (S3.1) (Table S3.3, model 2, eq. (S3.1); 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.

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

TABLE S3.3: 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

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

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.[3]

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 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.[4]

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: $\beta_2 = 1 - \beta_1$:

$$p_{it} = \alpha_0 + \beta_1 m_{it} + (1 - \beta_1) m_{it-1} + \textstyle\sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S3.2}$$

The latter eq. can be written as:

$$p_{it} - m_{it-1} = \alpha_0 + \beta_1 [m_{it} - m_{it-1}] + \textstyle\sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S3.3}$$

The left-hand-side is the dependent variable of the restricted model. In the estimated restricted

[3]E.g., for rye, model 2 (Table S3.3), the BIC is -1177; with 2 lags, model 6, the BIC is -1160.

[4]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.

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 S3.4.

TABLE S3.4: 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

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. 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 S3.4 contains several other specifications as robustness checks that yield overall similar results. Dropping very small and very large calendar year prices (model 9) and applying the first-differenced estimator (Wooldridge, 2009, 458) yields roughly similar results. Furthermore, increasing the sample by synthetic[5] *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. (S3.1), 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. (S3.1)). 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 S3.4 (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 S3.5, because the results were very similar to those for rye.

[5]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.

TABLE S3.5: 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

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. 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. (S3.1): $p_t = \alpha_0 + \beta_1 m_t + \beta_2 m_{t-1} + u_t$. 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 S3.6– S3.11. 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 S3.6: 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

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

TABLE S3.7: Time series relationship calendar year price and Martini price (part B)

	Göttingen barley	oats	Halle 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

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

TABLE S3.8: Time series relationship calendar year price and Martini price (part C)

	Hanover wheat	rye	oats	barley	Lüneburg 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

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

TABLE S3.9: Time series relationship calendar year price and Martini price (part D)

	Lüneburg rye	wheat	Minden 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

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

TABLE S3.10: Time series relationship calendar year price and Martini price (part E)

	Münster barley	oats	rye	wheat	Paderborn 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

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

TABLE S3.11: 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

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

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:

$$p_t = \alpha_r \cdot \bar{m}_{\text{time series Martini}} + 0.5 \cdot m_t + 0.5 \cdot m_{t-1}. \tag{S3.4}$$

Note that $\alpha_r \neq \alpha_0$ from the restricted panel regression (eq. (S3.3); Tables S3.4 and S3.5, 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.[6]

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. (S3.4)) 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 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. (S3.4)) for those series for which we have no or insufficient parallel information on annual prices (Osnabrück and

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

Celle[7]), 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 S3.12 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*: $r = 0.60$ and calendar year price $r = 0.64$; Münster: $r = 0.49$ and $r = 0.61$). In one case price synchronization is overestimated if *Martini* prices are used and the correlation of converted calendar year prices with the rest of the stable sample decreases (Xanten).

TABLE S3.12: 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	0.33
Berlin	0.78	0.85	0.50	0.65
Braunschweig	0.60	0.64	0.48	0.63
Cologne	0.56	0.56	0.63	0.76
Dresden	0.79	0.81	0.40	0.52
Göttingen	0.55	0.60	0.38	0.59
Halle	1.00	1.00	0.43	0.59
Hamburg	0.48	0.64	0.45	0.65
Munich	0.54	0.66	0.20	0.31
Münster	0.59	0.59	1.00	1.00
Osnabrück	0.64	0.69	0.74	0.91
Paderborn	0.55	0.55	0.75	0.83
Würzburg	0.66	0.75	0.37	0.49
Xanten	0.44	0.33	0.79	0.70
Average correlation	0.60	0.64	0.49	0.61

Note: Pearson correlations of price series in column with all price series in rows. 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.

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

SA3.3.2 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.

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. (S3.1)):

$$p_{it} = \alpha_0 + \beta_1 h_{it} + \beta_2 h_{it-1} + \beta_3 h_{it-2} + \sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S3.5}$$

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 (S3.1), 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 SA3.2). 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 S3.13). 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 S3.13: 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

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. 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 S3.13 is generally higher than for estimates based on Martini prices (Table S3.3). 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 S3.13, 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 S3.13). 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 S3.13), 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 ($\beta_2 = 1 - \beta_1 - \beta_3$):

$$p_{it} = \alpha_0 + \beta_1 h_{it} + (1 - \beta_1 - \beta_3) h_{it-1} + \beta_3 h_{it-2} + \sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S3.6}$$

The latter eq. can be written as:

$$p_{it} - h_{it-1} = \alpha_0 + \beta_1 [h_{it} - h_{it-1}] + \beta_3 [h_{it-2} - h_{it-1}] + \sum_{i=1}^{I-1} \alpha_i c_i + u_{it}. \tag{S3.7}$$

The results for model 8 in Table S3.14 directly refer to eq. (S3.7).

TABLE S3.14: 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 (crop y.)	-0.0762^{***}	-0.0756^{***}				
	(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

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. 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 S3.13. 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 S3.14). 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[8]). We report only the two most relevant specifications for each cereal in Table S3.15, because the results were very similar to those for rye, although small differences in the particular weights are evident.

[8]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.

TABLE S3.15: 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

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. 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 S3.16–S3.17 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 S3.16: Time series relationship calendar and crop year price (part A)

	Augsburg rye 1750–1799	Munich oats 1692–1773	Munich rye 1693–1773	Munich barley 1698–1820	Munich wheat 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

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

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

TABLE S3.17: Time series relationship calendar and crop year price (part B)

	Würzburg rye 1656–1777	Würzburg wheat 1658–1776	Munich oats 1691–1799	Munich rye 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

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. 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.

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.[9] Thus, we applied the commodity specific parameters from the restricted pooled model 7 rounded to two digits (Table S3.18). 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.[10] Thus, the equation for extrapolation that is applied if no local time series relationship is available is:

$$p_{it} = \beta_{1r} \cdot h_{it} + \beta_{2r} \cdot h_{it-1} + \beta_{3r} \cdot h_{it-2}. \tag{S3.8}$$

TABLE S3.18: 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 S3.14 and S3.15); $\beta_{2r} = 1 - \beta_{1r} - \beta_{3r}$.

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

[9]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 $p < 0.10$.

[10]Model 7 for wheat (Table S3.15) 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.

similar to the values reported in Table S3.18. 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.

SA3.4 Plots of data coverage and nominal rye price series

Figure S3.1 shows the number of cross-sectional observations over time.

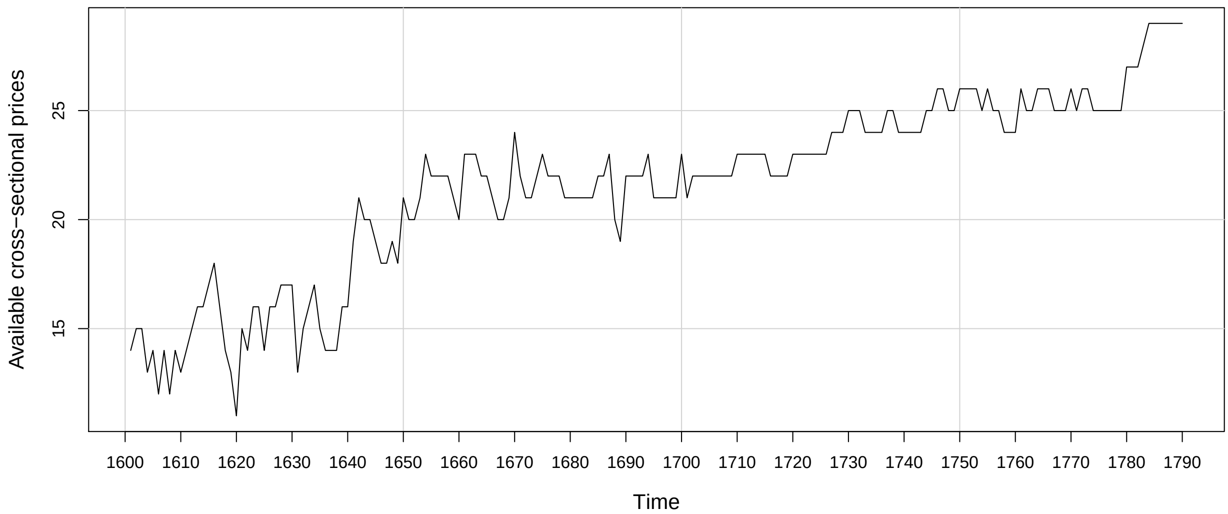

FIGURE S3.1: Data coverage over time, rye prices in individual cities, unbalanced sample 1601–1790

The following Figures S3.2–S3.5 plot the rye price series described in SA3.2 for the period 1601-1850.

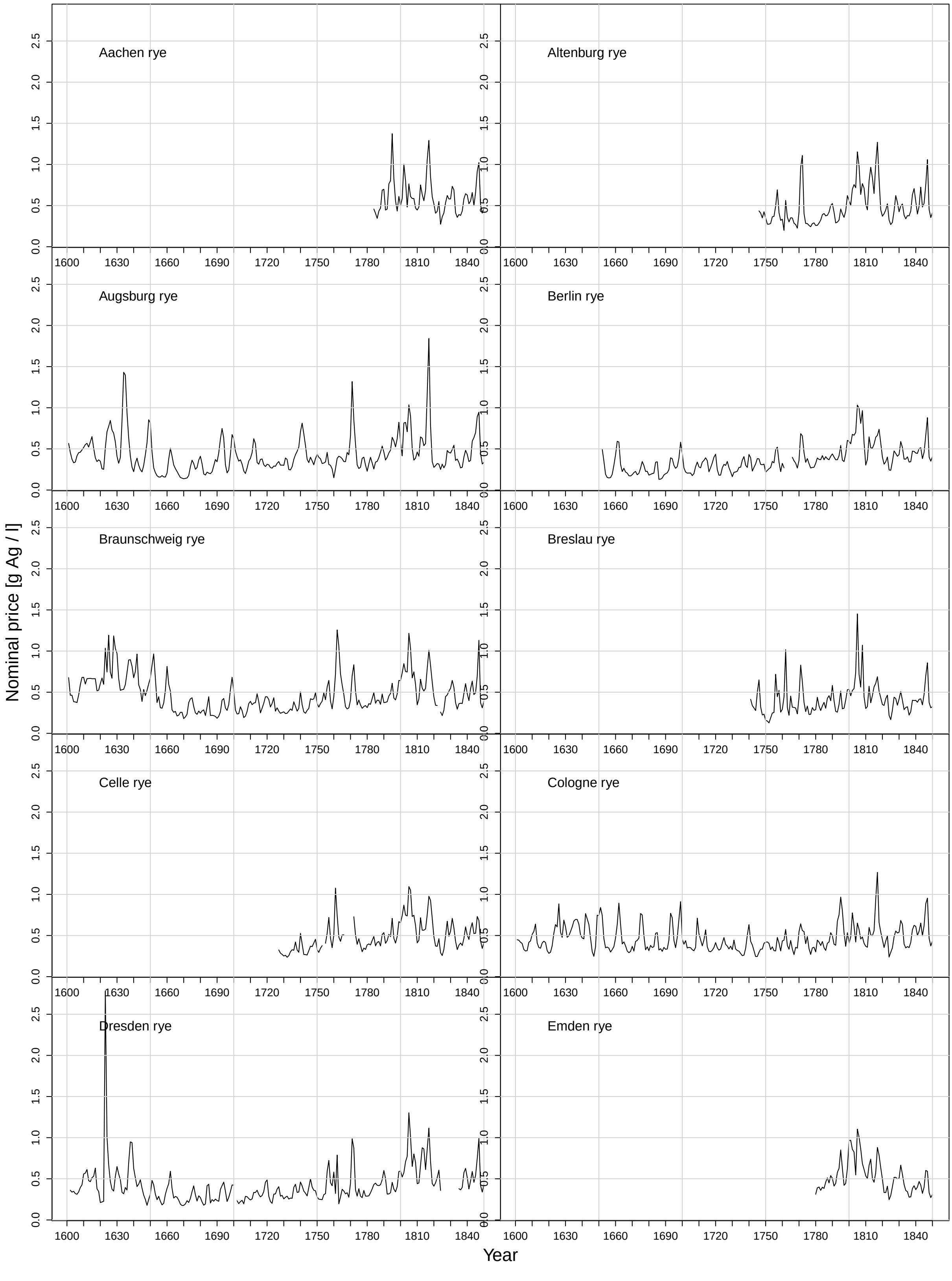

FIGURE S3.2: Nominal rye prices in grams of silver per litre, 1601–1850, part 1.

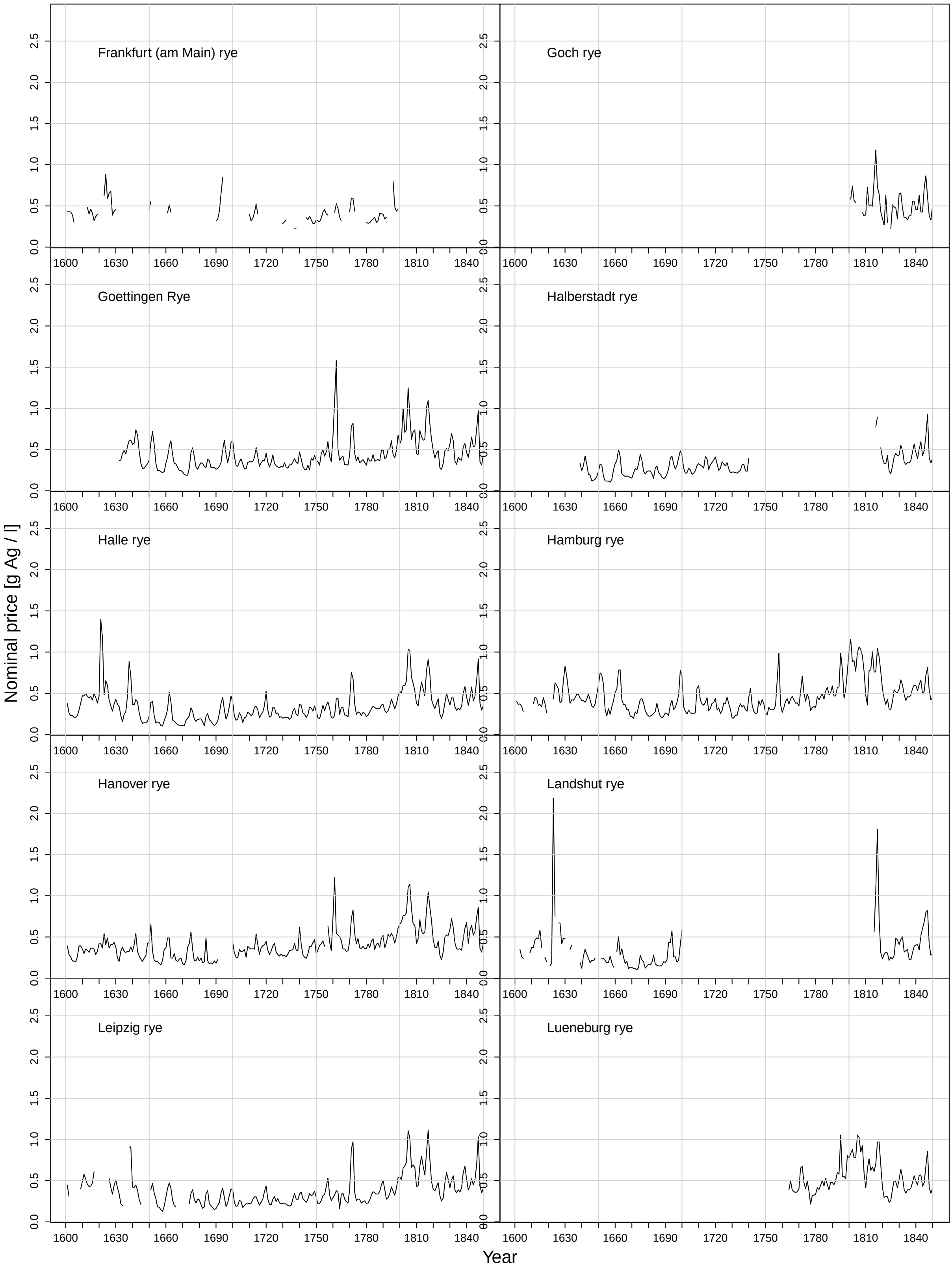

FIGURE S3.3: Nominal rye prices in grams of silver per litre, 1601–1850, part 2.

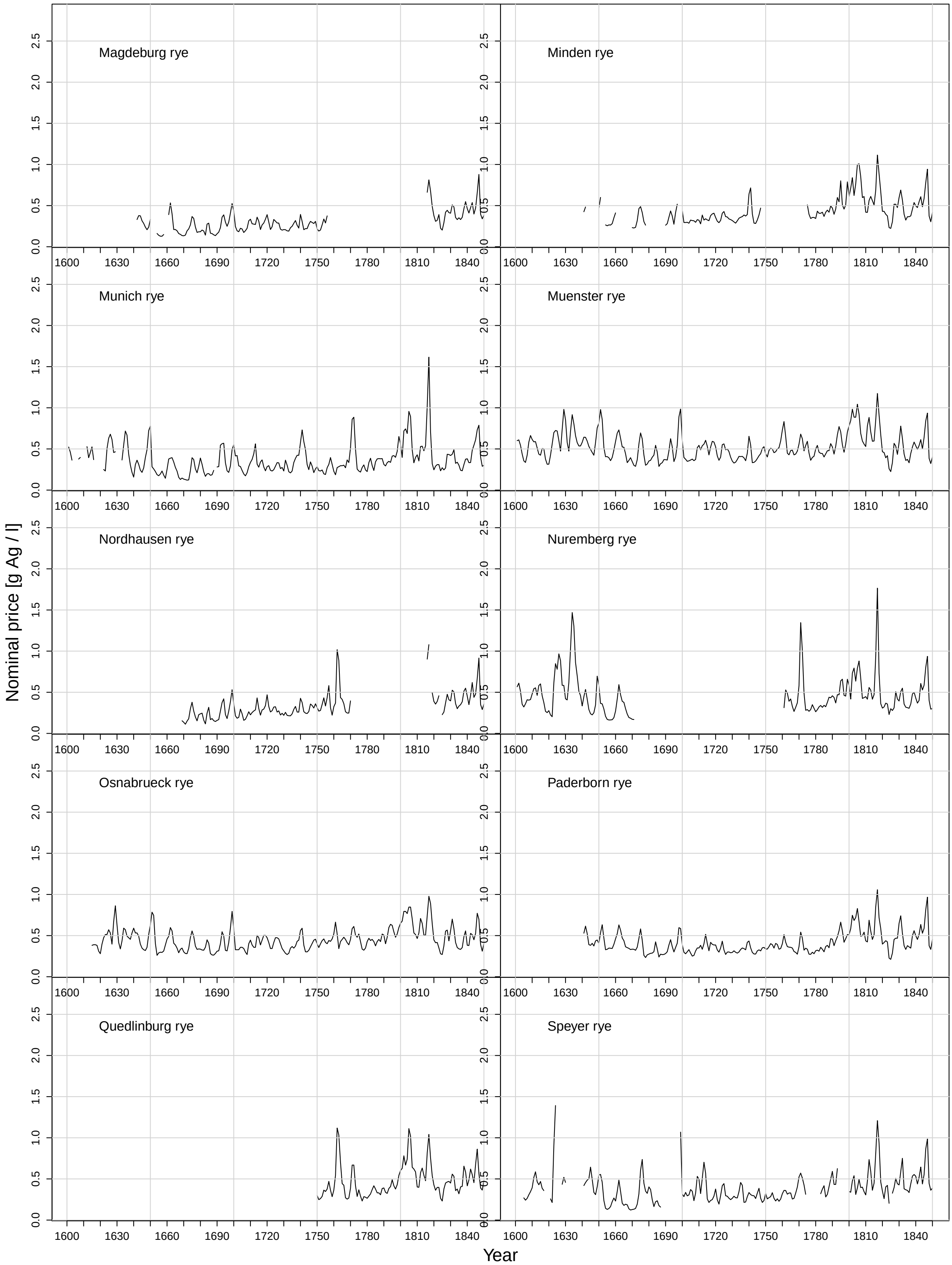

FIGURE S3.4: Nominal rye prices in grams of silver per litre, 1601–1850, part 3.

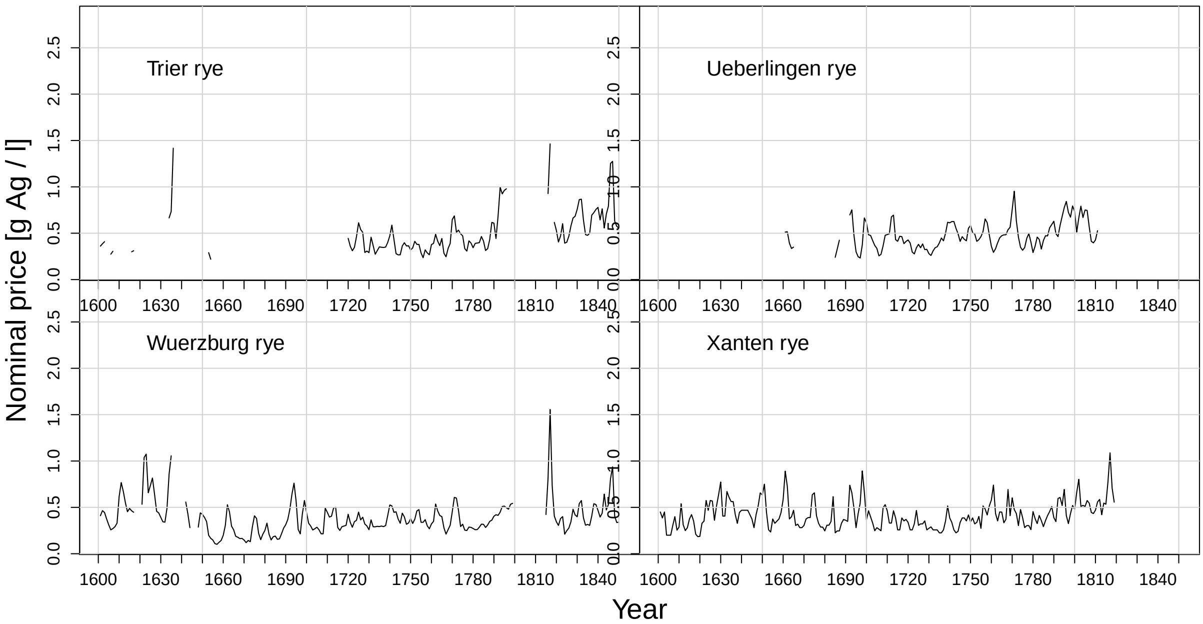

FIGURE S3.5: Nominal rye prices in grams of silver per litre, 1601–1850, part 4.

SA3.5 Time series properties of rye prices

Section SA3.5.1 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 SA3.5.2 contains corresponding results for the aggregate rye price.

SA3.5.1 Rye price series of stable sample

For the 14 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 S3.19). Similarly, the majority of price series is classified as stationary or trend stationary based on KPSS-tests (Table S3.22).

If we split the sample period, a unit root can be rejected at the 10% level in all except one city for the years 1721–1790 (Paderborn; Tables S3.20 and S3.21). The KPSS test results suggest that more series contain a unit root compared to the ADF test (four of fourteen 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 S3.23 and S3.24).

In what follows, we provide more details. Table S3.19 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 S3.19: Augmented Dickey Fuller tests, rye prices, stable sample 1651–1790

	Nominal t-statistic	ρ	log(nominal) t-statistic	ρ	real t-statistic	ρ	log(real) 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
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 fourteen cases not at the 5% level; Table S3.20). 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 S3.21) at the 5% level. Only for one city, namely Paderborn, a unit root cannot be rejected at the 10% level.

TABLE S3.20: 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
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 S3.21: 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
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 S3.22: KPSS tests, real rye prices, stable sample 1651–1790

Null-hypothesis	real rye price, *p*-values level st.	trend st.	log(real rye price), *p*-values 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
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 S3.23: KPSS tests, real rye prices, stable sample 1651–1720

Null-hypothesis	real rye price, *p*-values level st.	trend st.	log(real rye price), *p*-values 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
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 S3.24: KPSS tests, real rye prices, stable sample 1721–1790

Null-hypothesis	real rye price, p-values level st.	trend st.	log(real rye price), p-values 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
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.

SA3.5.2 Aggregate rye price

Table S3.25 shows the results of ADF-tests for the aggregate rye price (for the construction of this variable, see SA3.6) 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 S3.25: Augmented Dickey Fuller tests, aggregate rye price

	Nominal t-statistic	ρ	log(nominal) t-statistic	ρ	real t-statistic	ρ	log(real) t-statistic	ρ
1651–1790								
Stable sample	-5.51	0.58	-5.50	0.61	-5.52	0.57	-5.64	0.57
Unbalanced	-5.68	0.56	-5.64	0.60	-5.69	0.56	-5.82	0.56
Sub-period 1651–1720								
Stable sample	*-3.85*	0.64	*-3.91*	0.65	-4.17	0.57	-5.30	0.58
Unbalanced	*-3.62*	0.66	*-3.87*	0.64	-5.52	0.55	-5.30	0.58
Sub-period 1721–1790								
Stable sample	-4.98	0.50	-4.57	0.56	***-3.02***	0.66	*-3.95*	0.61
Unbalanced	-4.06	0.50	*-3.89*	0.53	*-3.60*	0.58	*-3.60*	0.58

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 S3.26 (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 S3.26: 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	<0.01	> 0.1	<0.01	> 0.1
Sub-period 1651–1720				
Stable sample	> 0.1	> 0.1	> 0.1	> 0.1
Unbalanced	> 0.1	> 0.1	0.0727	> 0.1
Sub-period 1721–1790				
Stable sample	0.0947	0.0129	0.0903	0.0129
Unbalanced	0.0612	0.03	0.0511	0.0304

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.

SA3.6 The aggregate real price of rye in Germany, 1600–1850

In SA3.6.1, we first provide details concerning the construction and deflation of the aggregate real price for rye in pre-industrial Germany. Section SA3.6.2 discusses an alternative approach to calculate the aggregate price (panel regression instead of arithmetic mean). In Section SA3.6.3, 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. Section SA3.6.4 discusses the aggregate price for the period 1651–170 against the background of subsistence crises. Section SA3.6.5 investigates basic relationships between mortality and the aggregate price.

SA3.6.1 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).[11] 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 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.

[11] 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.

SA3.6.2 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:

$$p_{it} = \alpha_0 + \sum_{i=1}^{N-1} \alpha_i c_i + \sum_{t=1}^{T-1} \beta_t y_t + u_{it}. \tag{S3.9}$$

Herein, p_{it} denotes the mean price in city $i = 1, ..., N$ $(N = 34)$[12] 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).

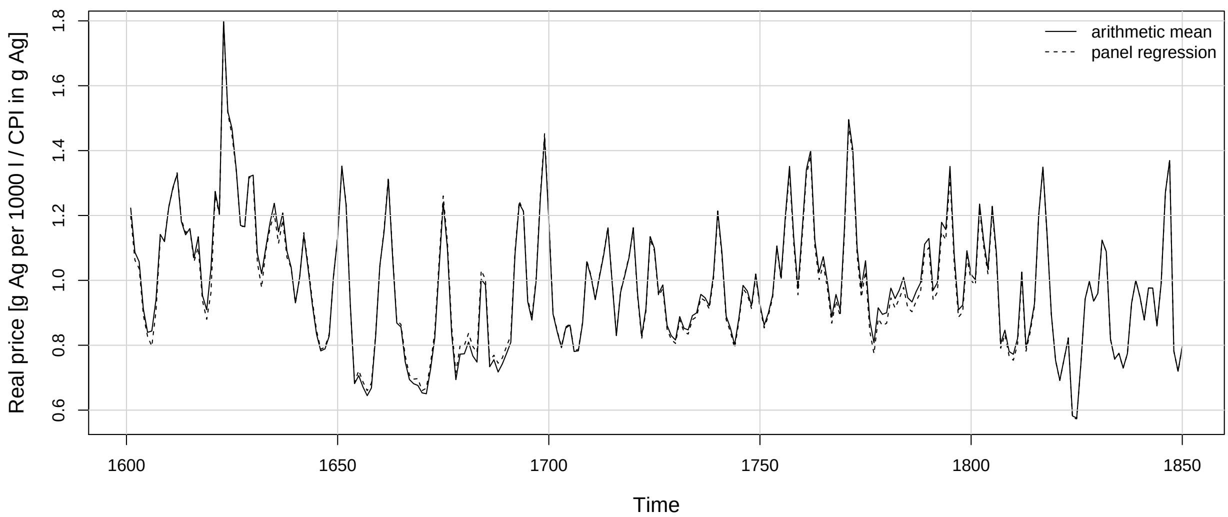

FIGURE S3.6: Two approaches to calculate the aggregate real rye price Germany. Unbalanced sample, 34 cities 1601–1850. Source: own representation.

[12]Inclusion of data for the 19th century increases the unbalanced sample from 33 to 34 cities as the city of Goch enters the sample.

We estimated eq. (S3.9) 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 ($R^2 = 0.67$) is slightly higher than of the variant in logs ($R^2 = 0.65$; see Wooldridge, 2009, 213 on calculation of comparable R^2). Hence, we only consider the results for prices in levels. Figure S3.6 shows the trajectory of the aggregate real price; the latter is very similar irrespective of the methodology used.

SA3.6.3 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.

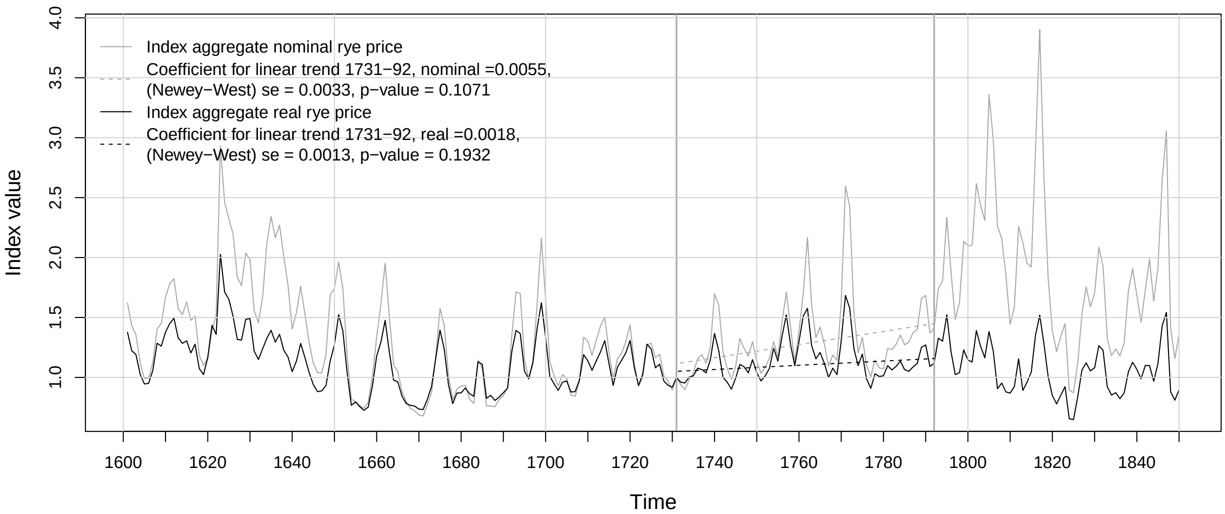

FIGURE S3.7: Index of nominal and real rye price, 1731 = 1; unbalanced sample, 34 cities 1601–1850. Vertical lines highlight sub-period of upward trend in nominal price discussed by Abel (1978, 196–8); 1731 was chosen by Abel; in 1792 the Napoleonic Wars began. Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Source: own representation.

Figure S3.7 shows the nominal rye price, 1601–1850 for the unbalanced sample. There is a just not significant upward trend in 1731–92 ($p = 0.11$), 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 not significant.

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 ($R^2 = 0.73$; stable sample 1651-1790; slope parameter with data in logs but no differences: 0.35; $R^2 = 0.77$). 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)[13] 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 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.

SA3.6.4 Subsistence crises in pre-industrial Germany: evidence from the aggregate real rye price

Panel (a) of Figure S3.8 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. The stable sample for 1651–1790 contains rye prices for 14 cities for 140 years (1953

[13] 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.

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; 3298 observations from 33 cities).

The peaks of the cyclical component (marked with vertical bars and corresponding years in Figure S3.8) indicate serious food or subsistence crises, because they can be linked to increasing mortality and decreasing fertility rates. The majority of peaks occurred during

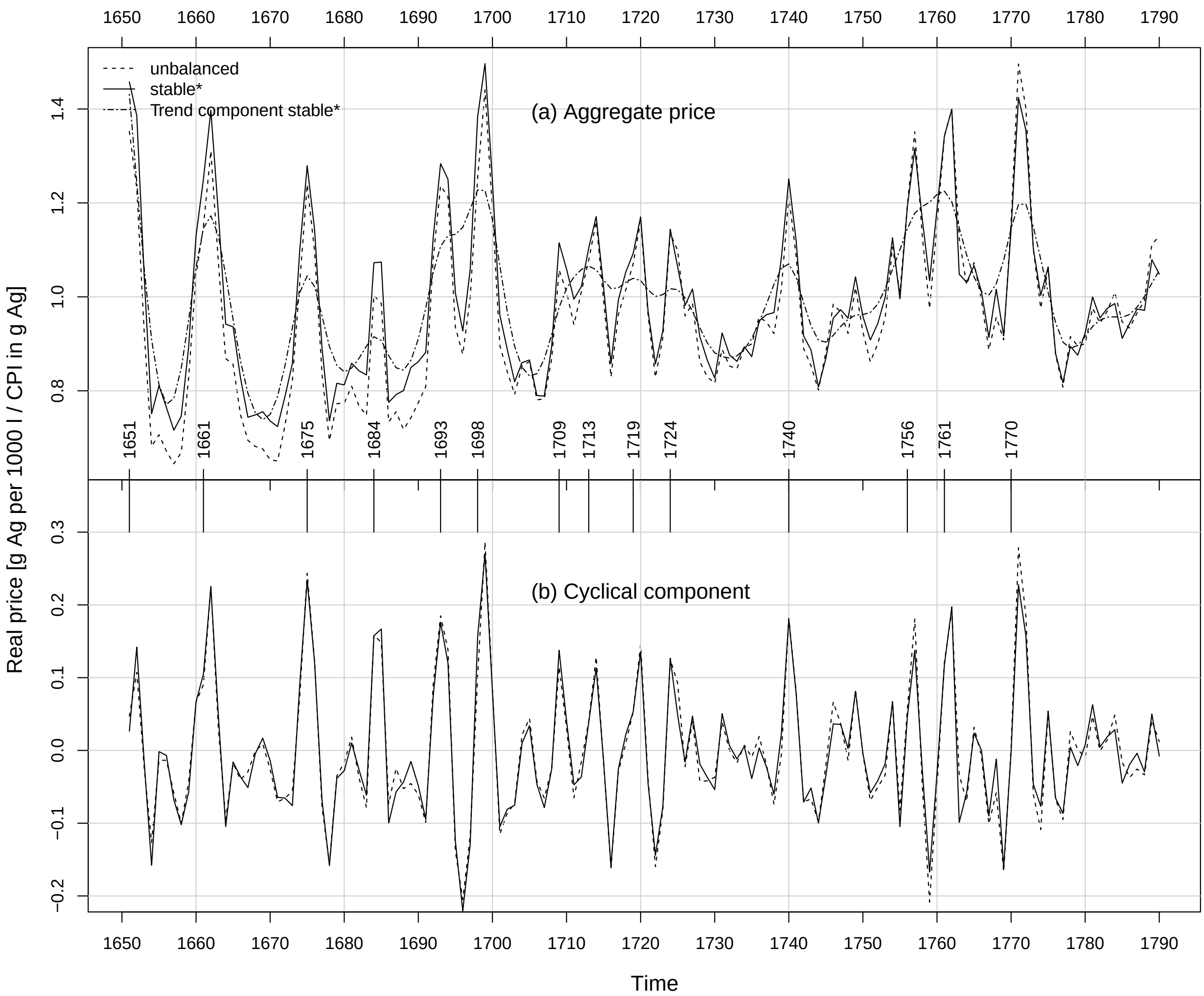

FIGURE S3.8: Aggregate real rye price Germany, 1651–1790. Trend (shown for stable sample series only) and cyclical component from Hodrick-Prescott-Filter, $\lambda = 6.25$ (Ravn and Uhlig, 2002). Vertical lines and given years on upper horizontal axis in panel (b) mark major price peaks associated with 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.

the first half of the sample period, that is, until 1720 (nine out of a total of fourteen peaks). Furthermore, the average deviation of the price peaks from the trend is higher in the 17th than in the 18th 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 are present 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 has been described for North-Western Germany (Lassen, 2016, 90–1, 370) and Saxony (Collet and Krämer, 2017, 106). Based on the aggregate price, the crisis generalizes to the rest of Germany and appears to have been particularly severe.

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 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). Several of the described subsistence crises have been associated with weather shocks (Post, 1985; Collet and Krämer, 2017, 105–8, 117; Lassen, 2016, 90–1, 132–40).

SA3.6.5 Relationship of aggregate rye price and death rate

The *level* of mortality decreases in Germany during the 18th century (Figure S3.9; 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 S3.9, right axis). The main exception to this pattern is the outstanding subsistence crisis around 1770.

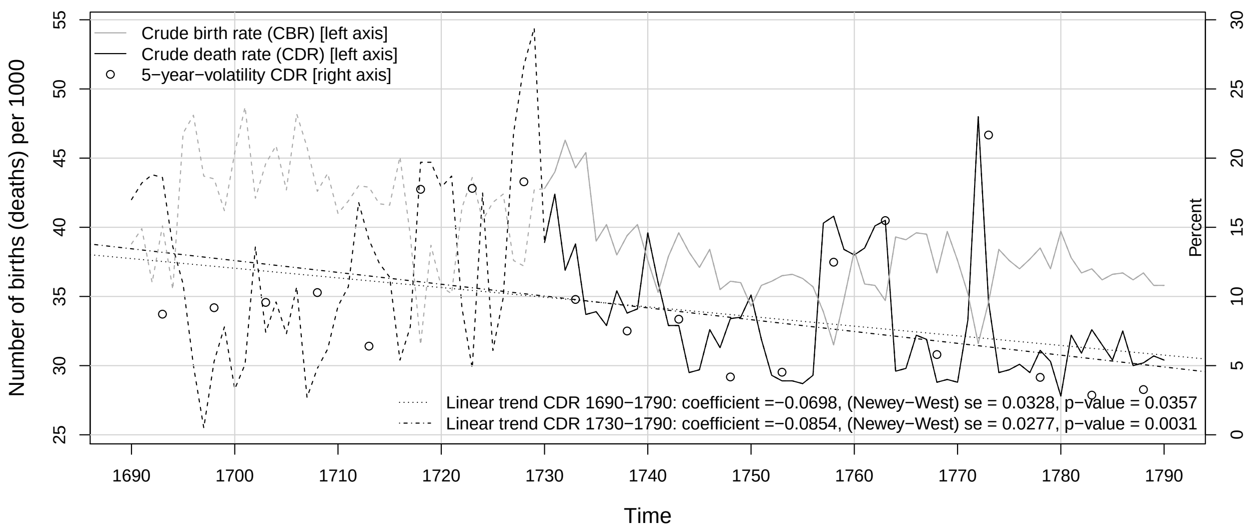

FIGURE S3.9: Crude birth and death rates, Germany 1690/1730–1790. Dashed lines refer to periods with unreliable data which are included for a broader picture. Source: data from Pfister and Fertig (2010); similar as their figure 4, p. 31. Regressions for linear trends include a dummy variable for Seven 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).

We also investigated the short-term relationship of the aggregate rye price and available mortality data during the sub-period 1730–1790 from Pfister and Fertig (2010).

Correlation of aggregate rye price and mortality: 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 0.3% increase of the death rate (1730–1790; $p < 0.05$, $\mathrm{R}^2 = 0.16$; dummy variable for Seven Years' War included). The correlation of log-differenced rye price and death rate is robust to inclusion or exclusion of dummy variables. Dropping the war dummy increases the parameter to 0.34% and reduces $R^2 = 0.12$. 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 0.22% ($p < 0.05$).

Granger causality test: We considered an exploratory bivariate vector autoregressive (VAR) model of the log-differences of the crude death rate and the real rye price. The VAR model 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 (0.51% with additional dummy variable for the crisis year 1772; $p < 0.01$). Vice versa, the second and fourth lag of the CDR are negatively related to the rye price with -0.23% and -0.25%, respectively (-0.24% and -0.27%; $p < 0.05$), which is in line with the intuition that a reduction of population size should decrease demand for grain (the third lag is positive but smaller: 0.19%).

SA3.7 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).[14] 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

[14] 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.

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 $\leq$ 17.0°C (Müller-Westermeier et al., 2001, map 7; Deutscher Wetterdienst, 2018). Breslau (Polish: Wroclaw) is not part of the latter classification. The WMO climate normal 1961–90 yields a continentality of 19.5°C (WMO, 2017a). Thus, Breslau is allocated to continental Germany.

SA3.8 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.

SA3.8.1 Time series properties of weather variables

Table S3.27 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 S3.27).

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 (2014a, 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; $R^2 = 0.13$).

TABLE S3.27: Time series properties weather variables 1651–1790

	ADF-test t-statistic	ρ	*p*-values KPSS-test 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 S3.10 shows the two temperature variables which exhibit a changing pattern during the period 1651–1720 corroborating the results of the KPSS tests from Table S3.27. For example, Xoplaki et al. (2005, 2) have associated the very low spring temperatures with the Maunder Minimum. 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 Northern Hemisphere (Masson-Delmotte et al., 2013, 389, 409; Kelly and Ó Gráda, 2014a, 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).

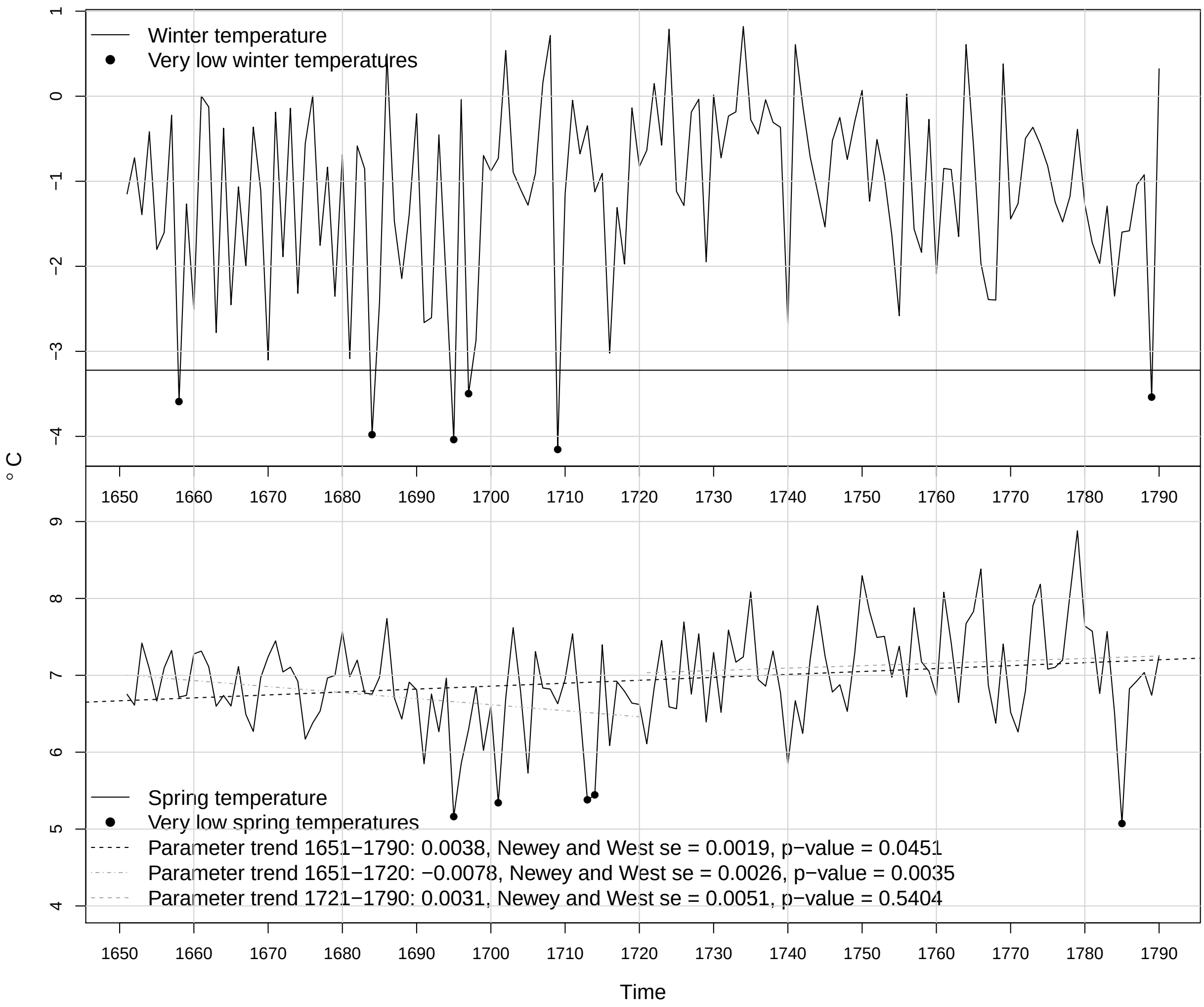

FIGURE S3.10: Temperature reconstructions for Europe 1651–1790. Very low: < mean minus two times the standard deviation. Own representation. Horizontal line in upper panel is drawn at the average winter temperature of 1963 (-3.22°C), when major rivers (Elbe, Rhine, Main) were frozen. Data sources: Luterbacher et al. (2004); Xoplaki et al. (2005); data retrieved from NOAA (2017).

SA3.8.2 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 S3.25 and S3.26). Spring temperature shows a positive trend, winter temperature a changing variance pattern as discussed above and visibile in Figure S3.10. 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} and spring temperature x_{3t}. Additionally, we include two lags of the dependent variable (to account for serial correlation). Hence, the model is: $\log(p_t) = \alpha_0 + \alpha_1 t + \rho_1 \log(p_{t-1}) + \rho_2 \log(p_{t-2}) + \sum_{i=1}^{3} \beta_i x_{it} + u_t$. We also tested an additional interaction term of winter temperature and precipitation, which adds the term $\beta_{12}(x_{1t}x_{2t})$ to the above model. 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 S3.28). 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 1.6% increase of the aggregate price. The interaction term is not significant. Furthermore, the Akaike information criterion (AIC) was lower without interaction term and thus, we favored the latter. The regression specification error test (RESET) was passed at 10% level. Higher spring temperature are associated with lower prices; but the result is not significant at conventional levels ($p = 0.14$; with interaction term $p = 0.11$). However, the Akaike information criterion (AIC) is lower for model 1 compared to model 3 and thus, we favored the latter.

TABLE S3.28: Relation of aggregate real rye price and weather variables 1651–1720

	(1) Baseline	(2) interaction	(3) w/o spring temp.	(4) w/o weather variables
Intercept	-1.4452	-1.5744	-2.2258^{**}	-2.1354^{*}
	(1.2086)	(1.2161)	(1.0977)	(1.1402)
Lag 1 price	1.1129^{***}	1.1172^{***}	1.1144^{***}	1.0935^{***}
	(0.0983)	(0.0984)	(0.0992)	(0.1027)
Lag 2 price	-0.5610^{***}	-0.5776^{***}	-0.5334^{***}	-0.5364^{***}
	(0.0957)	(0.0972)	(0.0948)	(0.0981)
Winter temp.	-0.0199^{*}	-0.0208^{*}	-0.0239^{**}	
	(0.0109)	(0.0110)	(0.0107)	
Winter prec.	0.0016^{*}	0.0017^{*}	0.0015^{*}	
	(0.0009)	(0.0009)	(0.0009)	
Spring temp.	-0.0378	-0.0428		
	(0.0255)	(0.0260)		
Trend	0.0010	0.0011	0.0013^{**}	0.0013^{*}
	(0.0007)	(0.0007)	(0.0007)	(0.0007)
Winter temp. x winter prec.		-0.0008		
		(0.0008)		
R^2	0.7333	0.7375	0.7237	0.6908
Adj. R^2	0.7070	0.7068	0.7014	0.6763
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.

TABLE S3.29: 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	0.0011	0.0015	0.0007	0.0020	0.0015	−0.0031
	(0.0125)	(0.0124)	(0.0132)	(0.0114)	(0.0127)	(0.0097)
Lag 1 price	0.1975	0.1654	0.1414	0.0913	0.2011	0.2849***
	(0.1356)	(0.1370)	(0.1411)	(0.1131)	(0.1373)	(0.0968)
Lag 2 price	−0.1870	−0.2145	−0.1237	−0.4323***	−0.2333	−0.3984***
	(0.1649)	(0.1653)	(0.1720)	(0.1145)	(0.1749)	(0.1031)
Winter prec.	0.2922**	0.2947**		0.2445**	0.2819**	0.2312**
	(0.1130)	(0.1123)		(0.1100)	(0.1149)	(0.0893)
Lag 1 CDR	−0.2253	−0.1973	−0.2134		−0.1921	−0.1630*
	(0.1508)	(0.1514)	(0.1590)		(0.1571)	(0.0821)
Lag 2 CDR	−0.2261*	−0.2148	−0.2158		−0.1932	−0.1373
	(0.1299)	(0.1294)	(0.1369)		(0.1371)	(0.0829)
Lag 3 CDR	0.1322	0.1302	0.1236		0.1284	−0.0060
	(0.1138)	(0.1130)	(0.1199)		(0.1200)	(0.0811)
Lag 4 CDR	−0.2025*	−0.1894*	−0.1829		−0.1929*	−0.2023**
	(0.1124)	(0.1122)	(0.1182)		(0.1146)	(0.0778)
7 Years' War		0.0872		0.0895		
		(0.0678)		(0.0681)		
Spring temp.				−0.1863*	−0.1126	
				(0.1052)	(0.1212)	
Lag 1 autumn temp.				−0.1767	−0.0261	−0.1462
				(0.1191)	(0.1358)	(0.1154)
R^2	0.3056	0.3283	0.2128	0.2752	0.3182	0.3593
Adj. R^2	0.2084	0.2186	0.1202	0.2027	0.1903	0.3004
Num. obs.	58	58	58	67	58	96

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. 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 (for models 2 and 3 rejection at 10% level), 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.

For the 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.[15] The model is:
$\Delta \log(p_t) = \alpha + \rho_1 \Delta \log(p_{t-1}) + \rho_2 \Delta \log(p_{t-2}) + \beta \Delta \log(x_t) + \sum_{i=1}^{4} \gamma_i \log(d_{t-i}) + \Delta u_t$. 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 temperature was significant in some specifications; autumn temperature only marginally significant with p-values close to 0.10 but not below. Both variables were not robust to the inclusion of lagged mortality from Pfister and Fertig (2010). Additionally, we tested a dummy variable for the Seven Years' War (1756–63).

In line with the results for the previous sub-period (1651–1720; Table S3.28), winter precipitation is again significant and of the same sign. This association is also robust to the inclusion of data from 1690 (model 6). Model 1 in Table S3.29 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 S3.28 and S3.29 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, which corresponds to higher prices in our results in both sub-periods (Chmielewski, 1992, 27).[16] Overall, the presented empirical evidence, although it is based merely on reduced form regression models, shows that weather shocks affected the German aggregate rye price, an important macroeconomic indicator of the pre-industrial era.

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[15]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.

[16]In addition, the signs for the variables which are not significant at conventional levels are reasonable. For the years 1721–1790, higher autumn temperatures are related to lower prices; 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, 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.

SA3.9 Supporting information on main results

In what follows, we provide additional details on the main results. Section SA3.9.1 provides a plot of the regional mean prices, Section SA3.9.2 shows further evidence of the price volatility moderation based on a panel regression.

SA3.9.1 Regional real mean prices

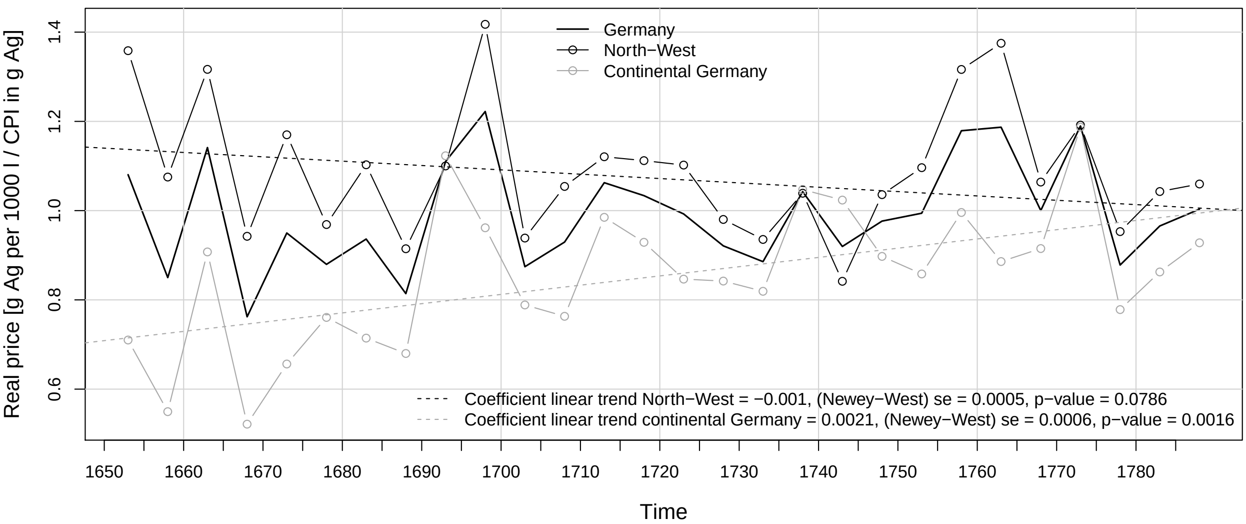

FIGURE S3.11: Real 5-year-mean-prices for Germany, North-Western and continental Germany, rye (stable sample) 1651/5–1786/90. Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA2.

SA3.9.2 Panel estimates for the moderation of grain price volatility

Table S3.30 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 S3.30: Panel analysis of real rye price volatility

	Dependent variable: Volatility
Trend	-0.0656 $(0.0271)^{**}$
R^2	0.0758
Adj. R^2	0.0413
Num. obs.	391

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

SA3.10 Additional results for unbalanced sample 1651–1790

In what follows, we provide the main results for a much larger unbalanced sample of 33 cities for the period 1651–1790. Both results, price convergence before the French Revolution and the reduction of grain price volatility, remain robust.

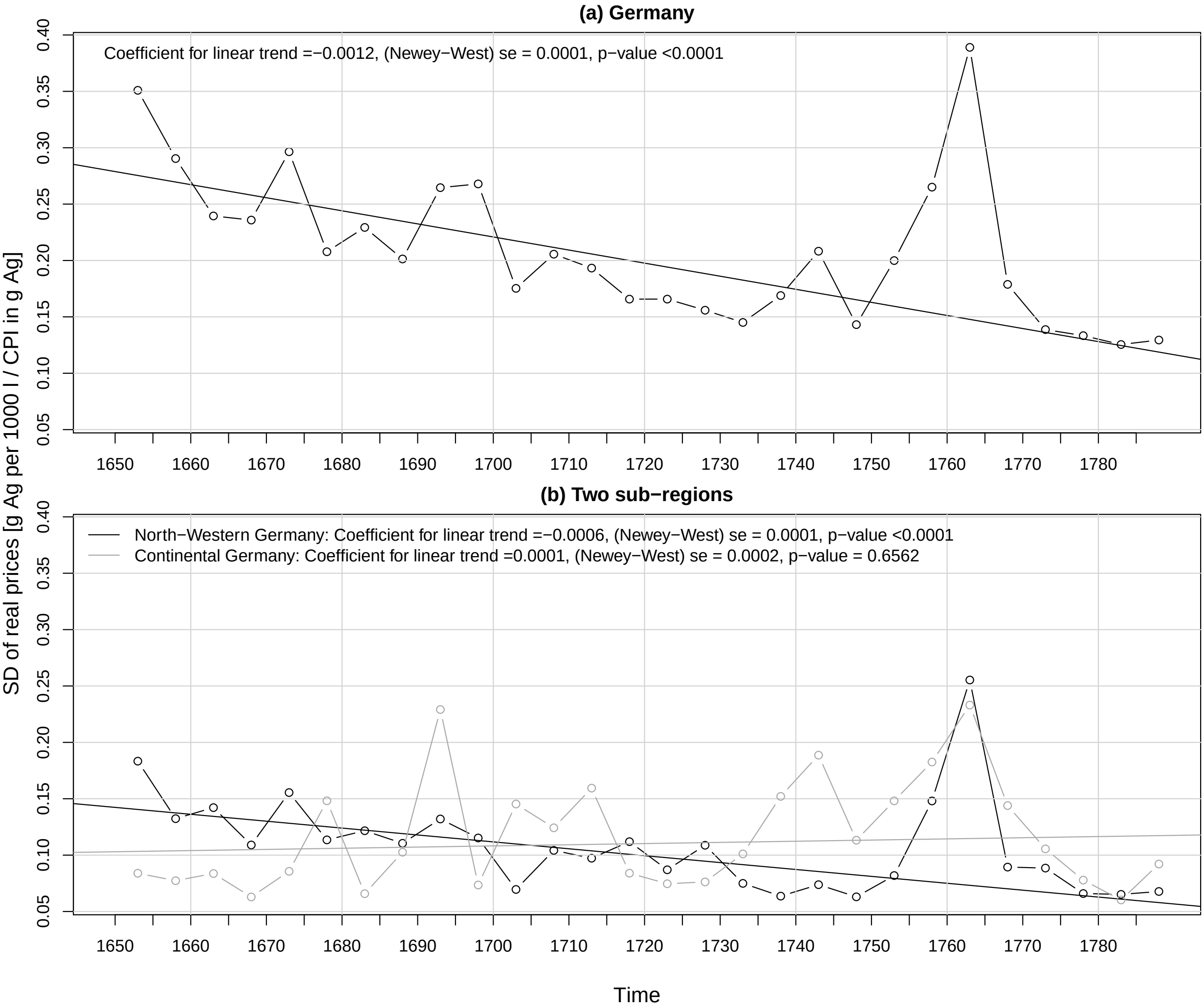

FIGURE S3.12: Inter-urban price dispersion 1651/5–1786/90. Cross-sectional standard deviation of real 5-year-mean-prices, rye (unbalanced sample). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

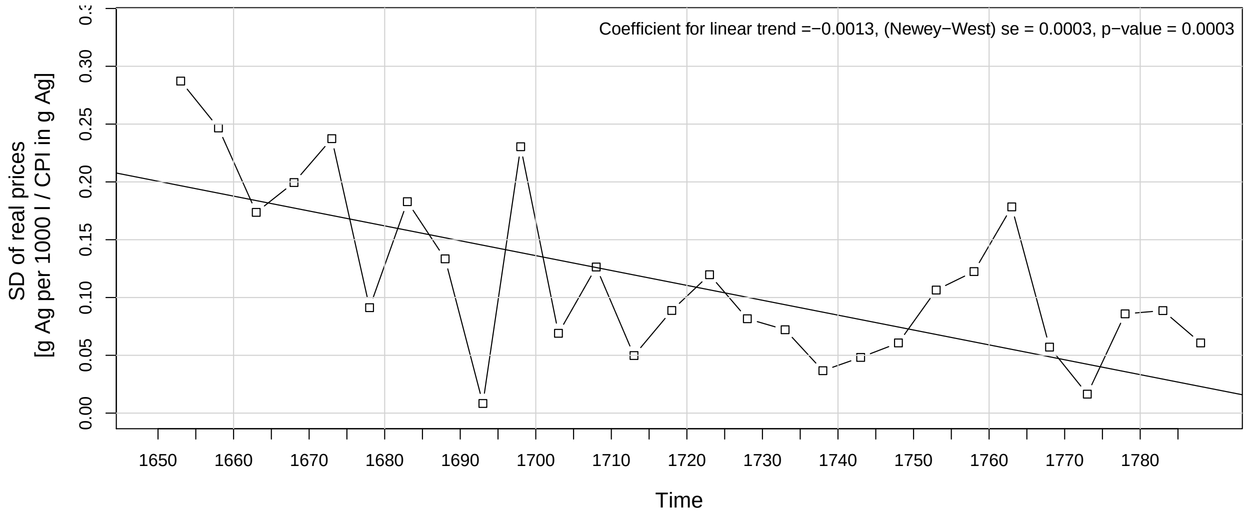

FIGURE S3.13: Inter-regional price dispersion 1651/5–1786/90. Between-region standard deviation (square root of variance between North-Western and continental Germany); real 5-year-mean-prices, rye (unbalanced sample). Regression for linear trend includes a dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

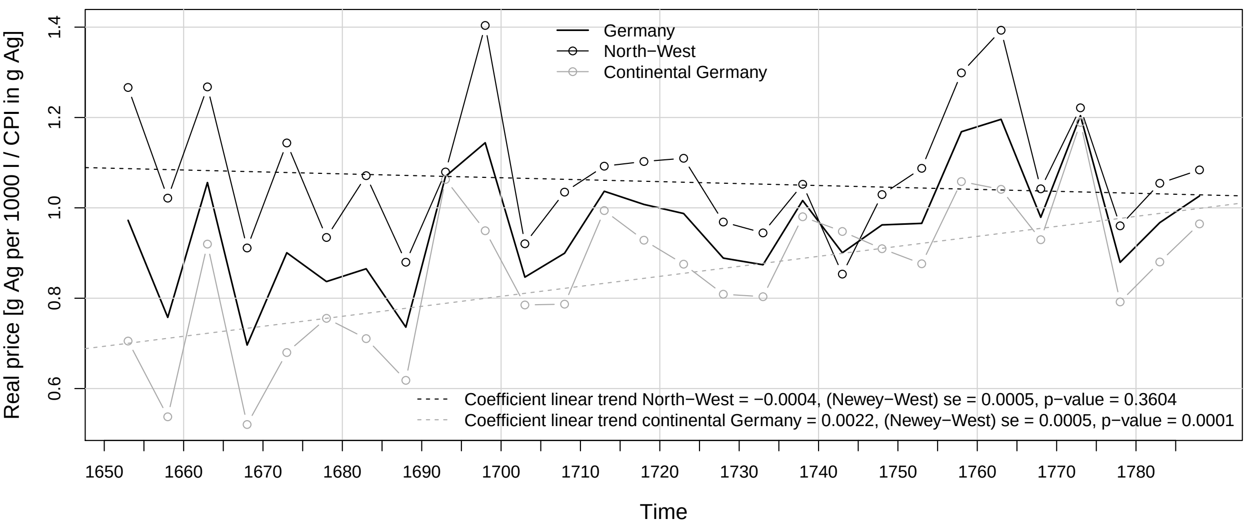

FIGURE S3.14: Real 5-year-mean-prices for Germany, North-Western and continental Germany, rye (unbalanced sample) 1651/5–1786/90. Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

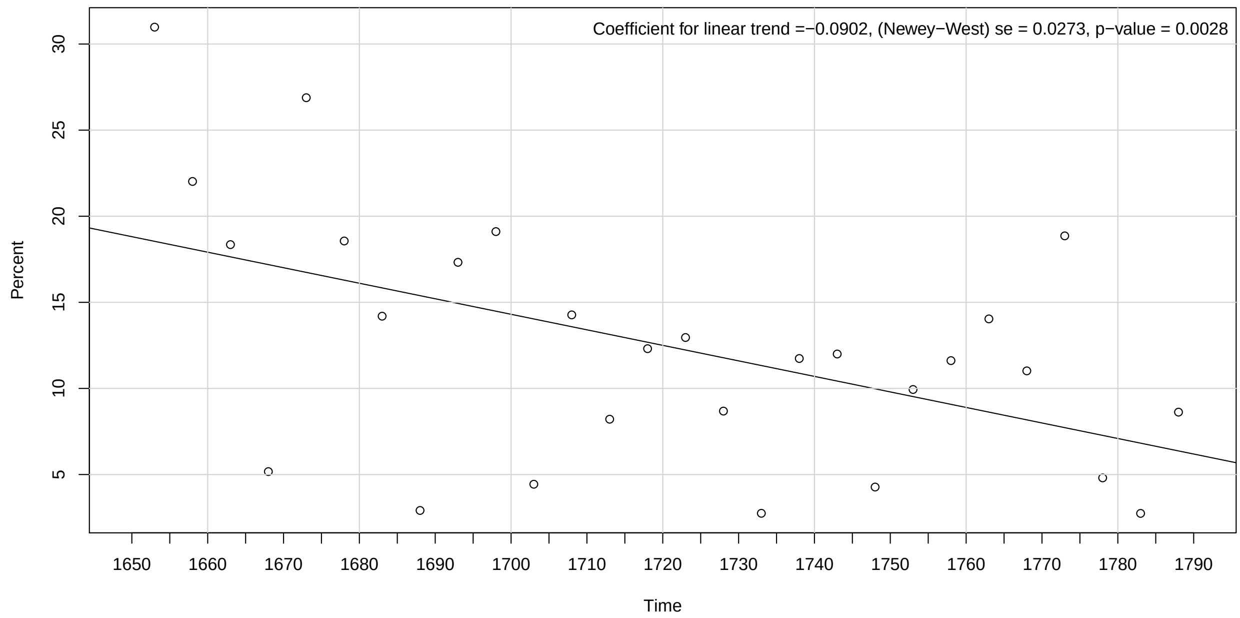

FIGURE S3.15: Volatility of aggregate real rye price Germany (unbalanced sample), 1651–1790. Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

SA3.11 Additional results for price convergence at annual frequency

Figure S3.16 shows price convergence measured as cross-sectional SD in 1651–1790 (stable sample, 14 cities, black solid line).[17] We can exclude analytically that *symmetric absolute* weather shocks or climate change leading to *symmetric absolute* price changes affect the SD.

[17]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.

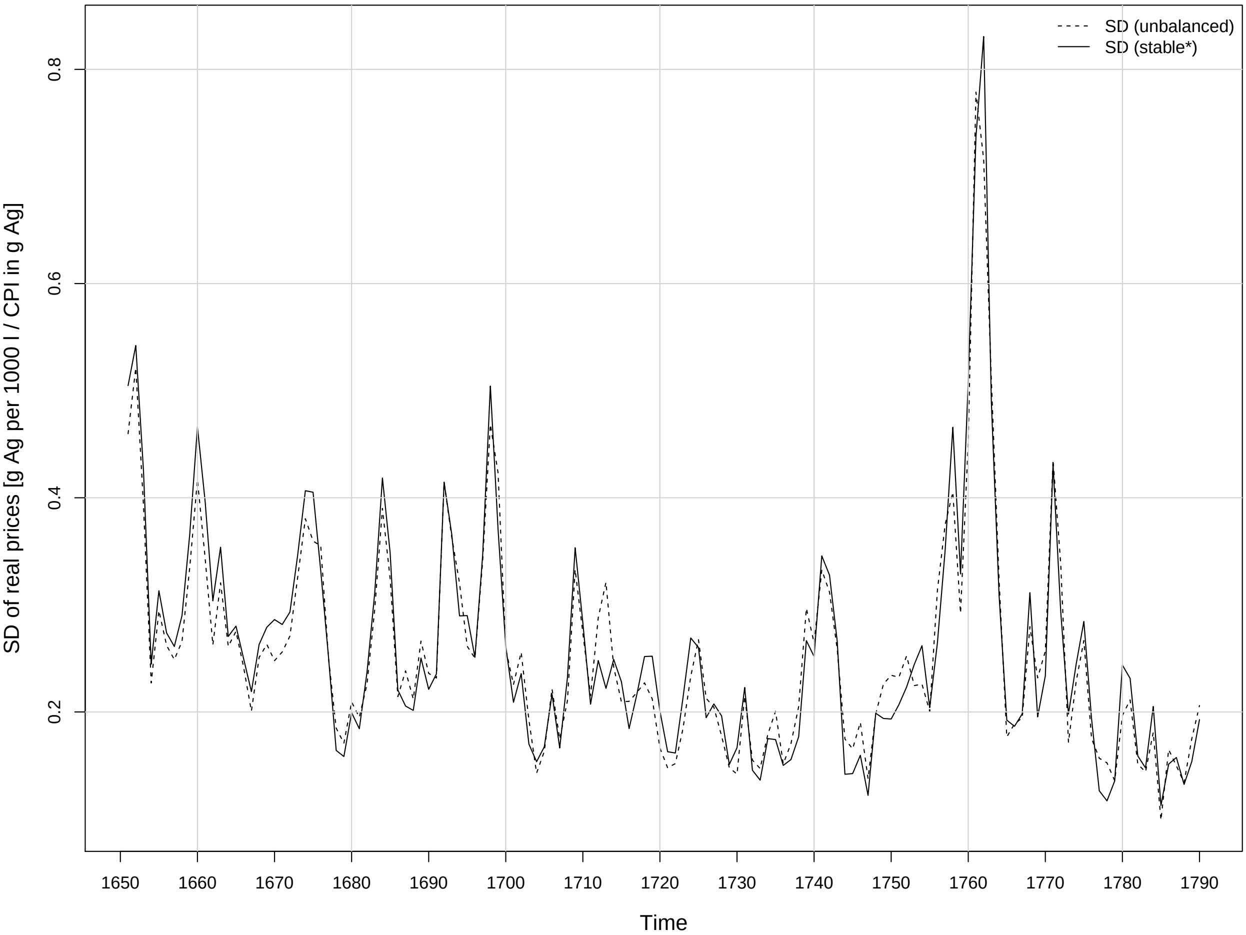

FIGURE S3.16: Inter-urban price dispersion: cross-sectional standard deviation (SD) of real rye prices Germany, annual frequency. *: ≤5% missing observations per individual series. Each cross-sectional nominal price series is deflated with the national CPI from Pfister (2017).

The following plot shows the result for not indexed CV (Figure S3.17).

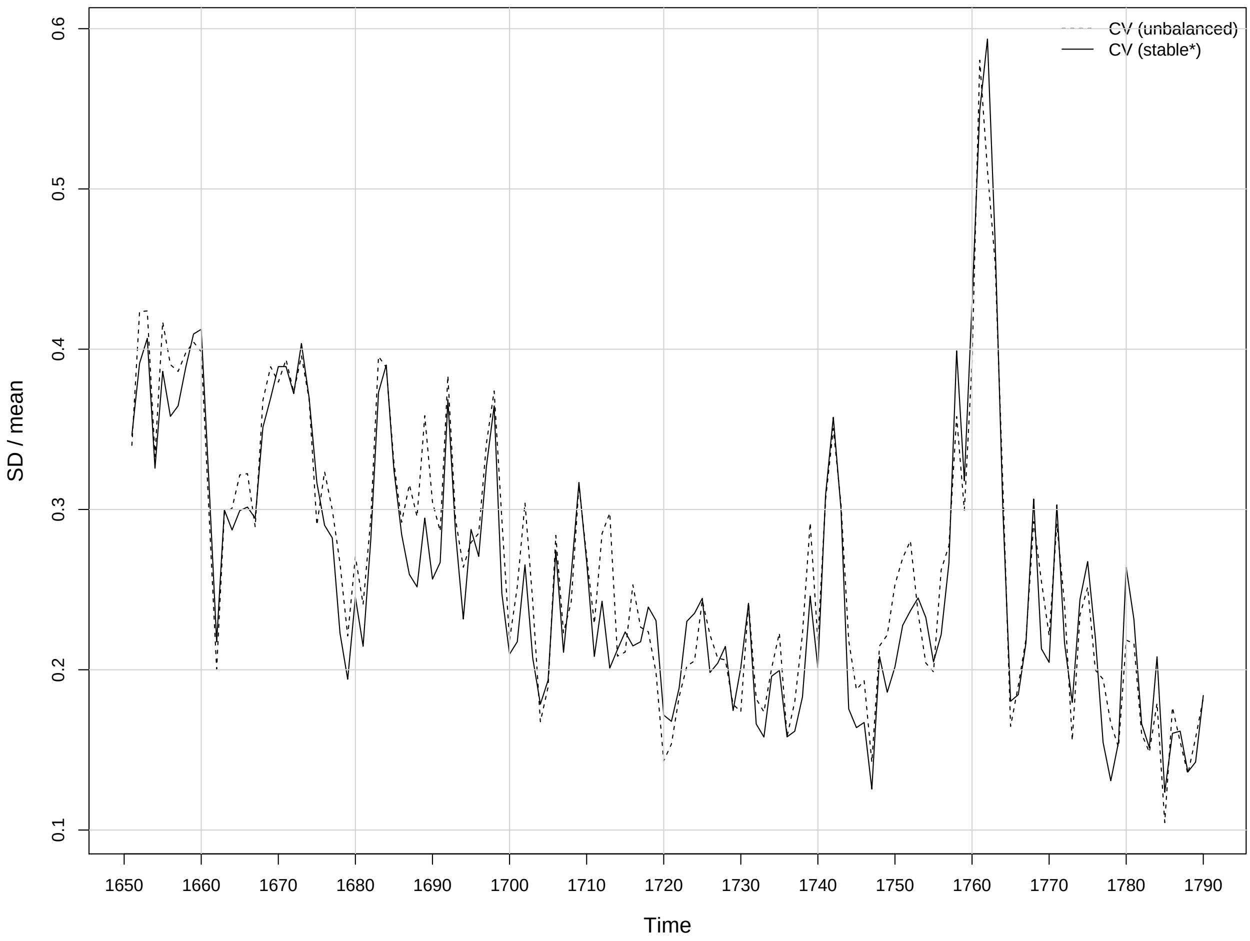

FIGURE S3.17: Inter-urban price dispersion: cross-sectional coefficient of variation (CV) of real rye prices Germany. *: ≤5% missing observations per individual series. Each cross-sectional nominal price series is deflated with the national CPI from Pfister (2017).

SA3.12 Additional results: long-run sample of rye prices 1601–1790

Sample description: Data availability before 1650 is very limited. A stable sample (≤5% missing observations per individual series) of rye prices includes nine, mostly North-Western cities for the period 1601–1790.

- North-West: Braunschweig, Cologne, Hamburg, Hanover, Münster, Xanten
- Continental: Augsburg, Dresden, Halle.

On average there are about six additional cities in the unbalanced sample for the period 1601–1650 (Figure S1 in SA3.4).

Detailed discussion of results for price convergence: Compared to the baseline result, a smaller but still relevant reduction of price convergence and volatility relative to the 15-year-period before the Thirty Years' War occurred. On average, price dispersion decreases by 0.36% per year[18] The corresponding trend estimate of the baseline result with regard to the first 15 years 1651–1665 indicates a reduction by 0.4% per year. We estimated the trend for the long-run sample while accounting for the increased heterogeneity with additional dummy variables (one dummy for the entire war period 1618–48 and an additional one for the Kipper and Wipper inflation in 1620–23). In relative terms, cross-sectional dispersion was 23% from 1601–1615 and it decreased to 14% in 1776–1790. The baseline result with reference periods of 15 years is a reduction from 31% in 1651–1790 to 14% in 1776–1790.

Detailed discussion of disintegration following the Thirty Years' War: Markets disintegrated at the national level following the Thirty Years' War compared to the period before the War: In 1601–1615, the cross-sectional SD of five-year prices was on average 0.25 while it was 0.35 for the years after the War 1651–1665 (unbalanced sample: 0.21 and 0.29). A roughly similar result obtains in the recent work on European wheat prices by Federico et al. (2018, 9, 11–12, figures 2 and 3). Our results for a smaller geographical area show a stronger convergence until 1790 compared to their balanced sample, which covers large parts of Europe but not North-Western Germany. In accordance with the baseline result, the latter region again shows a trend of price convergence. Furthermore, North-Western Germany does *not* exhibit disintegration following the Seven-Years' War (SD 1601–15: 0.22; SD 1650–65: 0.22; SD 1776–1790: 0.09). Continental Germany shows no price convergence but also no disintegration following the War (SD 1601–15: 0.09; SD 1650–65: 0.07; SD 1776–1790: 0.06). However, the dispersion between both sub-regions increases substantially after the War (SD 1601–15: 0.06; SD 1650–65: 0.25; SD 1776–1790: 0.08). That is, the disintegration after the War at the national level stems from these two sub-regions regions falling apart.

Figures S3.18 to S3.22 show the results.

[18] Parameter of trend (-0.0009) divided by the 15-year-average-SD (0.25) for 1601–1615, multiplied by 100.

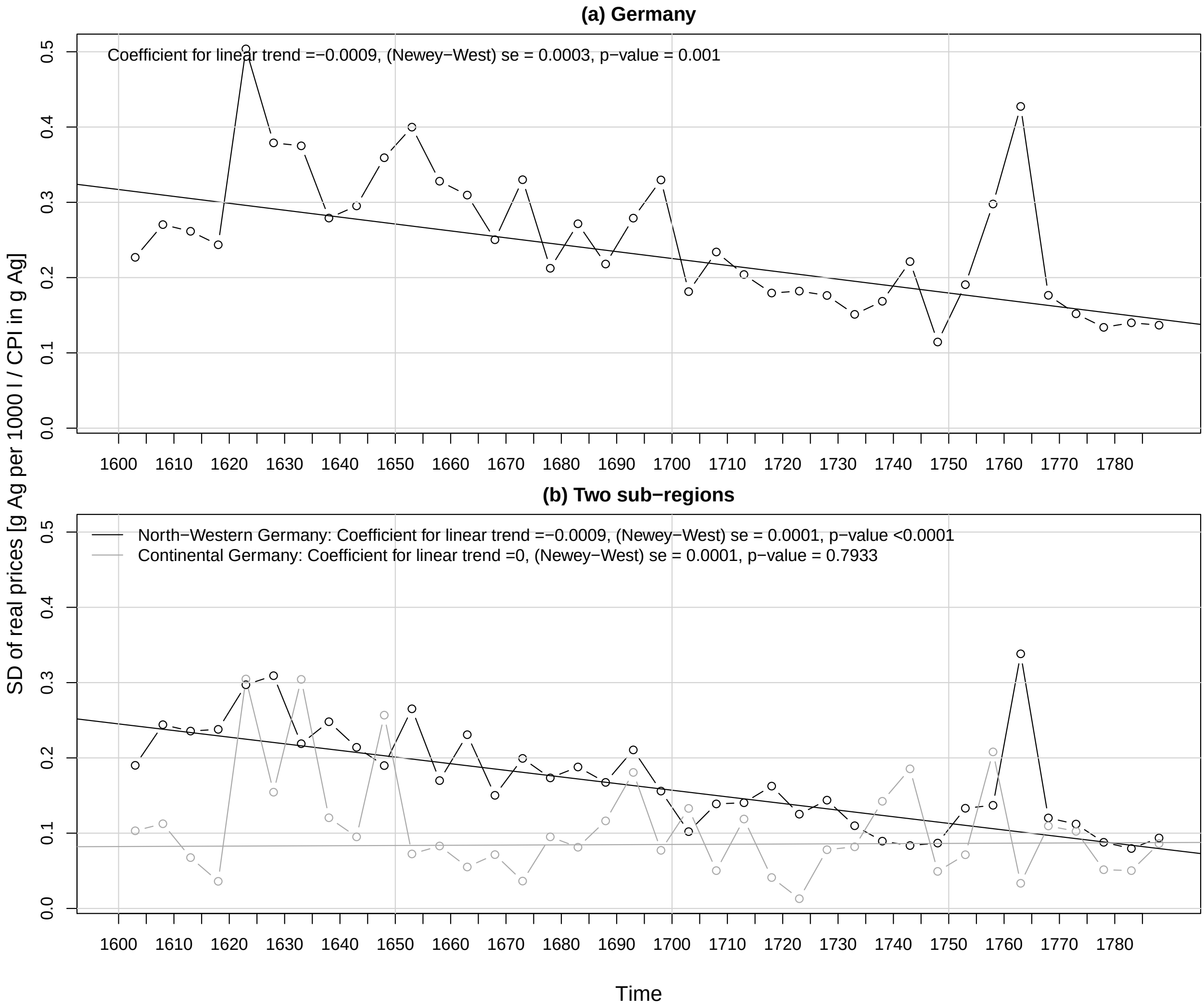

FIGURE S3.18: Inter-urban price dispersion 1601/05–1786/90. Cross-sectional standard deviation of real 5-year-mean-prices, rye (stable sample, 9 cities). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Regressions for linear trends 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 SA3.2.

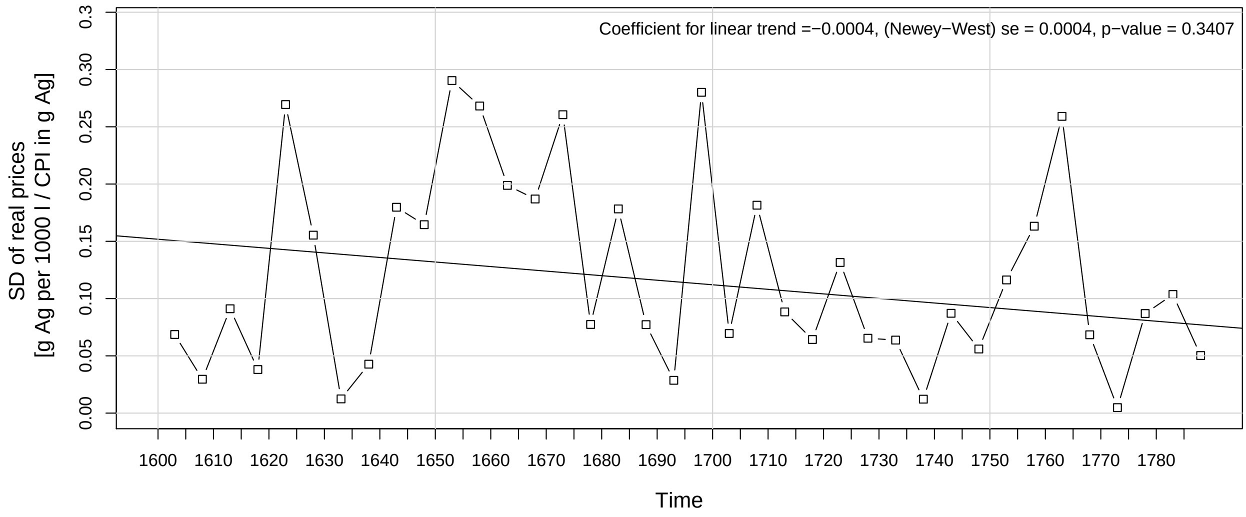

FIGURE S3.19: Inter-regional price dispersion 1601/05–1786/90. Between-region standard deviation (square root of variance between North-Western and continental Germany); real 5-year-mean-prices, rye (stable sample, 9 cities). Regression for linear trend includes dummy variables for Thirty Years' War (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 SA3.2.

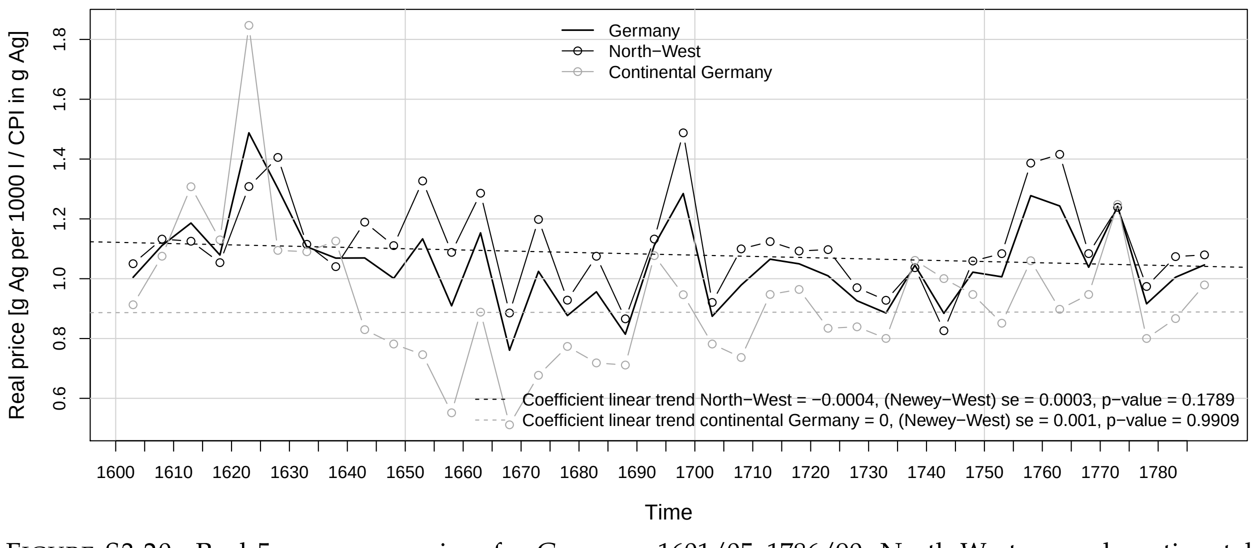

FIGURE S3.20: Real 5-year-mean-prices for Germany 1601/05–1786/90, North-Western and continental Germany, rye (stable sample, 9 cities). Regressions for linear trends include dummy variables for Thirty Years' War (1618–48), the Kipper and Wipper inflation (1620–23) and the Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

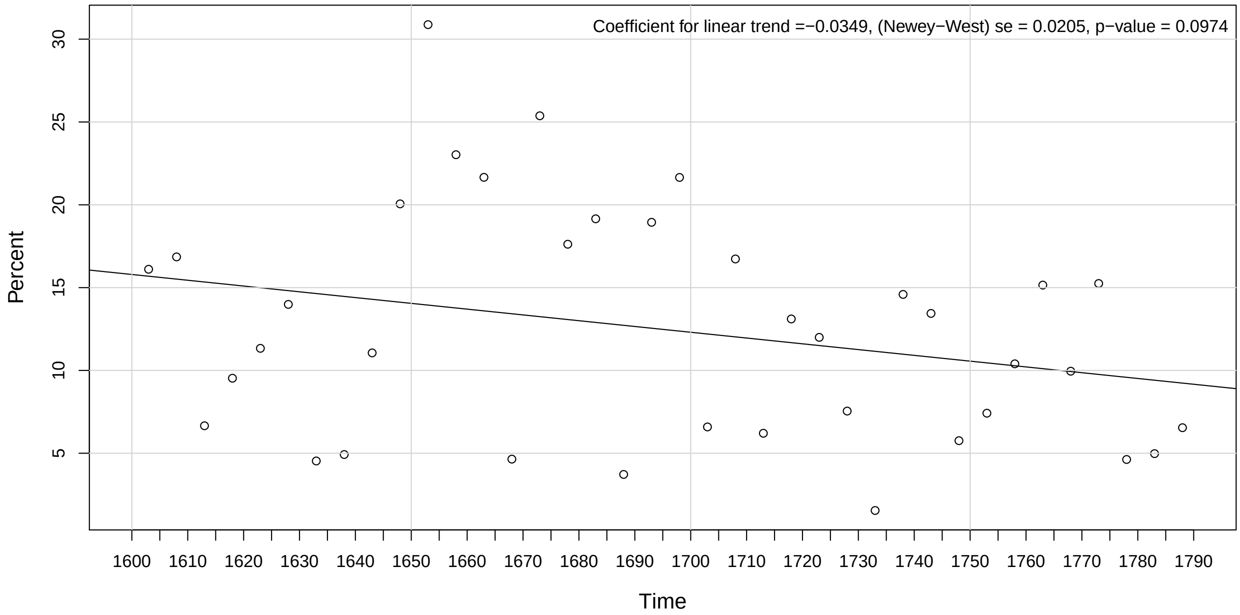

FIGURE S3.21: Volatility of aggregate real rye price Germany 1601/05–1786/90 (stable sample, 9 cities). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

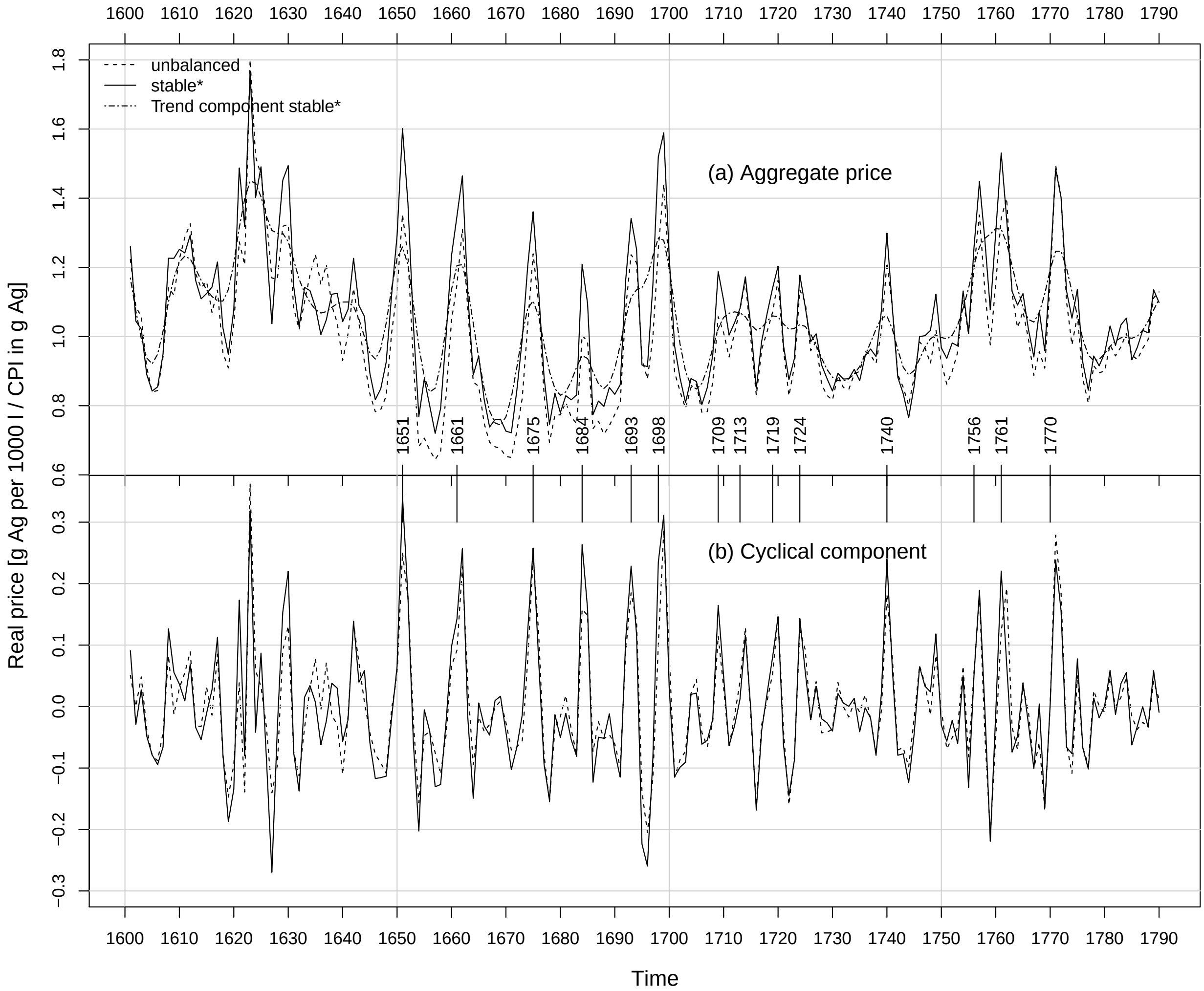

FIGURE S3.22: Aggregate real rye price Germany. Stable sample includes 9 cities, 1601–1790. Trend (shown for stable sample series only) and cyclical component from Hodrick-Prescott-Filter, $\lambda = 6.25$ (Ravn and Uhlig, 2002). Vertical lines and given years on upper horizontal axis in panel (b) mark 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 SA3.2.

SA3.13 Additional results based on the river systems Elbe and Rhine

Sample description: The four sub-regions outlined in the main paper contain the following cities. Elbe region: Hamburg, Lüneburg, Dresden, Magdeburg, Halle, Berlin and also Breslau. Due to the *Friedrich-Wilhelm* canal that was finalized in 1668 and established a connection between the rivers Oder and Spree (Berlin), Breslau is also allocated to the latter group.
'North-East': Northern cities are: Braunschweig, Celle, Emden, Göttingen, Hannover, Minden, Münster, Osnabrück, Paderborn; Eastern cities are: Altenburg, Halberstadt, Leipzig, Nordhausen, Quedlinburg.
Rhine region: Cologne, Xanten, Speyer, Überlingen as well as Frankfurt, Würzburg and Trier.
'South-West': South: Augsburg, Landshut, Munich, Nuremberg; West: Aachen.

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.

The following figures show the results of the variance decomposition with four sub-regions and two aggregate regions discussed in Section 3.5 of the main paper.

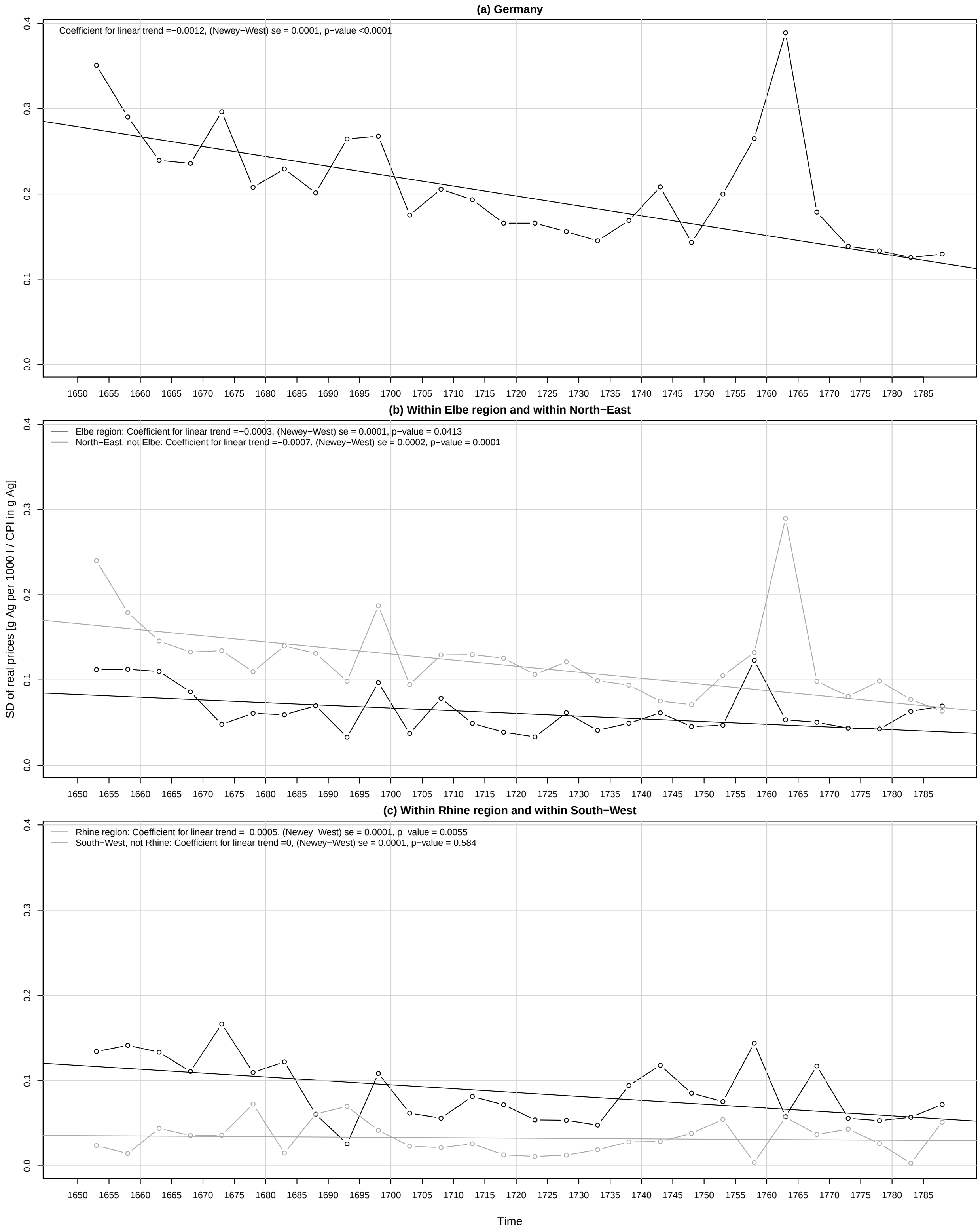

FIGURE S3.23: Inter-urban price dispersion 1651/5–1786/90. Cross-sectional standard deviation of real 5-year-mean-prices, rye (unbalanced sample). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Regressions for linear trends include dummy variable for the Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

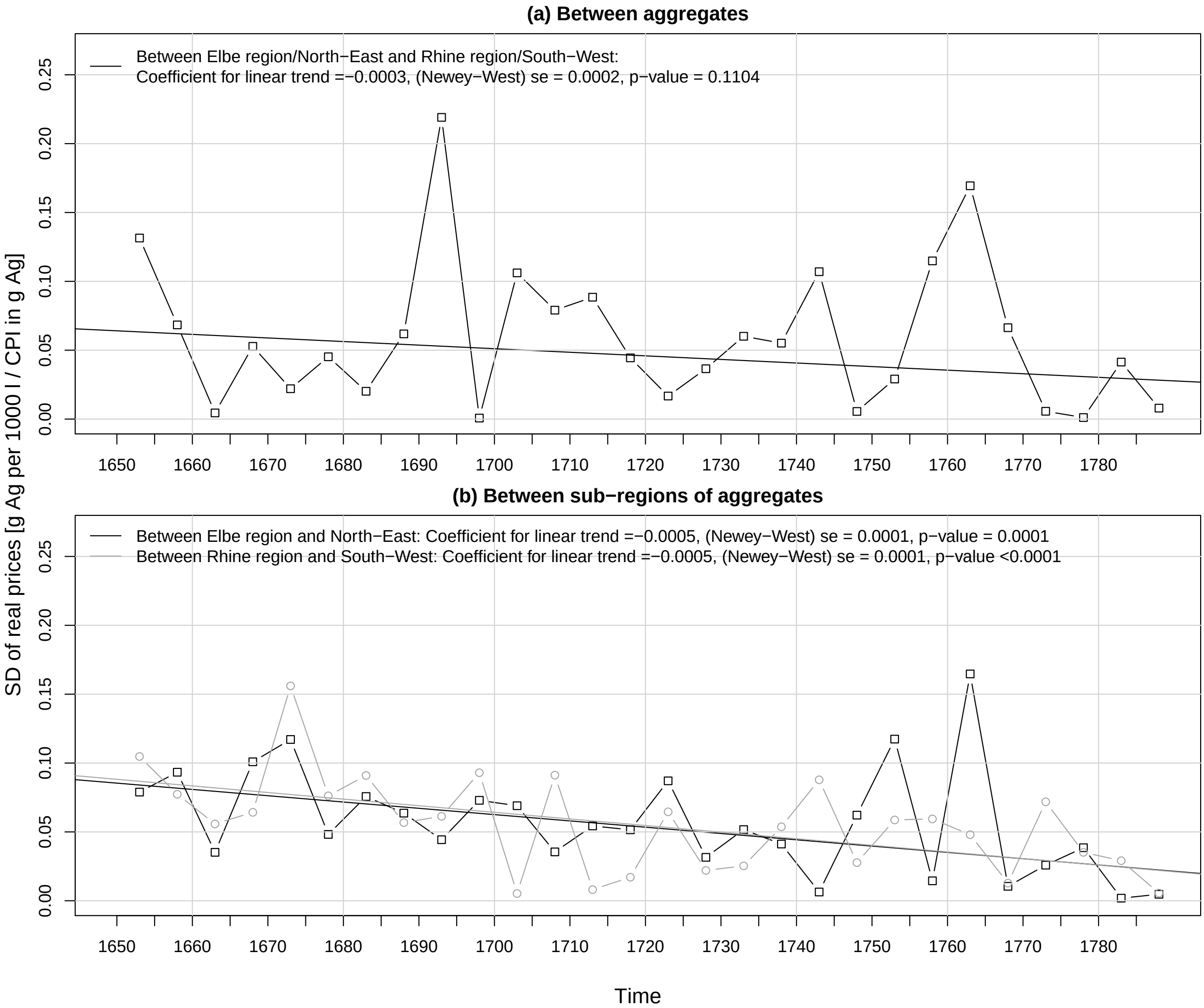

FIGURE S3.24: Inter-regional price dispersion 1651/5–1786/90. Between-region standard deviations (square root of between variances); real 5-year-mean-prices, rye (unbalanced sample). Regressions for linear trends include a dummy variable for the Seven Years' War (1756–1763); for panel (a) also for the crisis of 1690/91. Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

SA3.14 Additional results: split North-Western and continental Germany

In the variance decomposition referring to the river systems of Elbe and Rhine (Section VI 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: Braunschweig, Celle, Emden, Göttingen, Hamburg, Hannover, Lueneburg, Minden, Münster, Osnabrück, Paderborn.
2. West: all non-continental cities south of Xanten which are located in Western Germany: Aachen, Köln, Trier, Xanten.
3. South: all continental cities in Southern Germany: Augsburg, Frankfurt, Landshut, München, Nuremberg, Speyer, Wuerzburg, Ueberlingen.
4. East: Altenburg, Berlin, Breslau, Dresden, Halberstadt, Halle, Leipzig, Magdeburg, Nordhausen, Quedlinburg.

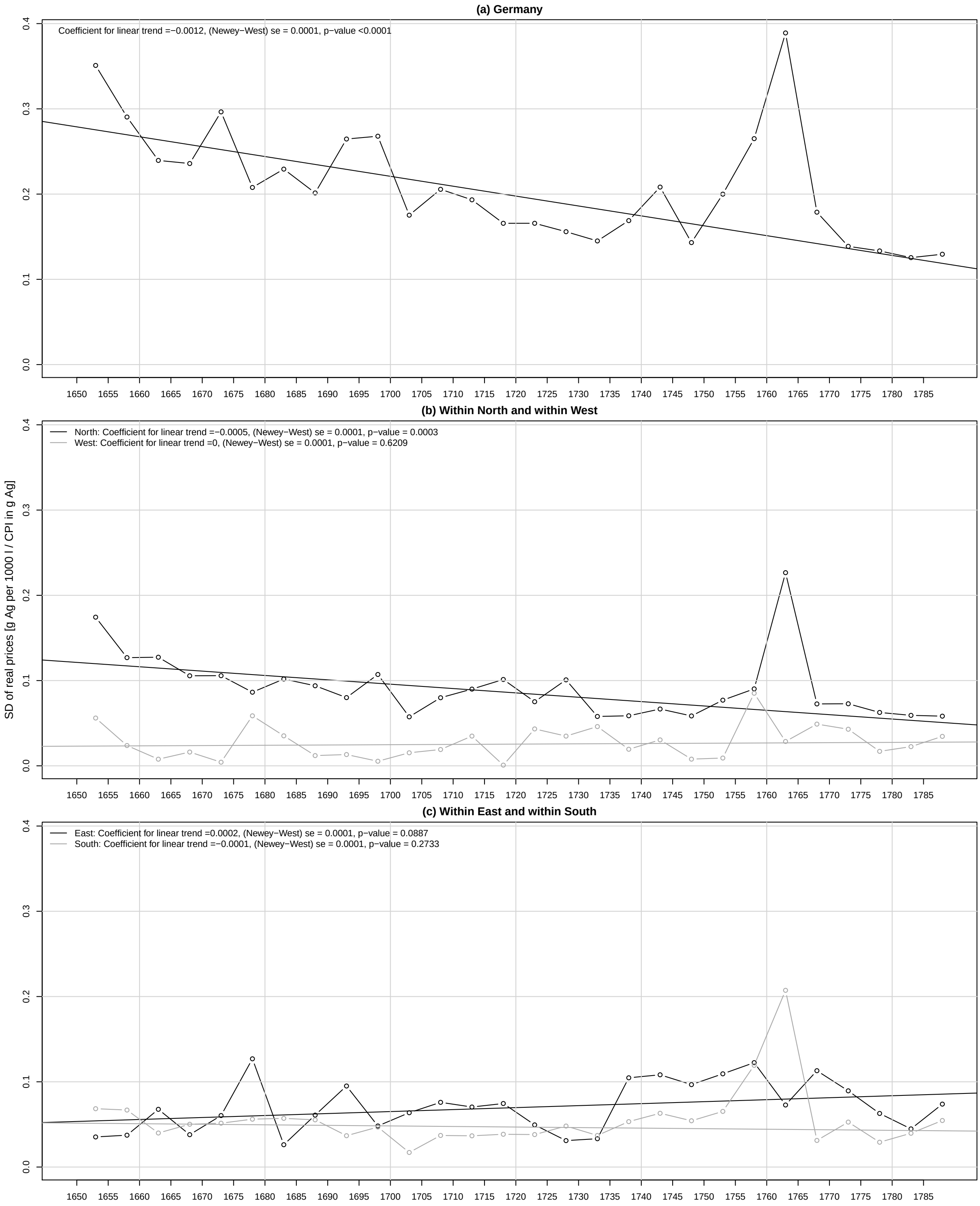

FIGURE S3.25: Inter-urban price dispersion 1651/5–1786/90. Cross-sectional standard deviation of real 5-year-mean-prices, rye (unbalanced sample). Each circle represents a 5-year-period centered at the given year (e.g., circle for year 1653 represents period 1651–55). Regressions for linear trends include dummy variable for Seven Years' War (1756–1763). Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

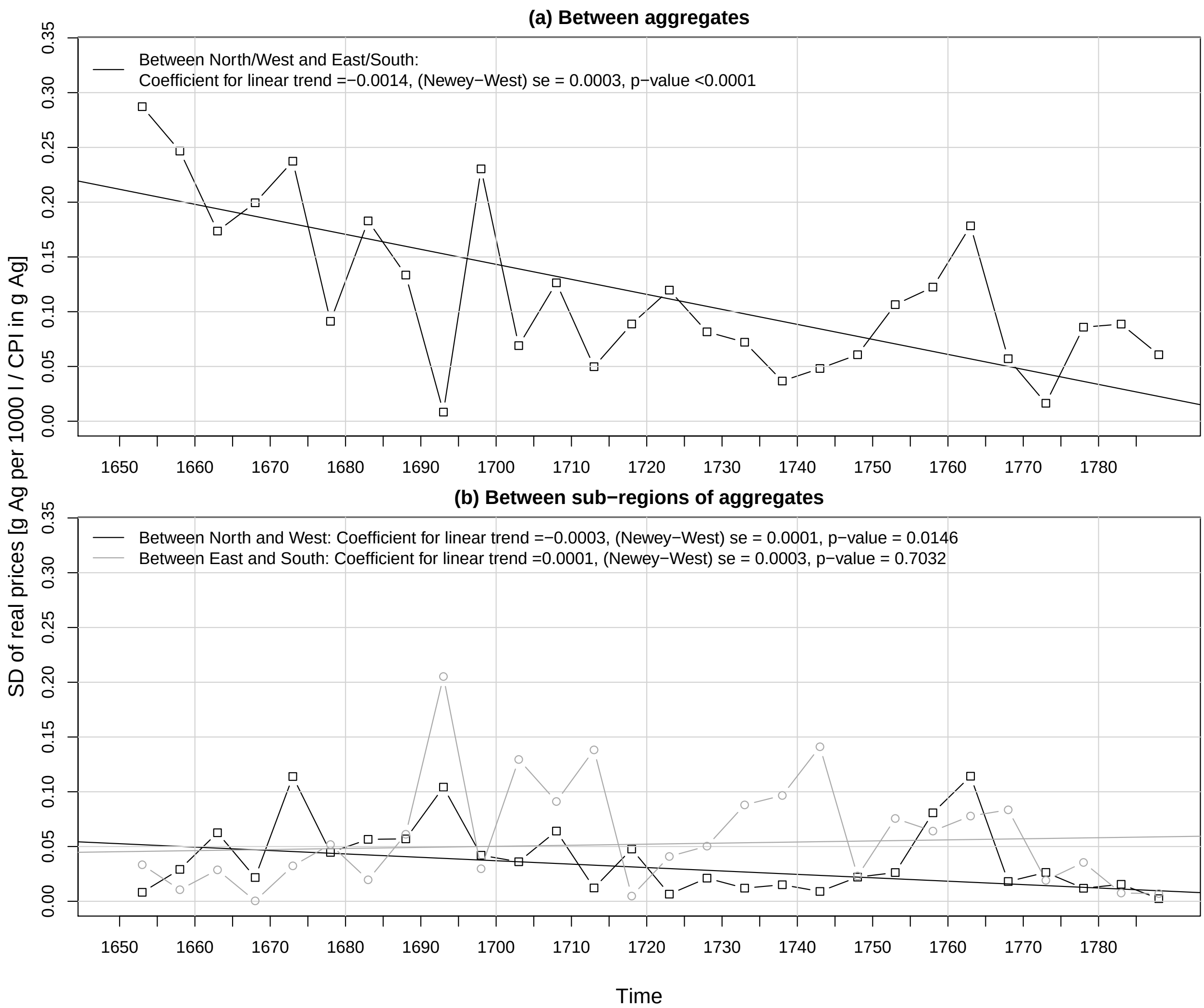

FIGURE S3.26: Inter-regional price dispersion 1651/5–1786/90. Between-region standard deviations (square root of between variances); real 5-year-mean-prices, rye (unbalanced sample). Regressions for linear trends include dummy variables for the Seven Years' War (1756–1763) and the crisis of 1690/91. Nominal prices are deflated with the CPI from Pfister (2017). Data sources: see SA3.2.

SA3.15 Additional results: 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 0.44 for wheat, 0.34 for rye, 0.26 for barley and 0.17 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 S3.31.

TABLE S3.31: 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 National absolute	in %	within NW	within continental	between NW and continental	Annual trend of volatility of aggregate price
Wheat						
St., $N = 11$	-0.0018^{***}	-0.45%	-0.0009^{***}	-0.0003^{***}	-0.0016^{***}	-0.0551^{***}
Unb., $N = 30$	-0.0016^{***}	-0.41%	-0.0009^{***}	-0.0002	-0.0016^{***}	-0.0486^{***}
Rye						
St., $N = 14$	-0.0013^{***}	-0.42%	-0.0005^{***}	0.0001	-0.0015^{***}	-0.0925^{***}
Unb., $N = 33$	-0.0012^{***}	-0.41%	-0.0006^{***}	0.0001	-0.0013^{***}	-0.0902^{***}
Barley						
St., $N = 10$	-0.0009^{***}	-0.34%	-0.0003^{***}	0	-0.0012^{***}	-0.0362
Unb., $N = 29$	-0.0008^{***}	-0.33%	-0.0004^{***}	0.0001	-0.0011^{***}	-0.0407
Oats						
St., $N = 9$	-0.0008^{***}	-0.45%	-0.0003^{**}	-0.0001	-0.0010^{***}	-0.0101
Unb., $N = 26$	-0.0007^{***}	-0.46%	-0.0005^{***}	-0.0001^{*}	-0.0005^{***}	-0.0093

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$; based on Newey-West standard errors. NW: North-West; 'St.' refers to stable sample (≤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.

We find price convergence also for all cereals other than rye at the national level, within North-Western Germany, between this region and the continental parts of Germany, and to a much weaker extent within continental Germany. 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 %' displaying the annual rate of change of the cross-sectional SD in Table S3.31).

Price volatility also declined among all three grains other than rye, but the trend is not significant at conventional levels for barley and much weaker and insignificant 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. 25% 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.

SA3.16 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:

$$\begin{aligned} \bar{p}_t^z &= \frac{(p_{1t}\tau_t) + ... + (p_{Nt}\tau_t)}{N} \\ \bar{p}_t^z &= \frac{p_{1t} + ... + p_{Nt}}{N}\tau_t = \bar{p}_t\tau_t \end{aligned} \tag{S3.10}$$

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

$$\begin{aligned} \textstyle\sum_{i=1}^{N}(p_{it} - \bar{p}_t^z)^2 &= (p_{1t}\tau_t - \bar{p}_t\tau_t)^2 + ... + (p_{Nt}\tau_t - \bar{p}_t\tau_t)^2 \\ &= (p_{1t} - \bar{p}_t)^2\tau_t^2 + ... + (p_{Nt} - \bar{p}_t)^2\tau_t^2 \\ &= [(p_{1t} - \bar{p}_t)^2 + ... + (p_{Nt} - \bar{p}_t)^2]\tau_t^2 \end{aligned} \tag{S3.11}$$

Consequently, the symmetric shock is in both numerator and denominator of the CV and cancels so that the initial CV is obtained:

$$CV_t = \frac{\sqrt{\frac{1}{N-1}[(p_{1t} - \bar{p}_t)^2 + ... + (p_{Nt} - \bar{p}_t)^2]\tau_t^2}}{\bar{p}_t\tau_t} = \frac{\sqrt{\frac{1}{N-1}\sum_{i=1}^{N}(p_{it} - \bar{p}_t)^2}}{\bar{p}_t}. \tag{S3.12}$$

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).

Contrary, the SD increases by factor τ_t for $\tau_t > 1$ and decreases for $0 < \tau_t < 1$. 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.

$$\begin{aligned}\textstyle\sum_{i=1}^{N}(p_{it} - \bar{p}_t^z)^2 &= (p_{1t} + \bar{p}_t(\tau_t - 1) - \bar{p}_t\tau_t)^2 + ... + (p_{Nt} + \bar{p}_t(\tau_t - 1) - \bar{p}_t\tau_t)^2 \\ &= (p_{1t} - \bar{p}_t)^2 + ... + (p_{Nt} - \bar{p}_t)^2.\end{aligned} \tag{S3.13}$$

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}.

$$\begin{aligned}\bar{p}_t^z &= \frac{p_{1t} + ... + p_{Nt}}{N} + \frac{(p_{1t}[\tau_{1t} - 1])}{N} \\ &= \frac{p_{1t} + ... + p_{Nt}}{N} + \frac{p_{1t}r_{1t}}{N} = \bar{p}_t + \frac{p_{1t}r_{1t}}{N}\end{aligned} \tag{S3.14}$$

This shock affects the sum of squared deviations as follows:

$$\begin{aligned}\textstyle\sum_{i=1}^{N}(p_{it}-\bar{p}_t^z)^2 &= (p_{1t}[1+\frac{\tau_{1t}-1}{N}]-[\bar{p}_t+p_{1t}\frac{\tau_{1t}-1}{N}])^2 \\ &+ ... + (p_{Nt}+p_{1t}\frac{\tau_t-1}{N}-[\bar{p}_t+p_{1t}\frac{\tau_{1t}-1}{N}])^2 \\ &= (p_{1t}[1+\frac{\tau_{1t}-1}{N}-\frac{\tau_{1t}-1}{N}]-\bar{p}_t)^2+...+(p_{Nt}-\bar{p}_t)^2.\end{aligned} \quad \text{(S3.15)}$$

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. (S3.14)) 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:

$$\begin{aligned}\textstyle\sum_{i=1}^{N}(p_{it}-\bar{p}_t^z)^2 &= (p_{1t}\tau_{1t}-[\bar{p}_t+\frac{p_{1t}(\tau_{1t}-1)}{N}])^2+...+(p_{Nt}-[\bar{p}_t+\frac{p_{1t}(\tau_{1t}-1)}{N}])^2 \\ &= (p_{1t}[\tau_{1t}-\frac{\tau_{1t}-1}{N}]-\bar{p}_t)^2+...+(p_{Nt}-\bar{p}_t-\frac{p_{1t}(\tau_{1t}-1)}{N})^2.\end{aligned} \quad \text{(S3.16)}$$

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*.

SA3.17 Proof: 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 SA3.18.

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 S3.8).

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 $\bar{p}^z$:

$$\bar{p}^z = \frac{\bar{p}_1 + [\bar{p}_2 + as_{12}] + \bar{p}_3 + ... + \bar{p}_T}{T} = \frac{\bar{p}_1 + \bar{p}_2 + \bar{p}_3 + ... + \bar{p}_T}{T} + \frac{as_{12}}{T} = \bar{p} + \frac{as_{12}}{T}. \tag{S3.17}$$

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 $u(.)$:

$$\begin{aligned} u(.) &= \textstyle\sum_{t=1}^{T}(\bar{p}_t - \bar{p}^z)^2 \\ &= (\bar{p}_1 - \bar{p} - \frac{as_{12}}{T})^2 \quad +(\bar{p}_2 + as_{12} - \bar{p} - \frac{as_{12}}{T})^2 + (\bar{p}_3 - \bar{p} - \frac{as_{12}}{T})^2 + ... \\ &+ (\bar{p}_T - \bar{p} - \frac{as_{12}}{T})^2 \\ &= ... \qquad\qquad\qquad +(\bar{p}_2 - \bar{p} + \frac{(Ta-a)s_{12}}{T})^2 + ... \end{aligned} \tag{S3.18}$$

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 $t = 2$ (see SA3.18 for details). 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}. In other words, we assume no intertemporal arbitrage, e.g., through storage, which would reduce volatility.

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. (S3.18) simplifies to (details in eq. (S3.21)):

$$u(.) = (T-1)\frac{a^2 s_{12}^2}{T}. \quad \text{(S3.19)}$$

The expression for $u(.)$ in eq. (S3.19) is clearly positive, because $T > 1$. Based on this result and further analysis, SA3.18.1 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:

$$\frac{\partial u/\bar{p}^z}{\partial a} = \frac{2(T-1)\frac{as_{12}^2}{T}\cdot[\bar{p}+\frac{as_{12}}{T}] - [(T-1)\frac{a^2 s_{12}^2}{T}]\cdot\frac{s_{12}}{T}}{[\bar{p}+\frac{as_{12}}{T}]^2} > 0, \quad \text{(S3.20)}$$

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 SA3.18.1). Thus, a larger number of cities N, which are available for spatial arbitrage, means that $a = \frac{1}{N}$ decreases and reduces volatility. In SA3.18.2, 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, a larger a means a larger volatility; $a = 1$ (implying $n = N$) corresponds to the case of a perfectly symmetric shock.

SA3.18 Supporting information: spatial arbitrage reduces aggregate volatility

We first provide additional details on the case of an absolute shock, which we discuss in SA3.17. Second, we repeat the analysis for the case of a proportional shock.

SA3.18.1 Absolute shock

Additional details on the flat line assumption

Figure S3.27 illustrates the 'flat line' assumption.

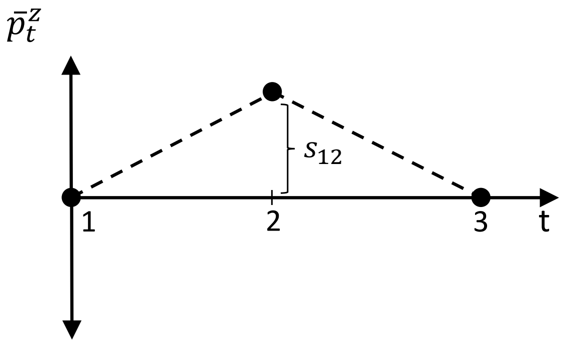

FIGURE S3.27: Graphical illustration of analytical framework. $\bar{p}_t^z$: cross-sectional average price in year t inlcuding the shock s_{12} to city 1 in year 2. Source: own representation.

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 S3.27, 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.

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

$$\begin{aligned}
u(.) &= (\bar{p} - \bar{p} - \frac{as_{12}}{T})^2 + \Big[\bar{p} - \bar{p} + \frac{(Ta-a)s_{12}}{T}\Big]^2 + (\bar{p} - \bar{p} - \frac{as_{12}}{T})^2 + ... \\
&+ (\bar{p} - \bar{p} - \frac{as_{12}}{T})^2 \\
&= (T-1)\cdot(-\frac{as_{12}}{T})^2 + \Big[\frac{(Ta-a)s_{12}}{T}\Big]^2 \\
&= (T-1)\cdot\frac{a^2 s_{12}^2}{T^2} + \frac{(Ta-a)^2\cdot s_{12}^2}{T^2} \\
&= (T-1)\cdot\frac{a^2 s_{12}^2}{T^2} + \frac{(T^2a^2 - 2Ta^2 + a^2)\cdot s_{12}^2}{T^2} \\
&= (T-1)\cdot\frac{a^2 s_{12}^2}{T^2} + \frac{(T^2 - 2T + 1)\cdot a^2 \cdot s_{12}^2}{T^2} \\
&= (T - 1 + T^2 - 2T + 1)\cdot\frac{a^2 s_{12}^2}{T^2} \\
&= (T^2 - T)\frac{a^2 s_{12}^2}{T^2} \\
&= (T-1)\frac{a^2 s_{12}^2}{T}.
\end{aligned} \tag{S3.21}$$

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:

$$\frac{\partial u/\bar{p}^z}{\partial s_{12}} = \frac{\overbrace{2(T-1)\frac{a^2 s_{12}}{T}}^{u'>0}\cdot\overbrace{[\bar{p} + \frac{as_{12}}{T}]}^{\bar{p}^z>0} - \overbrace{[(T-1)\frac{a^2 s_{12}^2}{T}]}^{u>0}\cdot\overbrace{\frac{a}{T}}^{\bar{p}^{z\prime}>0}}{\underbrace{[\bar{p} + \frac{as_{12}}{T}]^2}_{\bar{p}^{z2}>0}} > 0, \tag{S3.22}$$

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.[19] 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.

Effect of spatial arbitrage on aggregate volatility

Additional step on the simplification of the numerator

The numerator can be simplified to:

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

[19]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.

SA3.18.2 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.

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

$$\begin{aligned} \bar{p}^z &= \frac{\bar{p}_1 + [\bar{p}_2(1 + ar_{12})] + \bar{p}_3 + ... + \bar{p}_T}{T} = \frac{\bar{p}_1 + \bar{p}_2 + \bar{p}_3 + ... + \bar{p}_T}{T} + \frac{ar_{12}\bar{p}_2}{T} \\ &= \bar{p} + \frac{ar_{12}\bar{p}_2}{T}. \end{aligned} \tag{S3.23}$$

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

$$\begin{aligned} u(.) &= \textstyle\sum_{t=1}^{T} (\bar{p}_t - \bar{p}^z)^2 \\ &= (\bar{p}_1 - \bar{p} - \frac{ar_{12}\bar{p}_2}{T})^2 + \Big[\bar{p}_2 + ar_{12}\bar{p}_2 - \bar{p} - \frac{ar_{12}\bar{p}_2}{T}\Big]^2 \\ &+ (\bar{p}_3 - \bar{p} - \frac{ar_{12}\bar{p}_2}{T})^2 + ... + (\bar{p}_T - \bar{p} - \frac{ar_{12}\bar{p}_2}{T})^2. \end{aligned} \tag{S3.24}$$

The term for the second period in [.] can be simplified to $\Big[\frac{\bar{p}_2 \cdot (T + Tr_{12}a - ar_{12})}{T} - \bar{p}\Big]^2$. 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. (S3.24):

$$\begin{aligned} u(.) &= (\bar{p} - \bar{p} - \frac{ar_{12}\bar{p}}{T})^2 + \Big[\frac{\bar{p} \cdot (T + Tr_{12}a - ar_{12} - T)}{T}\Big]^2 + (\bar{p} - \bar{p} - \frac{ar_{12}\bar{p}}{T})^2 + ... \\ &+ (\bar{p} - \bar{p} - \frac{ar_{12}\bar{p}}{T})^2 \\ &= (T-1) \cdot (-\frac{ar_{12}\bar{p}}{T})^2 + \Big[\frac{\bar{p} \cdot ar_{12}(T-1)}{T}\Big]^2 \\ &= (T-1) \cdot \frac{a^2 r_{12}^2 \bar{p}^2}{T^2} + \frac{\bar{p}^2 \cdot a^2 r_{12}^2 (T-1)^2}{T^2} \\ &= \Big[(T-1) + (T-1)^2\Big] \cdot \frac{a^2 r_{12}^2 \bar{p}^2}{T^2} \\ &= \Big[T - 1 + T^2 - 2T + 1\Big] \cdot \frac{a^2 r_{12}^2 \bar{p}^2}{T^2} \\ &= (T-1) \cdot \frac{a^2 r_{12}^2 \bar{p}^2}{T}. \end{aligned} \tag{S3.25}$$

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:

$$\frac{\partial u/\bar{p}^z}{\partial r_{12}} = \frac{\overbrace{2(T-1)\frac{a^2 r_{12}\bar{p}^2}{T}}^{u'>0} \cdot \overbrace{[\bar{p} + \frac{ar_{12}\bar{p}}{T}]}^{\bar{p}^z>0} - \overbrace{[(T-1)\frac{a^2 r_{12}^2 \bar{p}^2}{T}]}^{u} \cdot \overbrace{\frac{a\bar{p}}{T}}^{\bar{p}^{z\prime}>0}}{\underbrace{[\bar{p} + \frac{ar_{12}\bar{p}}{T}]^2}_{\bar{p}^{z2}>0}} > 0, \tag{S3.26}$$

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:

$$\frac{\partial u/\bar{p}^z}{\partial a} = \frac{\overbrace{2(T-1)\frac{ar_{12}^2\bar{p}^2}{T}}^{u'\text{different}} \cdot [\bar{p} + \frac{ar_{12}\bar{p}}{T}] - [(T-1)\frac{a^2 r_{12}^2 \bar{p}^2}{T}] \cdot \overbrace{\frac{r_{12}\bar{p}}{T}}^{\bar{p}^{z\prime}\text{different}}}{[\bar{p} + \frac{ar_{12}\bar{p}}{T}]^2} > 0, \tag{S3.27}$$

because the numerator can be simplified to:

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

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

Chapter 4

War, state growth and market integration in the German lands, 1780–1830

Authors: Hakon ALBERS* and Ulrich PFISTER†

Note: This chapter is a revised version of the paper presented at the Annual Meeting of the Economic History Association: Albers and Pfister (2018).

Abstract: The Napoleonic Wars transformed Germany from a plethora of dominions with limited sovereignty to a federation of about 40 states in 1815. We test the hypothesis that state growth promoted market integration through larger internal markets with a novel data set of rye prices in 33 German towns. Price gaps between markets that belonged to different polities before the wars and that both became part of the same state after the wars fell significantly. We find this effect for Bavaria and a region that was part of the Kingdom of Westphalen during the wars and was later integrated into Prussia.

Keywords: state growth, market integration, Napoleonic Wars, economic development.

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"A large obstacle of interior transport are the restrictions, which one German state enacted against another one so far. Simply because the number of states was reduced, trade will benefit." (Winkopp, 1806, 66)

4.1 Introduction

Trade and institutions are important factors for economic development. The interplay of business-friendly institutions and Atlantic trade, for example, was pivotal for the rise of Western European countries like England or the Netherlands from 1500 to 1800 (Acemoglu et al., 2005). Continental

*__Corresponding Author__. Martin Luther University Halle-Wittenberg, Department of Economics, Universitätsring 3, D-06108 Halle, Germany

†University of Münster, Institute of Economic and Social History, Germany.

countries such as Germany, however, were rather isolated from Atlantic Trade. At the turn to the nineteenth century, continental Europe witnessed the Revolutionary and Napoleonic Wars (1792–1815). In the German lands, the wars led to the dissolution of the Holy Roman Empire, which arguably altered the institutional framework of this important pre-industrial economy. In a debated article, Acemoglu et al. (2011) relate institutional change which they approximate with an index of French reforms to regional city growth in Germany (cf. Kopsidis and Bromley, 2016). More specifically, Acemoglu et al. (2011) argue that institutional reforms enacted in territories under French rule or influence at the beginning of the nineteenth century had a beneficial impact on economic development after 1850.

Keller and Shiue (2016) employ a similar framework as Acemoglu et al. (2011) but combine the analysis of city growth with the one of price gaps as a proxy for market integration. Their findings suggest that institutional change impacted on economic growth both directly and indirectly through market integration. According to Keller and Shiue, the indirect channel prevailed over the direct channel; institutional reforms enacted at the beginning of the nineteenth century benefitted economic development primarily by fostering market integration, which in turn increased allocative efficiency.

While we share the view of Keller and Shiue that the wars impacted on market integration, we stress a different causal mechanism. We argue that the main consequence of the wars relevant for market integration was the creation of bigger states with larger internal markets. Our hypothesis is based on a theoretical model by Alesina and Spolaore (1997): The size of internal markets and the efficiency of government are both related to state size. While too large territories can have disadvantages caused by large and heterogeneous populations (e.g., secessions), too small states might have problems to cover the fixed costs of nonrival public goods (such as efficient taxation systems) and may suffer from market segmentation. For our case, we conjecture that, as a consequence of the boundary changes following from the Revolutionary and Napoleonic Wars, a few previously small German states expanded significantly and approached optimal state size from below. Much larger internal markets and new systems of external tariffs, which epitomized new rules—institutions—governing trade, were a direct consequence.

In the German lands, the number of political units shrank from about 200 polities, which all had limited sovereignty due to their subjection to Imperial law, to about 40 in 1815 (Kunz, 2004; 2014; see also Shiue, 2005, 131–2). 60 percent of the German population were now citizens of one of the two largest states, Prussia and Bavaria (Fertig et al., 2018). This had two effects: First, more trade partners belonged to the same territory and heterogeneous internal tariffs were abolished, which increased the size of internal markets. Second, public authorities in large, sovereign states were now better capable to provide public goods. Along with the development of modern states uniform systems of external tariffs were created. An additional case in point is infrastructure development: At the end of the eighteenth century there were few paved roads in Germany; after 1815, public authorities inaugurated programmes of state road construction (Müller, 2000; Uebele and Gallardo-Albarrán, 2015). In Bavaria, for example, the system of main routes (*Hauptstraßen*) was improved to a Bavaria-wide network after 1819, and the state became responsible for maintaining roads (Schäfer, 1985, 312; Gall, 2013).

Through these two channels—creation of internal markets and improved supply of public goods—the development of modern, much larger states benefitted market integration. We carry out a summary test of this hypothesis by implementing a difference-in-differences (DD) approach on a novel data set of rye prices for 33 markets in 1780–1830. We thereby follow a well-established tradition of employing grain price gaps as a proxy for market integration (e.g., Federico, 2012; Chilosi et al., 2013; Keller and Shiue, 2014).

We find that price gaps between markets that belonged to different polities before the wars and that both became part of Bavaria fell by about 62 percent between the 1780s and the 1820s. The effects in Prussia are heterogeneous and there is no uniform reduction of price gaps in newly included territories. But for those city-pairs that belonged to the former Kingdom of Westphalen, a French satellite state, and then became part of Prussia after the wars, we find a reduction of price gaps by about 41 percent. French presence as such had no uniform effect on price gaps. The latter finding indicates that reforms such as the introduction of modern, inclusive commercial law had no immediate impact on grain trade. State growth created internal markets and improved the supply of trade-related public goods. Thus, the effect of state growth, on market integration constituted the main channel through which the Revolutionary and Napoleonic Wars impacted on Germany's subsequent economic development.

The remainder of this study is organized as follows: We begin with an overview of the evolution of major institutions governing grain trade from the eighteenth to the early nineteenth century, namely, tariff regimes and grain market regulation. The following sections in turn describe data and empirical strategy, present our results and a summary of the robustness checks. Section 4.7 summarizes and interprets the results.

4.2 The transformation of institutions governing trade: tariff regimes and the regulation of grain markets

The formation of modern states in the wake of the Revolutionary and Napoleonic wars changed the institutions governing grain markets in two major respects: First, heterogeneous levies on trade gave way to a differentiation between indirect taxes and unified systems of external tariffs, which fostered the creation of internal markets. Second, liberalization of grain markets facilitated wholesale trade of grain. This section provides an overview of these processes of institutional change.

4.2.1 Tariff regimes

Until the late eighteenth century, trade in the German lands faced a plethora of internal tolls and levies. Tariff regulation was heterogeneous, and in some territories customs were blended with the excise, that is, indirect taxes on consumer goods levied at town gates. It has been speculated that in 1790 Germany had about 1800 internal tariff schemes. Along major rivers the average distance between two toll houses was at the order of magnitude of 10 to 20 kilometres (Stolz, 1954, 26, 33–5; Berding, 1980, 524–6). Consequently, the comparative advantage of individual trade routes over others was completely taxed away. On the Rhine, for instance, numerous customs stations, rights of staple and compulsory transfer of freight from one corporation of boatsmen to the next in Cologne and Mainz led to a massive rise in trade costs. Thus, a sizeable portion of north-south trade was displaced eastward to the Elbe river and the overland route from Magdeburg to Nürnberg and farther south to Augsburg (Newman, 1985, 74–5; Spaulding, 2011, 209–13).

Nevertheless, already during the second half of the eighteenth century individual states attempted to reform their excise and tariff regimes in the view of unifying their territory in economic terms and of bringing in line political frontiers with tariff borders (Stolz, 1954, 29–30, 35–6). The most important case is Bavaria, which introduced a unified system of external tariffs and concentrated the collection of the excise in major towns in 1765. The impact was limited, though, because of two reasons: First, the initial reform related only to the old part of the kingdom in the south and not to the recently acquired parts in the north, where the power of the crown was limited. Second, there existed numerous privileges of estates located outside Bavaria, both with respect to entitlements to and exemptions from tolls. Thus, attempts at introducing a unified tariff regime produced numerous conflicts that led to lawsuits in Imperial courts (Berding, 1980, 526). While some of these holes in the new system could be filled towards the end of the eighteenth century, only the dissolution of the Empire in 1806 and the accession to full sovereignty rendered it possible to complete the transformation of the tariff regime (Häberle, 1974). The crown of Prussia pursued a similar policy in several territories. Specifically, in the Duchy of Mark in Westphalia, public authorities strongly reduced the excise in 1791 and abolished internal tariffs in 1796 (Gorissen, 2002, 106). Still, the different parts of the kingdom remained under separate tariff regimes; east of the Elbe River there existed 67 tariff schemes at the beginning of the nineteenth century (Jacobs and Richter, 1935, 50).

War and French expansion were crucial for transforming tariff regimes. By 1795, France had conquered the left bank of the Rhine, and annexation was legalized with the Treaty of Lunéville in 1801. In the framework of the Imperial decree (*Reichsdeputationshauptschluss*) of 1803, France forced the abolition of all levies on river navigation on the Rhine. In the following year, France and the Empire concluded a treaty (*Octroi*) that internationalized navigation on the Rhine, instituted joint jurisdiction over river navigation and created an organization with twelve offices that fulfilled the function of a police and collected a fee destined to finance its own maintenance. Moreover, the *Octroi* contributed to a unification of measures and weights in its orbit and abolished rights of staple. In 1815, the Congress of Vienna created the Central Commission for Navigation on the Rhine, the first international governmental organization, as a successor of the *Octroi* of 1804. Further liberalization of river navigation, however, particularly the abolition of the corporate organization of river transport, occurred only with the adoption of the Convention of Mainz in 1831 (Borgius, 1899, 102–3; Diefendorf, 1980, 165–80; Spaulding, 2011, 213–8; Thiemeyer and Tölle, 2011, 180–4).

Implementation of the *Octroi* went hand in hand with a strong increase of freight quantities. Whereas the Continental system, which was inaugurated in 1806, led to a collapse of upstream trade after 1807, downstream freight quantities remained buoyant over the period 1806–1813. In the late 1800s, the rural hinterland of Cologne shipped sizeable quantities of grain to Holland, whereas this trade had been marginal in the early 1790s according to the ledgers of the Admiralty of Amsterdam that survive from that period (Eichhoff, 1814, 62–7; van Nierop, 1917; Spaulding, 2013; 2015).

The *Octroi* regime constituted the first element in the transformation of the Rhine from a location that offered a great number of points suitable for extracting levies from passers-by to a line posing barriers to trade between the east and the west bank. The second element consisted in a move of France towards protectionism and a tight regulation of international trade. In 1798 the tariff frontier was moved from the Maas to the Rhine, which meant that it became very difficult to export manufactures from the east bank to the west bank of the Rhine for about fifteen years (Rowe, 2015). Additionally, French authorities forbade the export of grain in 1798. The ensuing years saw several shifts between a liberal and a prohibitive stance towards grain trade; prohibition included high tariffs on exports and temporal trade bans (Springer, 1926, 367–8; Schultheis-Friebe, 1969, 254–61; Rowe, 2003, 197).

In the years following the dissolution of the Empire in 1806, most middle-sized German territories that were members of the French-sponsored Federation of the Rhine enacted laws instituting unified systems of external tariffs: the Grand-Duchy of Berg in 1806/08, Bavaria in 1807, Württemberg in 1808, the Kingdom of Westphalen in 1811 and Baden in 1812. The loss of revenue from taxing trade on the Rhine may have put a pressure towards this transformation, and the accession to sovereignty of most states provided reform-oriented bureaucracies with the necessary room for action. Compared with earlier schemes of internal customs the new regimes were uniform in the sense that they were rule-based, introduced a centralized customs administration, and abolished individual and corporate privileges and exemptions (Berding, 1980, 530–3). Prussia followed this step only after the war in the framework of a comprehensive programme aiming at fiscal rationalization and consolidation. Policy formulation benefitted from the experience of the Kingdom of Westphalen, a French satellite state, as the responsible minister, Hans von Bülow, had previously been minister of finance there in 1808–11. The Prussian tariff law enacted in 1818 (see also Shiue, 2005, 135), while stipulating separate tariff rates for the western and the eastern provinces, abolished all internal customs and tolls and thus created a large internal market. In the new provinces in the west the transition to external tariffs was easy because the reforms carried out under French rule during the war period had left behind a fiscal system resting solely on direct taxes. East of the Elbe trade had been hitherto taxed through a mixture of tariffs, tolls and the excise. Hence, the creation of a system of external tariffs in 1818 was followed by further legislation that replaced the excise with new indirect taxes (Ohnishi, 1973, 11–6, 30–1, 44–5; Siegert, 2001, 106–15; Spoerer, 2004, 47–51).

4.2.2 Regulation of grain markets

Until the early nineteenth century political authorities reacted to harvest failures with interventions on grain markets, such as export tariffs, quantitative restrictions or outright bans on exports. In southwestern Germany, where political fragmentation was particularly marked, public authorities cooperated through the Imperial circle (*Reichskreis*) of Swabia. The latter allocated export quotas and established a complex supervisory mechanism that included export certificates and procedures for their control. A predominantly rural territory such as the Palatinate, which possessed only relatively small towns, even tried to suppress exports into the neighboring Imperial cities on a permanent basis from 1775 (Borgius, 1899, 35–40; Huhn, 1987, 39–42, 58–66; Göttmann, 1991, part 2). The kingdom of Prussia pursued a liberal trade policy in its north-eastern possessions (the hinterland of Königsberg and Gdansk), where export-oriented grain production dominated, and in the western territories, which were too small to make a policy oriented on autarchy feasible. In the central provinces and Silesia, by contrast, the central government exerted tight control over grain trade using quantitative restrictions and periodic trade bans, except for an abortive liberal experiment in 1786–90 (Skalweit, 1931, 10–45, 165–210). Despite mounting critique from free-traders, export restrictions or bans continued to be practiced during the Revolutionary and Napoleonic wars (see above on French grain trade policy on the west bank of the Rhine) and the Tambora crisis of 1816/17. In the following food shortages from 1830 to the mid-1850s, by contrast, measures of this type fell in disuse (Borgius, 1899, 91–2; Huhn, 1987, 53; Bass, 1991, 122–3, 138).

Under the old regime, long-distance trade in grain was also hampered by market regulation that protected urban retail trade. Many cities and territories had enacted legislation that forced trade into formal markets in towns (staple rights), placed restrictions on wholesale trade, forbade forward contracts and reserved the morning hours to transactions of local consumers and craftsmen processing grain (bakers, millers). Only during a short time span between noon and the closure of the markets were wholesale traders allowed to conclude contracts (Borgius, 1899, 39–40, 45–6, 61–6; Skalweit, 1931; Huhn, 1987, 46, 80).

The era of the Revolutionary and Napoleonic Wars saw the onset of a profound liberalization of grain market regulation. In Baden, the concentration of grain trade on formal urban markets was abolished in 1807, in Bavaria in 1813. After the end of war, reform proceeded in a more gradual pace. In the Prussian province of Westphalia liberalized markets handling inter-regional grain trade coexisted with towns that reintroduced the restrictive legislation of the pre-revolutionary era. Only in 1845 the Prussian industrial code enforced universal deregulation. Around the same time we find evidence of an emerging trade with forward contracts on grain that spanned over a large area between Stettin, Berlin, Hamburg, Cologne and Mainz (Borgius, 1899, 84–8; Huhn, 1987, 80–5; Bass, 1991, 59).

4.3 Data and stylized facts of grain prices

To test the hypothesis that war-related state growth led to market integration, we construct a novel dataset of grain prices covering 33 towns in 1780–1830. We focus on rye because this was the staple food crop in most parts of Germany and convert prices to grams of silver per litre. Data for 25 towns are from Albers et al. (2018); series for additional eight towns are documented in Supplementary Appendix SA4.1. The map in Figure 4.1 shows the markets represented in our data set. Additionally, the map marks city-pairs that became newly included into the territory of Bavaria (black lines), Hanover (black dashed lines), and Prussia (grey dot-dashed lines) during the war period. In addition, we highlight city-pairs that were newly included into Prussia and had been part of the Kingdom of Westphalen (black dot-dashed lines). Pairs that are part of the control group are marked with grey lines.

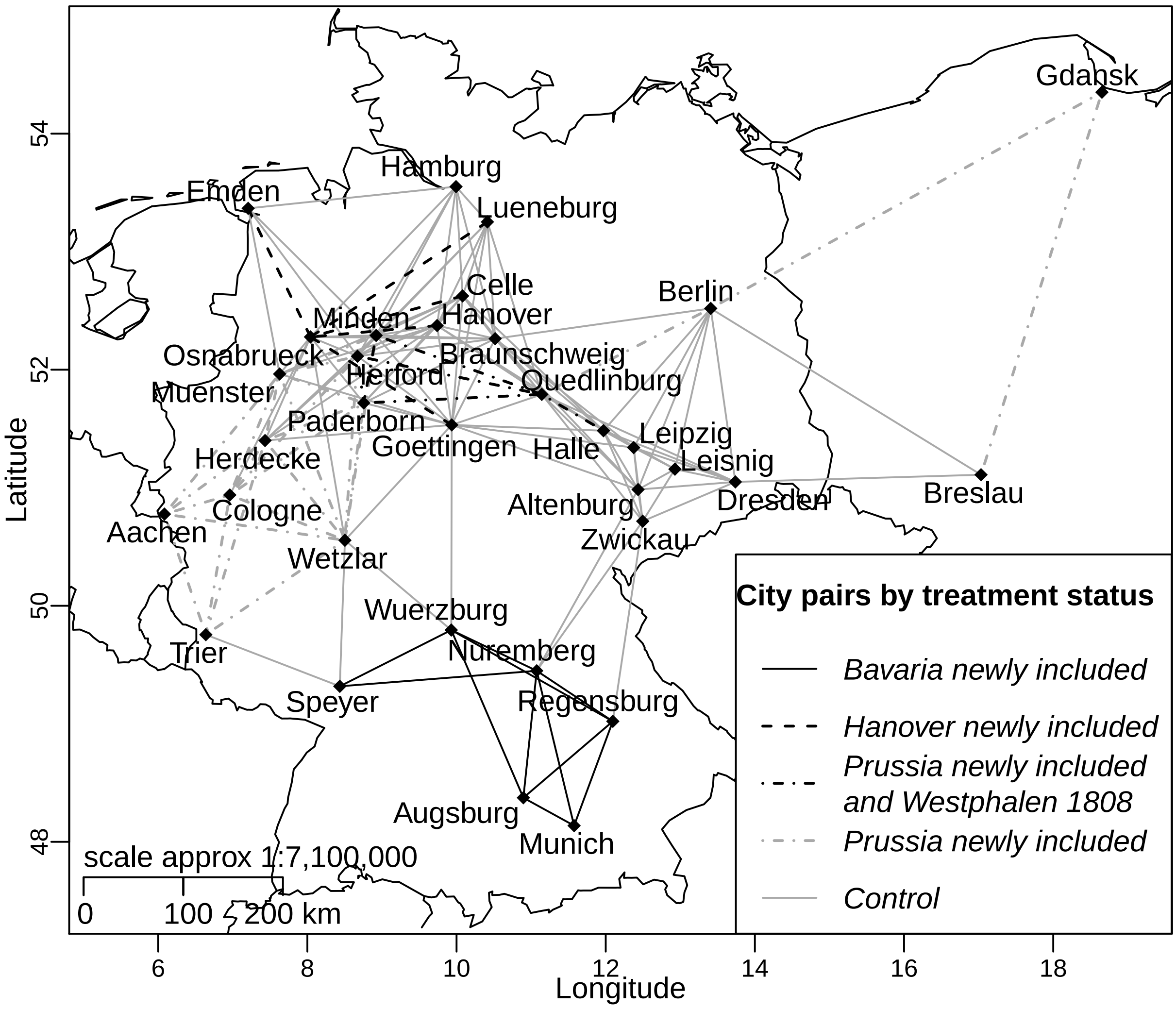

FIGURE 4.1: Towns and city-pairs in the data set.

Converting contemporary information on grain prices into a consistent data set faces several challenges. War-related dislocation and political change had temporarily negative effects on data quality. Specifically, wartime inflation partly took the form of currency debasement, which is incompletely reflected in the factors applied to convert historical currency units into silver equivalents (for details, see Albers et al. (2018), supplementary appendix SA2 and this study, SA4.1). Since we are unable to adjust for debasement in a uniform way across all towns, it is virtually impossible to study price gaps during the war period. The transformation of the political regime that occurred during the period under study also meant that the ways in which local and regional authorities observed rye prices underwent considerable change. In the majority of markets, old, local measures and currencies were replaced by state-wide standards, and reforms in regulation and information collection led to alterations with respect to the type of prices that were quoted. All this is a potential source of errors, and while we took every effort to handle the available information with care, data quality remains limited in many cases.

State growth that took place in the wake of the Revolutionary and Napoleonic wars is mirrored in our data by the fact that in 1815 30 towns – that is, 91 percent of all markets – belonged to one among four states, namely, Prussia (14), Bavaria (6), Hanover (6) and Saxony (4). By contrast, in 1791 18 towns — about 55 percent of all cities — were the only market of a specific territory represented in the sample. Accordingly, in our baseline specification with a distance restriction of 200 kilometres but including Gdansk and Breslau, 11 percent of all city-pairs involved markets from the same territory in 1791; by 1815 this proportion had risen to 44 percent.

To avoid misleading inferences about the effects of state growth and institutional change on grain markets, it is important to take account of three factors that affected regional demand conditions: geography, early industrialization, and conditions of third markets located outside Germany. Particularly north of the central mountain range (*Mittelgebirge*) that divides Germany in an east-westerly direction roughly at 50–51 degrees of latitude, these factors appear to have exerted a strong influence on regional grain price trajectories between the late eighteenth and the early nineteenth centuries. Specifically, while rye prices fell slightly in most parts of Germany between the 1780s and the 1820s, the towns that became part of the Prussian Rhine province in 1815 and a string of cities located farther east at the northern foot of the central mountain range experienced a significant rise in food prices ('String of cities Herdecke-Breslau' in Table 4.1).

Many of the towns in these two areas possessed a hinterland that had begun to industrialize from the turn of the nineteenth century (Saxony, several parts of the Rhineland, southern Westphalia). This and the concomitant growth of population raised food demand and, hence, local prices. The trajectory of rye prices in the adjacent lowland zone east of the Rhine and north of string of towns located on the foot of the central mountain range suggests that grain markets contributed little to satisfying food demand in emerging industrial regions. Between the 1780s to the 1820s eleven out of twelve markets in this group experienced a decline of rye prices; the average fell from 0.42 to 0.39 grams of silver ('Lowland' zone in Table 4.1).

TABLE 4.1: Average rye price, 1780s and 1820s, in individual markets and regions (grams of silver per litre)

	London wheat	Gdansk rye	Cologne rye	Hamburg wheat	Hamburg rye
1780–9	0.93	0.34	0.40	0.61	0.47
1820–9	1.19	0.28	0.42	0.55	0.44
	Rhine prov., Prussia	Lowland	String of cities Herdecke-Breslau	South (Bavaria 1815)	Arnhem
1780–9	0.43	0.42	0.36	0.34	0.49
1820–9	0.45	0.39	0.40	0.33	0.46

Sources: Own computation; data for London and Arnhem from Allen's Database (2001); series for Arnhem in Amsterdam file; data for Germany from Albers et al. (2018) and this study, SA4.1.
Notes: 'Rhine prov., Prussia' contains Aachen, Cologne, Trier and Wetzlar. 'String of cities Herdecke-Breslau' refers to cities located at the northern foot of the central mountain range (*Mittelgebirge*), namely, from west to east, Herdecke, Paderborn, Quedlinburg, Halle, Altenburg, all towns of Saxony (Dresden, Leipzig, Leisnig, Zwickau), and Breslau. 'Lowland' refers to the cities located east of the Rhine and north of the string of cities Herdecke-Breslau, that is, Braunschweig, Berlin, Celle, Emden, Gdansk, Hamburg, Hanover, Herford, Lüneburg, Minden, Münster and Osnabrück. 'Bavaria 1815' consists of Augsburg, Munich, Nuremberg, Regensburg, Speyer and Würzburg.

The doubling of the wheat price gap between London and Hamburg over the same period provides a potential clue for understanding this puzzle. German and Polish ports had emerged as important suppliers of grain for industrializing Britain during the last third of the eighteenth century, but British grain imports from Germany collapsed during the first quarter of the nineteenth century. After the end of the war period, domestic grain producers were shielded from import competition by the Corn Laws enacted in 1815 (Kutz, 1974, 276, 287–8; Sharp, 2010). For the period 1815–61 it has been shown that successive liberalization of the Corn Laws narrowed the price gap between English and Prussian wheat (O'Rourke and Williamson, 1999, 83); in parallel, British imports of German grain recovered from the late 1820s, and by 1840 Hamburg again handled exports from a large hinterland to overseas markets (Soetbeer, 1840, 103–5; see also Sharp and Weisdorf, 2013, 96). This opens the possibility that the grain price depression experienced by the lowland regions after 1817 was to a large extent caused by the temporary reduction of British demand in the wake of the more protectionist Corn Laws.

By contrast, the narrowing of the price gap between Arnhem and Cologne between the 1780s and the 1820s suggests that the opening of the Rhine trade during the Napoleonic era led to an integration of grain markets on the banks of the lower Rhine and the adjacent Dutch regions. In sum, our empirical strategy will have to take into account the possibility that regional industrialization and international trade may have influenced the trajectory of grain prices in markets located north of the central mountain range during the period under study.

4.4 Empirical strategy

We build on a well-established tradition of employing grain price gaps as a proxy for market integration (e.g., Federico, 2012; Chilosi et al., 2013; Keller and Shiue, 2014). With zero trade costs the Law of One Price holds, and arbitrage has the effect of equalizing product prices in two spatially separated markets. Conversely, the magnitude of the price gap between two places is an indicator of market fragmentation, and a decline of price gaps indicates market integration.

We build on the idea of Acemoglu et al. (2011) and consider the wars as a natural experiment; however, we employ a different empirical setup. We consider the price gap between two cities in a given year as depending on whether a city-pair became part of a new territory (our 'treatment') and a set of control variables. We implement this empirical strategy using a DD framework, that is, we also use a different identification strategy compared to Keller and Shiue (2014) who work with instrumental variables. Our baseline model is:

$$y_{ijt} = \sum_{l=1}^{L=4} \beta_l (g_{lij} p_t) + \sum_{k=1}^{K=4} \gamma_k (x_{kij} p_t) + \sum_{t=1}^{11} \delta_t d_t + \sum_{t=13}^{T=27} \delta_t d_t + a_{ij} + u_{ijt}. \quad (4.1)$$

The dependent variable y_{ijt} denotes the rye price gap between two cities i and j measured in percent, formally: $y_{ijt} = |\frac{\text{price}_{it}}{\text{price}_{jt}} - 1|$. The time index runs from $t = 1, ..., T$, with $T = 27$; this corresponds to the years 1780–91 and 1816–30. The dummy variables g_{lij} are the treatment group indicators. This variable is set to 1, if town pair ij became part of the same territory after 1815 but was not in the same territory in 1791; we refer to these city-pairs as 'newly included'. We differentiate between four territories, namely, Bavaria, Hanover, Prussia, and a part of the newly included territory in Prussia that also belonged to the former Kingdom of Westphalen during the wars. That is, we allow for heterogeneous effects in the territories newly included into Prussia. The dummy variable p_t denotes the post-treatment period 1816–1830. We expect the parameters β_l to be negative, signaling a decreasing price gap and an increase of market integration.

The variables x_{kij} are control variables capturing the third market and demand pressure effects discussed above. We employ the regional dummy variables 'Lowland' and 'String of cities Herdecke-Breslau' (see note to Table 4.1) and their interaction with time as control variables. Additionally, we construct a dummy variable that captures the movement of price gaps between these two groups (cf. Figure S4.1 in SA4.2). The fourth control dummy relates to those city-pairs that were part of the Kingdom of Westphalen but were not integrated into Prussia. While the Kingdom of Westphalen is, of course, a direct outcome of the war and thus could be regarded as a treatment group, this would be inconsistent with the treatment 'inclusion into a new territory that existed after the war' that we focus on. The reason is that the Kingdom disintegrated in 1813 and its cities became either part of the new territories of Prussia (Paderborn, Quedlinburg) and Hanover (Osnabrück) or returned to their previous territorial status (Halle, Herford and Minden were returned to Prussia, and Göttingen to Hanover; Braunschweig and its surroundings constituted an independent territory).

To account for unobserved time effects, we include a set of time dummy variables d_t, where we omit the one for 1791; that is, the year prior to the treatment is the base year. The estimator we use to estimate eq. (4.1) is the within-estimator, which eliminates the individual (city-pair) specific time constant error component a_{ij}.

Our sample of grain prices allows us to construct grain price gaps for 528 city-pairs. However, given the mostly regional character of grain trade before the age of railway construction one might worry that using all possible combinations includes many city-pairs in the sample that were irrelevant for trade given the large geographic area of the German lands. In addition, while one could argue from an econometric standpoint that also long-distance city-pairs add information, the latter also add heterogeneity that is difficult to capture with the control variables x_{kij} in eq. (4.1). For these two reasons we employ a distance restriction. Apart from Breslau and Gdansk, the two most remote cities in our data set, we retain only pairs of cities that are at most 200 kilometres apart from each other. We then add Breslau and Gdansk by connecting them both to their closest two neighbors. This yields $n = 148$ city-pairs (100 in the control group). Furthermore, with a restriction of 250 kilometers, we would pass the number of cross-sectional units for which critical values are implemented in the unit root test we apply ($n = 200$; Pesaran, 2007).

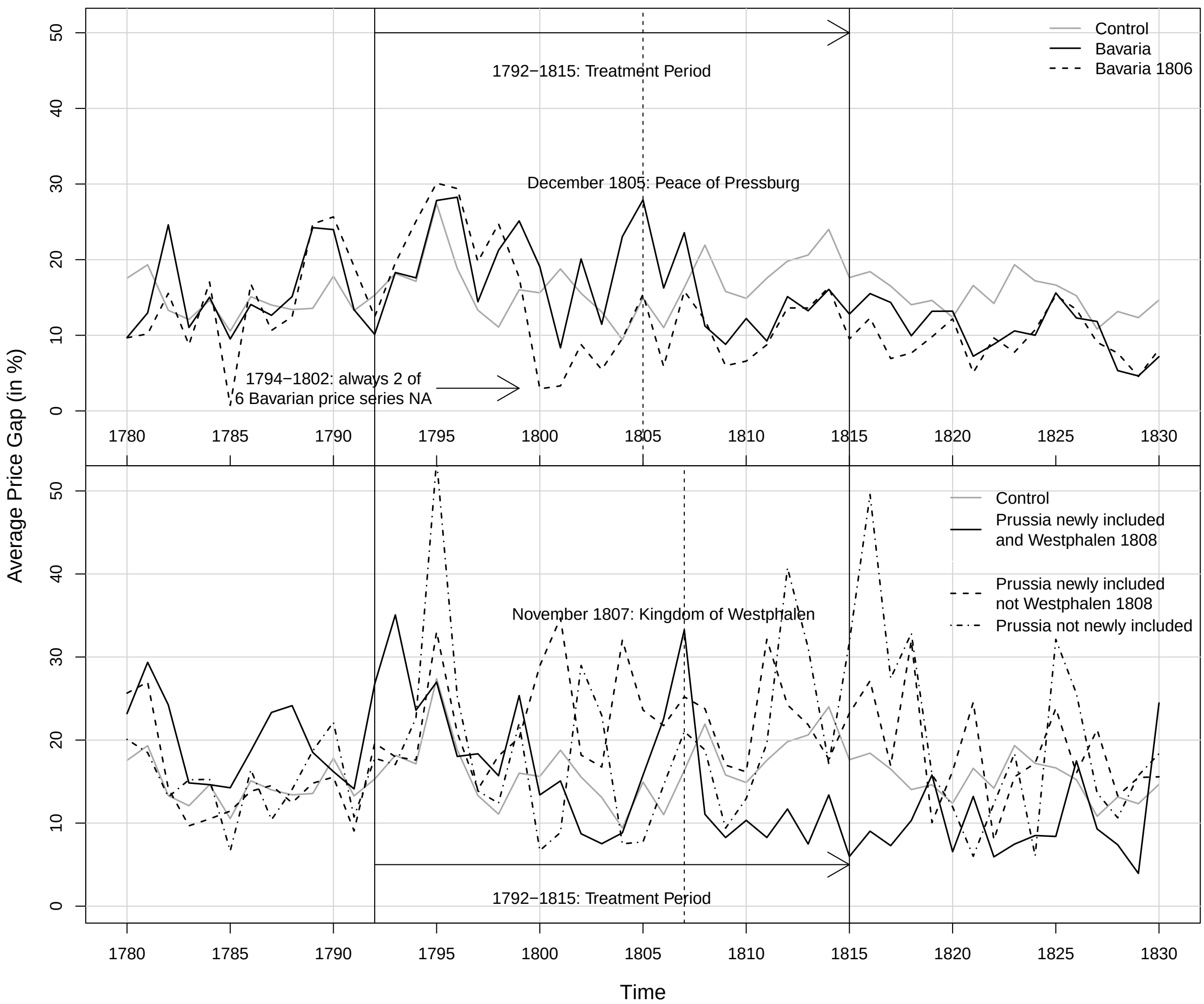

FIGURE 4.2: Average price gaps by treatment status, 1780–1830. In 'Bavaria' all city-pairs were newly included after the war; in 'Bavaria 1806' already a large fraction. 'Prussia newly included and Westphalen 1808' refers to city-pairs that were part of the Kingdom of Westphalen according to the borders prevailing in 1808 and were newly included into Prussia in 1815. 'Prussia newly included not Westphalen 1808': city-pairs where at least one city was not in Prussia in 1791 but both cities were in Prussia in 1815; does not include those city-pairs which were part of the Kingdom of Westphalen in 1808. 'Prussia not newly included' consists of all Prussian city-pairs in borders of 1815 that were also part of Prussia in 1791. NA: not available. Data sources: Albers et al. (2018, SA 1) and this study, SA4.1.

The panel data set we investigate is unbalanced; the time series dimension varies between $15 \leq t \leq 27$. Given the time series dimension, we performed the cross-sectionally augmented Im, Pesaran and Shin (CIPS) test. The test rejected the null hypothesis of a unit root with $p < 0.01$ in an interpolated sample and with $p = 0.011$ in a sample where series with missing observations are dropped. We summarize our empirical framework and the price gap data in Figure 4.2.

One concern with our empirical approach might be the fact that the Revolutionary and Napoleonic Wars and thus the treatment lasted a considerable time period. In this way the approach departs from a standard DD framework, where usually a particular date is assumed where a sudden change took place (Wooldridge, 2009, 453). We justify our approach with our focus on the net effect after the wars ended. The gap between the pre- and post-treatment periods could potentially affect our results, if another main intervention appeared during the War period, which reduced price gaps. We are unaware of such further changes. While there are additional reasons to exclude the war period (war-related destruction of markets, monetary disturbances due to war finance, and missing data), we also tested a specification that includes the war period (cf. Section 4.6). The main results remain robust.[1]

4.5 Effects of state growth on market integration

Table 4.2 shows parameter estimates for the effect of the inclusion of both elements of a city-pair into a new state on bilateral rye price gaps. The baseline estimate in column (1) includes the four treatment groups plus the control variables according to equation (1) in Section 4.4. The other specifications successively drop control variables and add further controls for states. Our main result is that the rye price gaps of those pairs where both cities were part of the territories of Bavaria in 1815 but not in 1791 were substantially lower after 1815 compared to the pre-war level. For Bavaria the estimated size of the effect is around 6.2 to 9.6 percentage points, which corresponds to 40-62 percent relative to the pre-war level, depending on the specification (models 2 and 1). For Prussia the effect is confined to those city-pairs that were also part of the ephemeral Kingdom of Westphalen and amounts to roughly 8 to 10 percentage points, equivalent to 41-51 percent relative to the pre-war level (models 1 and 2). City pairs that were part of the Kingdom of Westphalen but did not become integrated into Prussia also experienced a statistically significant reduction of the price gap. However, the effect is considerably smaller than for those pairs that were included into Prussia after the war. This suggests that the Kingdom of Westphalen had a trade creating effect and this effect was better preserved and/or potentially deepened by the integration into Prussia. By contrast, the city-pairs that were newly included into Prussia but were not part of the former Kingdom of Westphalen show no effect. Inclusion of both cities of

[1] Additionally, DD estimates could be driven by disintegration in the control group. However, the average price gaps in the control groups increases only slightly from 14.6 percent (1780–1791) to 15.1 percent (1816–1830). Both treatment groups show a substantial decline in the price gap. In the group 'Prussia newly included and Westphalen 1808' the price gap decreased from 19.6 to 10.3 percent, in Bavaria from 15.5 to 10.6 percent.

a pair into Hanover shows no stable effect, but this is also the state where only two cities were added to the territory.

TABLE 4.2: Negative effect of state growth on price gaps

	(1) Baseline	(2) as (1) w/o controls	(3) add states	(4) add controls
Bavaria x 1816-30	-0.0959^{***}	-0.0620^{***}	-0.0580^{***}	-0.0922^{***}
	(0.0201)	(0.0150)	(0.0153)	(0.0210)
Hanover newly included	-0.0105	0.0038	0.0078	-0.0068
x 1816-30	(0.0093)	(0.0103)	(0.0106)	(0.0099)
Prussia newly included and	-0.0798^{**}	-0.0995^{***}	-0.0955^{***}	-0.0736^{**}
Westphalen (1808) x 1816-30	(0.0315)	(0.0296)	(0.0298)	(0.0295)
Prussia newly included not	0.0081	0.0280	0.0318^{*}	0.0125
Westphalen (1808) x 1816-30	(0.0203)	(0.0172)	(0.0174)	(0.0208)
Westphalen (1808) not Prussia	-0.0366^{***}			-0.0357^{**}
newly included x 1816-30	(0.0135)			(0.0146)
Lowland x 1816-30	-0.0157			-0.0160
	(0.0151)			(0.0155)
String Herdecke-Breslau	-0.0473^{**}			-0.0407^{**}
x 1816-30	(0.0187)			(0.0201)
Between Lowland and String	-0.0570^{**}			-0.0642^{***}
Herdecke-Breslau x 1816-30	(0.0241)			(0.0242)
Hanover not newly included			0.0439^{***}	0.0174
x 1816-30			(0.0169)	(0.0194)
Prussia not newly included			0.1999	0.2267^{*}
x 1816-18			(0.1347)	(0.1354)
Prussia not newly included			0.0050	0.0322
x 1819-30			(0.0392)	(0.0410)
Saxony x 1816-30			-0.0207	-0.0146
			(0.0246)	(0.0262)
R^2	0.0658	0.0568	0.0686	0.0778
Adj. R^2	0.0184	0.0101	0.0213	0.0299
Num. obs.	3749	3749	3749	3749

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: city-pair price gap: |ratio of prices - 1|. Robust standard errors [Arellano] in (): heteroscedasticity consistent, clustered by group [=city pair] to account for serial correlation. The Pesaran-CD-test cannot reject the null hypothesis of no cross-sectional dependence.

How can we account for the heterogeneity of the effect of territorial expansion in Prussia? Most city-pairs newly integrated into this state that lay outside the Kingdom of Westphalen in 1808 are located in the west, more specifically in the lower Rhineland and adjacent Westphalia. This region, together with cities situated on the Rhine and Weser river systems, had experienced a process of grain market integration between the middle of the seventeenth and the late eighteenth century (Albers et al., 2018). The cities that belonged to the former Kingdom of Westphalen, by contrast, comprised inland cities that were not connected by navigable rivers. Consequently, price gaps in this region were higher than average in the pre-treatment period (SA4.3, Table S4.6, model 3). The territorial expansion of Prussia thus promoted market integration primarily in a region that had been left untouched by earlier processes of market creation. Grain markets in regions with a highly commercialized economy that had attained a considerable level of market integration at an earlier stage of development, by contrast, did not profit from the inclusion into a single state.

Comparison between specifications (1) and (2) in Table 4.2 shows that the inclusion of control variables does not alter the estimates of the effect sizes very much. The Lowland dummy variable captures the effect of decreasing British demand for German grain due to increased protection (a third market effect). The dummy 'String Herdecke-Breslau' focuses on the demand pressure associated with early industrialization in nearby inland zones from 1816–1830. As discussed in Section 4.4, we expect that the price gap between these two groups decreased over the war period due to these structural shifts in demand (in 'Lowland' prices decreased, in 'String Herdecke to Breslau' prices increased; cf. Table 4.1). The variable 'Between Lowland and String Herdecke-Breslau' captures exactly these two factors: price-gaps decrease by about 5.7 percentage points between pairs where one city is part of 'Lowland' and one city is part of 'String Herdecke-Breslau'.

Model (3) introduces dummy variables for the case in which both cities belonged to the same state in 1815, irrespective of their political status in the pre-war period. These dummy variables test whether belonging to a particular state both before and after the war had an effect on the trajectory of price gaps. If only territorial expansion mattered and institutional reforms within existing states were irrelevant, the city-pairs whose territorial status remained unchanged over the war period would not exhibit further integration. Models (3) and (4) confirm this for Hanover and Saxony;[2] in fact, model (3) suggests disintegration in Hanover, but the parameter is not robust to the inclusion of further controls in model (4). Bavaria does not appear in another form in model (3) because five of the six cities available in our sample were Bavarian in 1816 but were not part of Bavaria in 1791 so that no city-pairs on the old Bavarian territory exist in our sample.

For Prussia, model (3) includes an additional dummy variable that takes the value of 1 when both cities in a pair had been part of Prussia in 1791 and interacts it with the time dummies for 1816–8 and 1819–30. The former interaction effect captures the disintegration, which the old Prussian territory of 1791 saw during the Tambora crisis of 1816–1818.[3] The latter effect captures potential consequences of the implementation of the Prussian Customs Law, which took place only in 1818. From 1819 no further heterogeneity of price gaps between old and new Prussian territories appeared: The dummy variable for the territories of Prussia that were part of the latter in 1791 and 1815 interacted with the period dummy variable for 1819–30 remains insignificant and virtually zero (very similar for all of Prussia in borders of 1815, see Table S4.7). By contrast, the magnitude of the new territory effect found for the part that formerly belonged to the Kingdom of Westphalen, which was a forerunner of tariff reforms in the northern half of Germany, remains fairly stable and significant.

[2] All four markets located in Saxony from 1816 were already part of this polity before 1792.

[3] See also Figure 4.2 in Section 4.4. A sizeable amount of the heterogeneity in 1817 can be traced back to the city of Herdecke (see SA4.3, Table S4.3).

TABLE 4.3: Effects of French presence on market integration

	(1) FP like KS (2016) but dummy	(2) French territories	(3) as (2) add FP
Bavaria x 1816-30	-0.1230^{***}	-0.0925^{***}	-0.0854^{***}
	(0.0247)	(0.0207)	(0.0265)
Hanover newly included	-0.0090	-0.0250	-0.0237
x 1816-30	(0.0092)	(0.0179)	(0.0184)
Prussia newly included and	-0.0678^{**}	-0.0795^{**}	-0.0862^{**}
Westphalen (1808) x 1816-30	(0.0337)	(0.0315)	(0.0367)
Prussia newly included not	0.0100	0.0091	0.0068
Westphalen (1808) x 1816-30	(0.0198)	(0.0216)	(0.0223)
Westphalen (1808) not Prussia	-0.0322^{**}	-0.0346^{**}	-0.0378^{**}
newly included x 1816-30	(0.0138)	(0.0141)	(0.0168)
Lowland x 1816-30	-0.0136	-0.0141	-0.0163
	(0.0146)	(0.0149)	(0.0157)
String Herdecke-Breslau	-0.0606^{***}	-0.0441^{**}	-0.0378
x 1816-30	(0.0191)	(0.0192)	(0.0243)
Between Lowland and String	-0.0623^{**}	-0.0541^{**}	-0.0528^{**}
Herdecke-Breslau x 1816-30	(0.0244)	(0.0243)	(0.0255)
'FP 1 city' dummy x 1816-30	-0.0379^{**}		
	(0.0172)		
Hanover x 1816-30		0.0162	0.0136
		(0.0191)	(0.0201)
Left Bank of Rhine		0.0133	0.0114
x 1816-30		(0.0463)	(0.0459)
FP dummy x 1816-30			0.0109
			(0.0248)
R^2	0.0684	0.0660	0.0662
Adj. R^2	0.0208	0.0181	0.0180
Num. obs.	3749	3749	3749

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: city-pair price gap: |ratio of prices - 1|. Robust standard errors [Arellano] in (): heteroscedasticity consistent, clustered by group [=city pair] to account for serial correlation. The Pesaran-CD-test cannot reject the null hypothesis of no cross-sectional dependence. KS (2016): Keller and Shiue (2016); FP: French Presence.

Table 4.3 revisits the argument by Keller and Shiue (2016) that French rule impacted via institutional reform on market integration (see also Tables S4.8 and S4.9 in SA4.3). We use a variety of specifications of the French presence indicator. We distinguish a version which requires that only one city in a pair was French-ruled from a version where both cities were French-ruled. Both versions are considered in years or as dummy variable. Additionally, we test dummies for major regions that experienced specific forms of French presence. These are the left bank of the Rhine, which was an integral part of France for about one-and-a-half decades, and the ephemeral Kingdom of Westphalen (1807–13; eight markets located roughly between eastern Westphalia [borders of the Prussian province] and the Elbe river), a French satellite state. Except for the Kingdom of Westphalen we find no evidence that French presence contributed to market integration. In only one specification French Presence is statistically significant with a negative sign (Table 4.3, model 1) without altering the estimates of our baseline specification (column 1 of Table 4.2) in an important way. Once considered separately the parameter for the 'French Presence one city' dummy variable becomes zero and insignificant (Table S4.8 in SA4.3). Our findings imply that institutional reforms, such as the introduction of modern, inclusive commercial law in regions under French rule in the form of the *Code de commerce*, had no immediate consequences for the operation of trade, at least with respect to grain markets.

4.6 Robustness checks

The effects found for the territorial expansion of Bavaria and Prussia remain negative in a variety of robustness checks: tests of the common trend assumption, inclusion of additional control variables, and variations of the sample as well as the cluster applied in the computation of the standard errors. We also checked whether the four treatments in Table 4.2 remain significant and of similar size when considered separately. Tables for all robustness checks are shown in SA4.3.

First, we tested the common trend assumption by inclusion of treatment group specific time trends (Angrist and Pischke, 2009, 238–41). The main effect for Bavaria exhibits a small decrease in magnitude (from -9.6 to -8.8) while the one for Prussia gains slightly (from -8 to -9.7; specification 1 in Table S4.2). Time trends for both groups are not significant.

At this point we also considered a different empirical setup where we include the War period as discussed in Section 4.4 above and differentiate the several stages of the expansion of Bavaria and the foundation of the Kingdom of Westphalia in 1808 (model 2 in Table S4.2). Both effects appear significant and of similar size, if the beginnings of the treatments are set to 1806 for Bavaria (first round of territorial gains) and 1808 (when the Kingdom of Westphalen came into existence). Thus, our main results remain insensitive to changes in the specification of the treatment period.

We also tested a specification where we include treatment group specific time dummy variables in the pre-treatment period similar as in Angrist and Pischke (2009, 237). In model (3) (table S4.2, SA4.3) the parameter for Bavaria shows an reduction in absolute magnitude to -8 percentage points (model 1: -9.6) but remains statistically significant. The effect for Prussia instead drops from -8 percentage points in model 1 to -3.9 percentage points. The coefficient estimate is also not statistically different from zero; however, even this much reduced effect is still of economic significance as it corresponds to a decrease of the price gaps by 19.6 percent compared with the pre-War level. The loss in size in specification (3) stems from the additional time dummies that absorb peaks of the price gap that are not shared by the control group (see also Figure 4.2 in Section 4.4). This specification points at additional time varying heterogeneity in the price gaps that our baseline model or the inclusion of linear trends does not capture. In our view, shocks that widened price gaps in weakly integrated markets rather than anticipated border changes might be the cause of this result. For example, the specification with the Prussian effect considered separately (SA4.3, Table S4.6, model 3) suggests that the price gaps are on average higher (by 4.8 percentage points) than in the control group. Seen in this light and together with the insensitivity of the parameter to the inclusion of linear trends, the drop in the parameter's size is less concerning.

A second way to investigate the sensitivity of the results is to include additional control variables such as distance, a dummy variable accounting for the distance to places of early coal extraction and another one that accounts for regional heterogeneity during the crisis year 1817 (Table S4.3). We also tested for the inclusion of region varying time effects, where the regions are defined based on a climatological criterion to capture weather shocks to production. Additionally, we considered a pooled panel model with city dummy variables and distance instead of city-pair fixed effects. Results are similar at annual frequency as well as in a version with five-year averaged data.

Third, we tested whether the estimated parameters varied in an important way when dropping influential observations, using a balanced sample or dropping the remote connections involving Gdansk and Breslau (Table S4.4). The results for model (1) shown in Table 4.2 passed all these robustness checks.

Fourth, we analysed in how far our results depend on the cluster used in the computation of the standard errors (Keller and Shiue, 2014, 1194). Table S4.5 shows that the statistical significance of the effects is invariant to the cluster (control/treatment group level, city level, city-pair level). Table S4.5 also includes a specification where we employ the first-differenced (FD) estimator to eliminate the unobserved error component. In this specification, the standard errors increase very much. While the Prussian effect is still significant at the 10% level, the Bavarian effect is not significant at all. Increasing standard errors can be reconciled with the relative properties of the fixed effects (FE) and FD estimator as discussed in Wooldridge (2009, 487–8). If the data are weakly dependent (as the evidence from unit root test reported above shows) both estimators are unbiased but the efficiency of the FE estimator is higher if the untransformed errors are not serially correlated. This suggests that serial correlation is a larger problem in first-differenced errors.

Finally, we ran regressions applying the (pooled) ordinary least square estimator, where the treatment groups are considered one-by-one. In these regressions, the data set is restricted to the treatment group considered and the control group only and no additional dummy variables (neither for city-pairs nor for time) are included. Thus, the intercept in these specifications measures the average price gap of the control group, the treatment group indicator shows the latter's deviation from the control group's average price gap, and the interaction of treatment group indicator with the post-treatment time indicator quantifies the treatment effect. The effects for Bavaria and the part of Prussia that belonged to the former Kingdom of Westphalen remain similar in size (albeit closer to the results reported for model 2 without controls) and statistically significant when using the spatial correlation consistent (SCC) standard error that is robust to both cross-sectional and serial correlation (Driscoll and Kraay, 1998; Millo, 2017).

4.7 Conclusion

The main immediate effect of the Revolutionary and Napoleonic Wars on Germany was a consolidation of territorial power. From about 200 polities with limited sovereignty in the early 1790s, the number of political units shrank to about 40 in 1815. 60 percent of the German population were now citizens of one of two large sovereign states, Prussia and Bavaria. We explore the effect of the political transformation of Germany on market integration with an analysis of price gaps that draws on a novel data set of rye prices in 33 towns.

We find that price gaps between markets that belonged to different polities before the wars and that both became part of Bavaria decreased by about 62 percent between the 1780s and the 1820s. The effect in Prussia is smaller (reduction by 41 percent) and confined to an area that was part of a French satellite state, the Kingdom of Westphalen. State growth impacted on market integration in two ways: First, a plethora of heterogeneous internal tariffs was abolished and more trade partners belonged to the same territory, which deepened internal markets. Second, public authorities in large states were better capable to provide public goods. One example is the creation of a more efficient taxation system of external trade through uniform systems of external tariffs. Another example relates to programmes of state road construction after 1815, which reduced transport costs. Admittedly, our empirical setup does not allow differentiating between tariff reforms and road construction. But it is plausible that through these two channels — creation of internal markets and improved supply of public goods — the development of modern, much larger states benefitted market integration and deepened the division of labor between regions.

We do not find a uniform effect of French presence on price gaps in our sample. This suggests that institutional reforms such as the introduction the *Code de commerce*, had no immediate consequences for the operation of grain trade. At the same time, we find that price gaps were influenced by structural forces that impacted on regional price levels. Specifically, the beginnings of modern industrialization in inland zones were associated with upward pressure on food prices there, whereas the insulation of British grain markets in the wake of the Corn Laws increased price gaps between London and German port cities and depressed grain prices in the lowland of northern Germany. Although we find a strong effect of war-related state growth on market integration in two particular areas, its magnitude was insufficient to enable grain markets to accommodate for these structural shifts in foodstuff demand. Finally, the heterogeneity found for the large Prussian territory stems from the fact that territorial unification reduced price gaps primarily in an area that had not been affected by earlier processes of market integration (Albers et al., 2018). Regions with a highly commercialized economy having attained a considerable level of market integration at an earlier stage of development, by contrast, did not profit from the state growth that took place around 1800.

Whereas the short-term market creating effect of the development of modern states in the wake of the Revolutionary and Napoleonic Wars was thus limited, trade diversion effects that may have followed from the deepening of domestic markets in these states possibly created a motive for integrating markets at an inter-state level in the long run. In fact, already contemporaries deplored the trade-diverting effects of the implementation of unified systems of external tariffs at the beginning of the nineteenth century (Berding, 1980, 536). Recent research has stressed endogenous forces in the creation of the customs union (*Zollverein*) in 1834, that is, the desire to improve market access and reduce the costs of transit trade (Keller and Shiue, 2014; Huning and Wolf, 2016). State growth occurring in the wake of the Revolutionary and Napoleonic Wars led to a limited round of economic integration mostly at a regional level, which in turn laid the basis for subsequent integration rounds. Nevertheless, it took Germany another 100 years to evolve as an integrated economy in the 1930s (Wolf, 2009).

Chapter SA4

Supplementary Appendix: War, state growth and market integration in the German lands, 1780–1830

Authors: Hakon ALBERS* and Ulrich PFISTER†

SA4.1 Description of individual price series

In what follows we list each city in alphabetical order and describe data sources and conversion to grams of silver per litre. Series which are contained in the data set but not included in the list below are described in the SA of Ch. 3 (SA3.2).

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) Data from Prussian sources in 1814–1871 are converted as for Berlin in SA3.2. To account for the vellon inflation in 1808–22, we apply the specific adjustment factors for Königsberg from Jacobs and Richter (1935, 20) also to Gdansk (reason: geographic proximity). See the description of Königsberg in Albers et al. (2018, 50–1) for details on adjustment factors.

Rye (1501–1871)

Calendar year prices for rye until 1815 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 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 in one the sources based on this observation 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.

*Corresponding Author. Martin Luther University Halle-Wittenberg, Department of Economics, Universitätsring 3, D-06108 Halle, Germany

†University of Münster, Institute of Economic and Social History, Germany.

Data for 1814–65 from Königliches Statistisches Bureau (1867, 120–1) and for 1866–71 from Königlich Preussisches Statistisches Bureau (1865; 1867–72); all data in *Silbergroschen* per *Scheffel*.

Herdecke

Currency and volume conversion

Herdecke had Prussian currency and Berlin measures; see conversion factors applied to Berlin in Albers et al. (2018). Because Herdecke was not under Prussian rule in 1808–1815 no allowance is made for war-time debasement of currency (Prussian vellon inflation), however.

Rye (1780–1830)

Original prices for 1780–1816 as given in the local market ledger in *Reichstaler* and changing fractions of *Reichstaler* (*Gute Groschen*, *Stüber* or *Silbergroschen*) per *Berliner Scheffel* from Reinert (1920, 259–65). Reinert shows two prices per year, namely, for market days close to May 15 and November 11 (Saint Martin). We average both prices for each calendar year.

For November 13, 1780 the author gives a rye price of only 4 *Gute Groschen*. This has been corrected by adding a *Reichstaler* (24 *Groschen*).

Original prices from 1817 in *Reichstalern*, *Silbergroschen* and *Pfennig* per *Berliner Scheffel* from reports by the regional administration (*Amtsblätter*). We average both given prices for May and October to obtain calendar year prices.

Herford

Currency and volume conversion

Currency conversion applies *Graumannscher Fuß* (Gerhard and Kaufhold, 1990, 416). Herford was not under Prussian rule in 1808–1815 and thus, we do not correct for the Prussian vellon inflation. Volume conversion applies the rates for *Berliner Scheffel* (Witthöft, 1993, 26).

Rye (1771–1850)

Nominal prices are in *Reichstaler* and *Guten Groschen* (from 1822 in *Silbergroschen*) per *Berliner Scheffel* from Gerhard and Kaufhold (1990, 174–5).

1771–74: Calendar year prices based on monthly averages; 1776–1816: calendar year prices extrapolated from *Martini* prices applying the method developed by Albers et al. (2018, Supplementary Appendix 3.1.2).

1817–50: Calendar year prices mostly based on monthly averages of monthly minimum and maximum prices.

Leisnig

Currency and volume conversion
Conversion to gram silver applies the rate for converting *Pfennig* developed by Pfister (2017) for Leipzig. Conversion of volumes based on Witthöft (1993, 141).

Rye (1772–1830)
We calculate calendar year average prices using monthly data in *Taler* per *Dresdener Scheffel* from Uebele et al. (2013). The original source is a weekly newspaper (Uebele et al., 2013, 3). Data for 1825–1828 are missing.

Regensburg

Currency and volume conversion
Currency system and silver equivalent of *Kreuzer* from Kruse (1771, 310–311), Witthöft (1993, 22) and Gerhard (2002, 213). To convert local measures to litres, we apply 585.5 liter per *Regensburger Schaff* until August 1811 (Verdenhalven (1993, 49); for the same value see Chelius (1830, 323). From September 1811, we apply the rate given for *Bayerischer Scheffel* in Witthöft (1993, 76).

Rye (1786–1830)
1786–1814: Data are transcriptions from the manuscript source by Nürnberger (2016). Calendar year prices are calculated as mean of daily observations in *Gulden* and *Kreuzer* per local *Schaff* covering the entire calendar year with mostly one observation per month. For a few months missing observations were replaced with values from a very close day of a preceding or following month. E.g., the monthly value for April 1803 refers to March 30, 1803. The average for 1791 rests only on observations for January until May.

1792–1802: Data are missing.

1815–1830: Calendar year prices calculated as mean of monthly average prices in *Gulden* and *Kreuzer* per *Bayerischer Scheffel* from Seuffert (1857, 161–164).

Data from both sources are consistent. On average (1815–20, rye) the difference between the values given by Nürnberger and Seuffert, respectively, is ca. -0.28%.

Wetzlar

Currency and volume conversion
Conversion to grams of silver follows Gerhard and Kaufhold (1990, 57, 416); volume conversion as for Berlin in Albers et al. (2018).

Rye (1784–1859)
Calendar year prices 1784–1815 in *Reichstaler* per *Berliner Scheffel* from Kopsidis (1994, Appendix Table V.a/8). Data for 1816–59 in *Guten Groschen* (until 1821) or *Silbergroschen* (from 1822) per *Berliner Scheffel* are from Geheimes Staatsarchiv Berlin (b) except rye price for 1824 which is from Geheimes Staatsarchiv Berlin (a).

Zwickau

Currency and volume conversion
See Leisnig.

Rye (1764–1824)
The description of Leisnig applies but data for 1806 are not available.

SA4.2 City characteristics

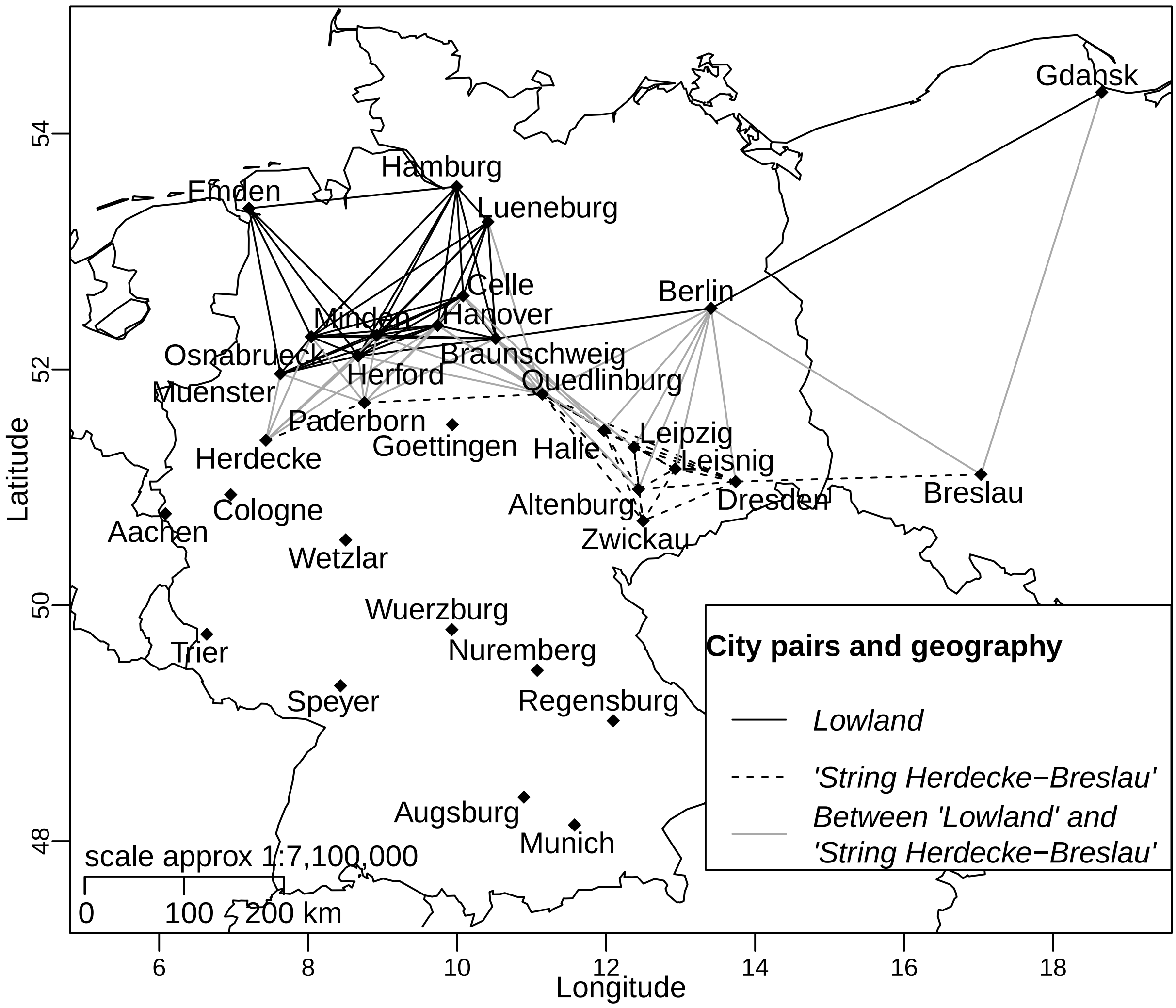

FIGURE S4.1: Geography and city-pairs. *Note*: see note to Table 4.1 in Chapter 4.3.

TABLE S4.1: City characteristics

Name	Latitude	Longitude	Lowland	Bavaria	Hanover	Prussia Prussia	Saxony	Westphalen	Annexed by France 1797	Polity_1791	Polity_1815	FP_ begin	FP_ end
Aachen	50.777	6.084	0	0	0	1	0	0	1	City_Aachen	Prussia	1797	1814
Altenburg	50.984	12.434	0	0	0	0	0	0	0	Gotha-Altenburg	Gotha-Altenburg		
Augsburg	48.373	10.896	0	1	0	0	0	0	0	City_Augsburg	Bavaria		
Berlin	52.519	13.408	1	0	0	1	0	0	0	Prussia	Prussia		
Braun-	52.263	10.518	1	0	0	0	0	1	0	Braun-	Braun-	1807	1813
schweig										schweig	schweig	1807	1813
Breslau	51.110	17.032	0	0	0	1	0	0	0	Prussia	Prussia		
Celle	52.625	10.081	1	0	1	0	0	0	0	Hanover	Hanover	1810	1813
Cologne	50.938	6.960	0	0	0	1	0	0	1	City_Cologne	Prussia	1797	1814
Dresden	51.049	13.738	0	0	0	0	1	0	0	Saxony	Saxony		
Emden	53.367	7.206	1	0	1	0	0	0	0	Prussia	Hanover	1807	1813
Gdansk	54.351	18.653	1	0	0	1	0	0	0	Poland	Prussia		
Goettingen	51.532	9.935	0	0	1	0	0	1	0	Hanover	Hanover	1807	1813
Halle	51.483	11.970	0	0	0	1	0	1	0	Prussia	Prussia	1807	1813
Hamburg	53.551	9.993	1	0	0	0	0	0	0	City_Hamburg	City_Hamburg	1811	1814
Hanover	52.374	9.739	1	0	1	0	0	0	0	Hanover	Hanover	1810	1813
Herdecke	51.400	7.432	0	0	0	1	0	0	0	Prussia	Prussia	1808	1813
Herford	52.114	8.671	1	0	0	1	0	1	0	Prussia	Prussia	1807	1813
Leipzig	51.340	12.375	0	0	0	0	1	0	0	Saxony	Saxony		
Leisnig	51.157	12.928	0	0	0	0	1	0	0	Saxony	Saxony		
Lueneburg	53.252	10.414	1	0	1	0	0	0	0	Hanover	Hanover	1811	1813
Minden	52.290	8.922	1	0	0	1	0	1	0	Prussia	Prussia	1807	1813
Muenster	51.964	7.628	1	0	0	1	0	0	0	Cleric_Muenster	Prussia	1808	1813
Munich	48.137	11.576	0	1	0	0	0	0	0	Bavaria	Bavaria		
Nuremberg	49.449	11.075	0	1	0	0	0	0	0	City_Nuremberg	Bavaria		
Osnabrueck	52.277	8.042	1	0	1	0	0	1	0	Cleric_Osnabrueck	Hanover	1807	1813
Paderborn	51.718	8.755	0	0	0	1	0	1	0	Cleric_Paderborn	Prussia	1807	1813
Quedlinburg	51.791	11.143	0	0	0	1	0	1	0	Cleric_Quedlinburg	Prussia	1807	1813
Regensburg	49.021	12.095	0	1	0	0	0	0	0	City_Regensburg	Bavaria		
Speyer	49.318	8.435	0	1	0	0	0	0	1	Cleric_Speyer	Bavaria	1797	1814
Trier	49.755	6.639	0	0	0	1	0	0	1	Cleric_Trier	Prussia	1797	1814
Wetzlar	50.555	8.503	0	0	0	1	0	0	0	City_Wetzlar	Prussia	1810	1813
Wuerzburg	49.794	9.929	0	1	0	0	0	0	0	Cleric_Wuerzburg	Bavaria		
Zwickau	50.716	12.496	0	0	0	0	1	0	0	Saxony	Saxony		

Note: Sources for all variables starting with Bavaria: Main source is Köbler (1999); the territory of the Kingdom of Westphalen at an early stage of its development is documented in Decret royal (1807); dating of French presence in Hamburg and Lüneburg follows Stubbe da Luz (2003, 78–81, 124–6). Detailed explanations: see next page.

Detailed Explanations of city characteristics in Table S4.1:

Lowland: See note to Table 1 in main text.

Bavaria, Hanover, Prussia, Saxony: Town was part of respective state in 1815.

Westphalen: Town was part of the Kingdom of *Westphalen*, a French satellite state after its creation late in 1807 (see Decret royal, 1807).

Annexed by France 1797: Left bank of the Rhine. See note on *FP_begin*, *FP_end* below.

Polity_1791, *Polity_1815*: Territory or state to which a town belonged in the respective year. In 1791, 'City_...' refers to an independent city state within the Holy Roman Empire, 'Cleric_...' to a clerical state, mostly a prince-bishopric.

FP_begin, *FP_end*: Year when French presence began and ended. Duration of French presence is calculated as the difference between these two years. Towns with no entries experienced no French presence. Years refer to the beginning or end of regular administration by French authorities or by those of a French satellite state (mostly the Kingdom of *Westphalen*, in two cases the Grand Duchies of Berg and of Frankfurt, respectively). Thus, military occupation is not considered as French presence. Following Fehrenbach (2008, 92), regular administration under French rule on the left bank of the Rhine began in 1797. If not stated otherwise in the sources mentioned, it is assumed that French presence ended in 1813 (e.g., Münster). Consultation of additional sources shows that this assumption is correct. The resulting years of French Presence are similar compared with Acemoglu et al. (2011). Note that Hamburg is not coded with French Presence in Keller and Shiue (2016, 35).

SA4.3 Additional result tables of robustness checks

TABLE S4.2: Test common trend assumption

	(1) as baseline + TG-specific trends	(2) as (1) with war period, split Bavarian territorial gains	(3) as baseline + TG-specific time dummy v. in pre-T-period
Bavaria x 1816-30	-0.0884^{**}	-0.0865^{*}	-0.0805^{**}
	(0.0378)	(0.0497)	(0.0382)
Hanover newly included	0.0200	-0.0387	0.0186
x 1816-30	(0.0315)	(0.0523)	(0.0251)
Prussia newly included and	-0.0973^{**}	-0.0500	-0.0385
Westphalen (1808) x 1816-30	(0.0470)	(0.0638)	(0.0305)
Prussia newly included not	0.0485	-0.0548	0.0467^{**}
Westphalen (1808) x 1816-30	(0.0310)	(0.0367)	(0.0229)
Westphalen (1808) not Prussia	-0.0367^{***}	-0.0101	-0.0365^{***}
newly included x 1816-30	(0.0135)	(0.0126)	(0.0135)
Lowland x 1816-30	-0.0153	-0.0035	-0.0145
	(0.0151)	(0.0131)	(0.0150)
String Herdecke-Breslau	-0.0459^{**}	0.0077	-0.0462^{**}
x 1816-30	(0.0187)	(0.0247)	(0.0186)
Between Lowland and String	-0.0564^{**}	-0.0091	-0.0555^{**}
Herdecke-Breslau x 1816-30	(0.0241)	(0.0205)	(0.0240)
Bavaria x Year	-0.0005	0.0005	
	(0.0022)	(0.0014)	
Hanover newly included x Year	-0.0023	0.0010	
	(0.0021)	(0.0016)	
Prussia newly included and	0.0012	-0.0013	
Westphalen (1808) x Year	(0.0025)	(0.0018)	
Prussia newly included not	-0.0031^{*}	0.0017	
Westphalen (1808) x Year	(0.0017)	(0.0011)	
Bavaria (1806) x 1806-15		-0.0869^{***}	
		(0.0301)	
Prussia newly included and		-0.1063^{**}	
Westphalen (1808) x 1808-15		(0.0439)	
R^2	0.0676	0.0750	0.0796
Adj. R^2	0.0191	0.0464	0.0208
Num. obs.	3749	7035	3749

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: city-pair price gap: |ratio of prices - 1|. Robust standard errors [Arellano] in (): heteroscedasticity consistent, clustered by group [=city pair] to account for serial correlation. The Pesaran-CD-test cannot reject the null hypothesis of no cross-sectional dependence in models (1) and (3) but rejects in model (2) with p=0.085 and thus the SEs are also made robust against cross-sectional correlation (Driscoll and Kraay spatial correlation consistent standard error).

TABLE S4.3: Other variables and distance specification

	(1) region-varying time effects effects	(2) distance to coal	(3) as model (3) in table 4.2 but city time effect for Herdecke	(4) add distance	(5) city FE pooling model with distance + time effects	(6) as (5) but with 5-year-averages
Bavaria x 1816-30	−0.0962***	−0.0893***	−0.0576***	−0.0944***	−0.0957***	−0.1014***
	(0.0303)	(0.0208)	(0.0152)	(0.0199)	(0.0199)	(0.0233)
Hanover newly included x 1816-30	−0.0160*	−0.0063	0.0082	−0.0122	−0.0114	−0.0042
	(0.0097)	(0.0100)	(0.0105)	(0.0105)	(0.0104)	(0.0115)
Prussia newly included and Westphalen (1808) x 1816-30	−0.0700**	−0.0757**	−0.0951***	−0.0786**	−0.0781**	−0.0773**
	(0.0335)	(0.0299)	(0.0297)	(0.0315)	(0.0316)	(0.0335)
Prussia newly included not Westphalen (1808) x 1816-30	−0.0011	−0.0081	0.0291*	0.0083	0.0100	0.0202
	(0.0202)	(0.0248)	(0.0171)	(0.0200)	(0.0201)	(0.0184)
Westphalen (1808) not Prussia newly included x 1816-30	−0.0370***	−0.0311**		−0.0343**	−0.0341**	−0.0291**
	(0.0131)	(0.0131)		(0.0140)	(0.0139)	(0.0133)
Lowland x 1816-30	−0.0293*	−0.0141		−0.0139	−0.0142	−0.0118
	(0.0166)	(0.0159)		(0.0151)	(0.0151)	(0.0150)
String Herdecke-Breslau x 1816-30	−0.0514*	−0.0420**		−0.0438**	−0.0456**	−0.0502**
	(0.0284)	(0.0190)		(0.0192)	(0.0192)	(0.0196)
Between Lowland and String Herdecke-Breslau x 1816-30	−0.0528**	−0.0560**		−0.0568**	−0.0567**	−0.0389*
	(0.0243)	(0.0237)		(0.0240)	(0.0240)	(0.0222)
Distance to coal		0.0328				
		(0.0257)				
Hanover not newly included x 1816-30			0.0443***			
			(0.0169)			
Prussia not newly included x 1816-18			0.1776			
			(0.1236)			
Prussia not newly included x 1819-30			0.0046			
			(0.0392)			
Saxony x 1816-30			−0.0197			
			(0.0245)			
Herdecke x 1817			0.1957***			
			(0.0513)			
log(distance) x 1816-30				0.0060	0.0043	0.0096
				(0.0105)	(0.0104)	(0.0109)
Intercept					−0.0912	−0.0327
					(0.0583)	(0.0592)
Bavaria					−0.0681*	−0.0834*
					(0.0378)	(0.0425)
Hanover newly included					−0.0162	−0.0171
					(0.0128)	(0.0117)
Prussia newly included and Westphalen (1808)					0.0167	0.0111
					(0.0312)	(0.0334)
Prussia newly included not Westphalen (1808)					0.0025	−0.0144
					(0.0206)	(0.0189)
Westphalen (1808) not Prussia newly included					0.0237*	0.0196
					(0.0142)	(0.0129)
log(distance)					0.0390***	0.0331***
					(0.0088)	(0.0088)
Lowland					0.0423*	0.0297
					(0.0233)	(0.0227)
String Herdecke-					0.0604*	0.0740**

Breslau					(0.0333)	(0.0323)
Between Lowland and String Herdecke-Breslau					0.1052***	0.0826***
					(0.0280)	(0.0264)
R^2	0.1443	0.0678	0.0759	0.0659	0.1839	0.3791
Adj. R^2	0.0881	0.0202	0.0288	0.0182	0.1670	0.3291
Num. obs.	3749	3749	3749	3749	3749	725

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: city-pair price gap: $\lceil$ratio of prices - 1$\rceil$.
Robust standard errors [Arellano] in (): heteroscedasticity consistent, clustered by group [=city pair] to account for serial correlation. Distance to coal: approximated by distance to town Witten, Ruhr area.

Additional explanations Table S4.3:

The region-varying time effects in specification (1) are based on a dummy variable capturing different climates. The dummy is coded 1, if continentality 1961–1990 $\leq$ 17°C; this corresponds roughly to North-Western Germany. The coding is based on Chapter 3.2 and Müller-Westermeier et al. (2001, map 7).

TABLE S4.4: Alternative specifications: other samples

	(1) drop influential obs.	(2) balanced by dropping series with NAs	(3) drop pair Herdecke-Muenster	(4) drop pair Osnabrueck-Hamburg	(5) drop all pairs Breslau or Gdansk
Bavaria x 1816-30	-0.0959^{***}	-0.0845^{***}	-0.0992^{***}	-0.0983^{***}	-0.0962^{***}
	(0.0201)	(0.0196)	(0.0200)	(0.0205)	(0.0201)
Hanover newly included	-0.0105	-0.0133	-0.0117	-0.0099	-0.0117
x 1816-30	(0.0093)	(0.0102)	(0.0092)	(0.0092)	(0.0093)
Prussia newly included and	-0.0798^{**}	-0.0902^{***}	-0.0773^{**}	-0.0772^{**}	-0.0800^{**}
Westphalen (1808) x 1816-30	(0.0315)	(0.0313)	(0.0310)	(0.0321)	(0.0315)
Prussia newly included not	0.0081	-0.0031	0.0001	0.0016	0.0075
Westphalen (1808) x 1816-30	(0.0203)	(0.0272)	(0.0195)	(0.0218)	(0.0203)
Westphalen (1808) not Prussia	-0.0366^{***}	-0.0232	-0.0373^{***}	-0.0368^{***}	-0.0373^{***}
newly included x 1816-30	(0.0135)	(0.0164)	(0.0134)	(0.0137)	(0.0135)
Lowland x 1816-30	-0.0157	0.0010	-0.0183	-0.0199	-0.0144
	(0.0151)	(0.0166)	(0.0150)	(0.0156)	(0.0151)
String Herdecke-Breslau	-0.0473^{**}	-0.0057	-0.0505^{***}	-0.0533^{***}	-0.0475^{**}
x 1816-30	(0.0187)	(0.0221)	(0.0187)	(0.0191)	(0.0187)
Between Lowland and String	-0.0570^{**}	-0.0423	-0.0643^{***}	-0.0621^{**}	-0.0571^{**}
Herdecke-Breslau x 1816-30	(0.0241)	(0.0296)	(0.0242)	(0.0260)	(0.0241)
R^2	0.0658	0.0792	0.0656	0.0665	0.0668
Adj. R^2	0.0184	0.0299	0.0181	0.0189	0.0193
Num. obs.	3749	2403	3722	3647	3722

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: city-pair price gap: $\lvert$ratio of prices - 1$\rvert$. Robust standard errors [Arellano] in (): heteroscedasticity consistent, clustered by group [=city pair] to account for serial correlation. The Pesaran-CD-test cannot reject the null hypothesis of no cross-sectional dependence.

Additional explanations Table S4.4:
Influential observations, which are dropped in specification (1), are detected using Cook's distance> 1 (e.g., Kleiber and Zeileis, 2008, 96–100).

The city-pair Herdecke-Muenster is dropped from the sample in specification (3), because both belonged to the Grand Dutchy of Berg (ephemeral French satellite state), where the code Napoleon was adopted (improved commercial law).

Some (parts of) paved roads were built during the time of French rule and of course also for military purposes (e.g., Pfeffer, 2012; see also Winkopp, 1806, 66). In particular, the city-pair Osnabrück-Hamburg is dropped from the sample in specification (4), because both were on a road constructed during the wars (Pfeffer, 2012; Walther, 2007).

TABLE S4.5: Baseline model: altering cluster level and first differences

	(1) by city-pair	(2) by city	(3) by CG/TG	(4) FD, by city-pair
Intercept	0.0875***	0.0875***	0.0875***	0.0108
	(0.0108)	(0.0139)	(0.0119)	(0.0071)
Bavaria x 1816-30	−0.0959***	−0.0959***	−0.0959***	−0.0861
	(0.0207)	(0.0172)	(0.0033)	(0.0621)
Hanover newly	−0.0105	−0.0105	−0.0105***	0.0581
included x 1816-30	(0.0095)	(0.0103)	(0.0021)	(0.0582)
Prussia newly included and	−0.0798**	−0.0798**	−0.0798***	−0.1190*
Westphalen (1808) x 1816-30	(0.0324)	(0.0381)	(0.0046)	(0.0707)
Prussia newly included not	0.0081	0.0081	0.0081***	0.0863
Westphalen (1808) x 1816-30	(0.0209)	(0.0144)	(0.0016)	(0.0686)
Westphalen (1808) not Prussia	−0.0366***	−0.0366***	−0.0366***	0.0040
newly included x 1816-30	(0.0139)	(0.0142)	(0.0029)	(0.0498)
Lowland x 1816-30	−0.0157	−0.0157	−0.0157***	−0.0859**
	(0.0156)	(0.0129)	(0.0050)	(0.0375)
String Herdecke-Breslau	−0.0473**	−0.0473***	−0.0473***	−0.1555**
x 1816-30	(0.0193)	(0.0140)	(0.0151)	(0.0685)
Between Lowland and String	−0.0570**	−0.0570	−0.0570***	0.0187
Herdecke-Breslau x 1816-30	(0.0248)	(0.0367)	(0.0141)	(0.0753)
R^2	0.2368	0.2368	0.2368	0.0642
Adj. R^2	0.1981	0.1981	0.1981	0.0553
Num. obs.	3749	3749	3749	3601

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: city-pair price gap: ∣ratio of prices - 1∣. Robust standard errors in (): heteroscedasticity consistent, clustered by group (cluster level varies, see model names) to account for serial correlation. Clustering by city exercised as two-way clustering by city 1 of the city-pair and by city 2 of the pair. The Pesaran-CD-test cannot reject the null hypothesis of no cross-sectional dependence. FD: first-differenced estimator.

TABLE S4.6: Treatment groups considered separately

	(1)	(2)	(3)	(4)
Intercept	0.1486^{***}	0.1486^{***}	0.1486^{***}	0.1486^{***}
	(0.0050)	(0.0050)	(0.0050)	(0.0050)
Bavaria	0.0106			
	(0.0167)			
Bavaria x 1816-30	-0.0533^{***}			
	(0.0200)			
Hanover newly included		-0.0514^{***}		
		(0.0164)		
Hanover newly included x 1816-30		0.0104		
		(0.0225)		
Prussia newly included and Westphalen (1808)			0.0477^{***}	
			(0.0163)	
Prussia newly included and Westphalen (1808) x 1816-30			-0.0929^{***}	
			(0.0197)	
Prussia newly included not Westphalen (1808)				-0.0074
				(0.0124)
Prussia newly included not Westphalen (1808) x 1816-30				0.0369^{*}
				(0.0192)
R^2	0.0068	0.0065	0.0081	0.0058
Adj. R^2	0.0061	0.0057	0.0074	0.0051
Num. obs.	2809	2679	2706	3187

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: city-pair price gap: |ratio of prices - 1|. Pooling model without city-pair fixed effects and without time effects. Driscoll and Kraay spatial correlation consistent standard error. Data are restricted to the considered group and the control group.

TABLE S4.7: Several groups considered separately

	(1)	(2)	(3)	(4)	(5)
Intercept	0.1525^{***}	0.1525^{***}	0.1470^{***}	0.1470^{***}	0.1496^{***}
	(0.0054)	(0.0054)	(0.0049)	(0.0049)	(0.0053)
Westphalen (1808) not Prussia newly included	-0.0158				
	(0.0138)				
Westphalen (1808) not Prussia newly included x 1816-30	-0.0160				
	(0.0161)				
Westphalen (1808)		-0.0003			
		(0.0135)			
Westphalen (1808) x 1816-30		-0.0361^{**}			
		(0.0153)			
Prussia			0.0057	0.0057	
			(0.0122)	(0.0122)	
Prussia x 1816-30			0.0157		
			(0.0199)		
Prussia x 1816-18				0.0806^{***}	
				(0.0201)	
Prussia x 1819-30				0.0011	
				(0.0170)	
Saxony					-0.0137
					(0.0090)
Saxony x 1816-30					-0.0162
					(0.0135)
R^2	0.0061	0.0091	0.0036	0.0126	0.0015
Adj. R^2	0.0054	0.0083	0.0030	0.0117	0.0007
Num. obs.	2571	2733	3349	3349	2544

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: city-pair price gap: |ratio of prices - 1|. Pooling model without city-pair fixed effects and without time effects. Driscoll and Kraay spatial correlation consistent standard error. Data are restricted to the considered group and the control group.

TABLE S4.8: Alternative specifications of French presence considered separately

	(1) FP two cities, with Hamburg	(2) FP like KS (2016)	(3) as (2) but dummy
Intercept	0.1699^{***}	0.1499^{***}	0.1548^{***}
	(0.0072)	(0.0066)	(0.0091)
FP dummy	-0.0362^{***}		
	(0.0100)		
FP dummy x 1816-30	0.0127		
	(0.0115)		
log('FP 1 city' years + 1)		-0.0009	
		(0.0011)	
log('FP 1 city' years + 1) x 1816-30		0.0028	
		(0.0059)	
'FP 1 city' dummy			-0.0094
			(0.0107)
'FP 1 city' dummy x 1816-30			0.0029
			(0.0107)
R^2	0.0115	0.0006	0.0005
Adj. R^2	0.0109	0.0000	-0.0001
Num. obs.	3403	3476	3476

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: city-pair price gap: |ratio of prices - 1|. Pooling model without city-pair fixed effects and without time effects. Driscoll and Kraay spatial correlation consistent standard error. Data are restricted to the considered group and the control group. KS (2016): Keller and Shiue (2016).

Overview of specification of French Presence (FP) in Tables 4.3, S4.8 and S4.9:
'FP 1 city' = FP like KS (2016): requires that one city of the pair was French-ruled. Counts number of years of French presence.
'FP 1 city dummy' = FP like KS (2016) but dummy: as 'FP 1 city' but sets positive number of years equal to 1.
'FP years' = FP in years: requires that both cities of the pair were French-ruled. Counts number of years of French presence.
'FP dummy': as 'FP years' but sets positive number of years equal to 1.
'FP two cities, with Hamburg': as 'FP dummy' but includes Hamburg. Hamburg was for some time ruled by the French but Keller and Shiue (2016, 35) explain why it should not be included in the FP variable.

TABLE S4.9: Additional specifications of French presence added to baseline model

	(1) add FP	(2) FP in years	(3) FP like KS (2016)
Bavaria x 1816-30	-0.0870^{***}	-0.0851^{***}	-0.1055^{***}
	(0.0266)	(0.0258)	(0.0241)
Hanover newly included	-0.0117	-0.0121	-0.0089
x 1816-30	(0.0095)	(0.0094)	(0.0097)
Prussia newly included and	-0.0875^{**}	-0.0897^{**}	-0.0728^{**}
Westphalen (1808) x 1816-30	(0.0363)	(0.0368)	(0.0346)
Prussia newly included not	0.0056	0.0030	0.0119
Westphalen (1808) x 1816-30	(0.0208)	(0.0223)	(0.0210)
Westphalen (1808) not Prussia	-0.0399^{**}	-0.0414^{**}	-0.0330^{**}
newly included x 1816-30	(0.0158)	(0.0166)	(0.0152)
Lowland x 1816-30	-0.0180	-0.0161	-0.0171
	(0.0157)	(0.0151)	(0.0150)
String Herdecke-Breslau	-0.0395	-0.0375	-0.0541^{***}
x 1816-30	(0.0244)	(0.0240)	(0.0203)
Between Lowland and String	-0.0550^{**}	-0.0532^{**}	-0.0604^{**}
Herdecke-Breslau x 1816-30	(0.0258)	(0.0266)	(0.0256)
FP dummy x 1816-30	0.0126		
	(0.0244)		
log(FP years + 1) x 1816-30		0.0076	
		(0.0122)	
log('FP 1 city' years + 1)			-0.0078
x 1816-30			(0.0104)
R^2	0.0661	0.0662	0.0662
Adj. R^2	0.0184	0.0185	0.0185
Num. obs.	3749	3749	3749

Note: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Dependent variable: city-pair price gap: |ratio of prices - 1|. Robust standard errors [Arellano] in (): heteroscedasticity consistent, clustered by group [=city pair] to account for serial correlation. The Pesaran-CD-test cannot reject the null hypothesis of no cross-sectional dependence. KS (2016): Keller and Shiue (2016); FP: French Presence.

Chapter 5

Conclusion

Economics studies, *inter alia*, how humans produce and trade consumption goods and how institutions guide these activities. But nature also impacts on economic activity through geography, weather and climate. This thesis contributes to these areas of research by analyzing the agricultural sector in Germany in three papers with different macroeconomic settings.

Paper 1 quantifies how weather impacts on wheat yield volatility conditional on input variations and agricultural policy in Germany in the modern growth regime where only few people work in agriculture. Paper 2 investigates if and where a market for grain existed in Malthusian Germany when the majority of Germans were employed in agriculture and eventually died in hunger crises. We show that a market for grain developed in North-Western Germany and along major rivers. Foodgrain prices became more stable, at least partly due to market integration. Paper 3 studies the effect of the Napoleonic and Revolutionary Wars on the grain market in Germany. We show that the induced state growth led to a limited round of integration in inland regions that had been left untouched during the earlier process of market integration. However, Germany was still far away from being an integrated economy in 1830.

Paper 1 shows how weather induces systemic fluctuations in agricultural wheat production in Germany. While the sub-periods we analyze in paper 1 witnessed an increase in volatility in aggregate German wheat yields from 4.1% (1996–2002) to 8.2% (2003–2009; calculated as CV, for data sources see Figure 5.1), we must be careful not to jump to quick conclusions: There is no clear trend towards increased volatility over time in aggregate wheat yields. The last seven years (2012–2018) exhibited the highest absolute variability in all consecutive seven-year-periods during 1963–2018 (SD: 0.62), however, only the second highest volatility (8%). The highest volatility was 11% in 1963–1969 (SD: 0.4).[1]

In the face of global warming, one potential concern might be that in the future large spatially correlated yield losses could possibly jeopardize food security, mostly in low-income countries (Gaupp et al., 2017). According to the IPCC report, "[...] climate change will increase crop yield variability in many regions (*medium evidence*, *medium agreement*) (emphasis in original, Porter et al.,

[1] Similarly, wheat yield volatility based on 14-year-periods is lower after 1990 than before (e.g., 11.5% in the years 1963–76; 7.1% in 2005–18). Note that Germany comprised two different economic systems after the second World War (capitalist Western Germany and socialist Eastern Germany). With regard to the different production systems, the data before 1990 are not directly comparable to the data since the reunification.

2014, 505)."[2] However, the predictions about future extreme events that are spatially correlated are even more difficult (Porter et al., 2014, 513). In paper 1, we argued that Gaupp et al. show that current global wheat production is independent so that trade is one important channel to sustain food security in case of large national level output losses.

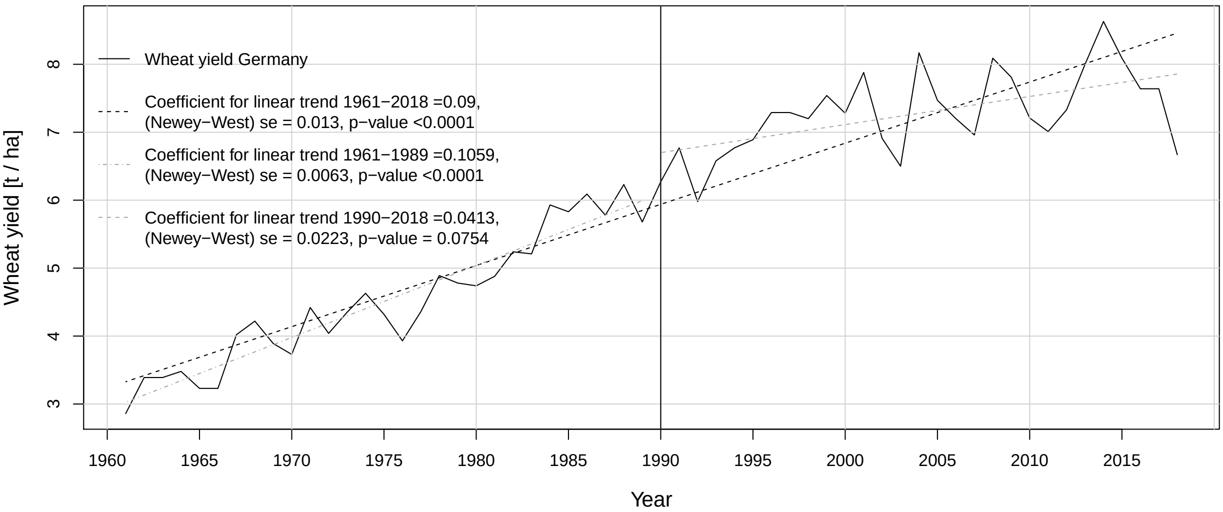

FIGURE 5.1: Wheat yields in Germany, 1961–2018. Vertical line highlights the year of the German reunification. Source: own representation. Data for 1961–2017 from FAOSTAT (2019); data for the year 2018 from BMEL (2018). According to FAOSTAT, the data before/after 1990 are calculated for the same geographical area.

However, the global trade system is a man-made institution and as such subject to human intervention. National trade policies might fall back to protectionism. An example is the Russian grain export ban in 2010, which followed a heatwave (Porter et al., 2014, 503; Zier et al., 2011, 9; see also Pies et al., 2013, 104–5 for further examples). A further issue for poor countries is that demand failure might potentially prevent trade from stabilizing food supply, because income losses (e.g., of farmers or agricultural wage laborers due to crop failure) do not allow net-consumers to demand food at the market (Ravallion, 1987b, here in the context of colonial India). Thus, while international trade provides an important security net, technological change, e.g., through improved seeds that are resistant to weather shocks (e.g., Emerick et al., 2016; Porter et al., 2014, 515), is a further important way to adapt to climate change and ensure global food security.

[2]In this context, yield variability is measured as volatility using the CV (Porter et al., 2014, 505–6). Note that the certainty in this statement about the future (as signaled by the wording in italics) takes a value in the middle of the used scale in both cases. It could take one the following values: low, medium, robust (for evidence) and low, medium, high (for agreement), see IPCC (2014, 37). The projections about yield levels also show a high degree of uncertainty. "[...] it is only *about as likely as not* that the net effect of climate and CO_2 changes on global yields will be negative by 2050, but *likely* that such changes will occur later in the 21st century (emphasis in original, Porter et al., 2014, 513)." The likelihoods associated with the used wording are 33–66% and 66–100%, respectively (IPCC, 2014, 37). A main problem for predictions is that increasing levels of CO_2 raise yields of certain crop types (e.g., wheat); however, negative effects of increasing levels of ozone (O_3) on yields counter this so-called fertilization effect of CO_2 (Porter et al., 2014, 499–500, 512). The relative strength of these opposing effects is unknown (ibid.).

Technological progress in agriculture is particularly relevant given the additional population pressure the world faces. Recently, a concern is that global population growth leads to increasing demand for food while land productivity *growth* in developed countries is declining (Madsen and Islam, 2016, 648–9) and global warming could make agricultural output potentially more risky as discussed above. Figure 5.1 revisits the observed trend of aggregate wheat yields in Germany, a particular form of land productivity. For the years after 1990, the trend estimate is smaller than before. Figure 5.1 shows absolute yields and corresponding trend parameters; the trend parameters based on logarithmised yields are 1.65% for the period 1961–2018, 2.38% for 1961–89, and 0.57% for 1990–2018. The latter estimates corroborate the observation by Madsen and Islam (2016).

Population growth and climate change have led to renewed interest in (agricultural) technology, agriculture, and global economic growth (Lanz et al., 2017; 2018). While these authors show that the Malthusian pressure that arises from the fixed factor land[3] appears manageable, it is clear that better technology will allow to use less land less intensively in agricultural production. This can help to dampen the negative external effects of agriculture on the environment such as the global loss in biodiversity (Foley et al., 2011, 338) or nitrate in surface waters and groundwater due to too much applied fertilizer (e.g., reported for Germany), which can lead to eutrophication (European Commission, 2016). It follows that a thorough understanding of the determinants of agricultural technology is vital to cope with the global challenge of providing a growing population with food in a changing climate.[4]

The recent price booms in agricultural commodities, e.g., in 2008 and 2010, motivated new research into cereal prices with at least two directions. One line of research aims at understanding whether the fundamentals in agricultural markets, that is supply and demand factors, are changing as discussed above or in how far financial speculation could be an explanatory factor (e.g., Algieri, 2014; Bruno et al., 2017; Conrad, 2015; Pies et al., 2013; Pies, 2015; Prehn et al., 2013; Will et al., 2013). An additional factor discussed in this literature are biofuel policies that have been adopted in many industrialized economies. A further question resulting from recent price increases in agricultural commodities is how many people were pushed into poverty by the surge in prices (e.g., Ivanic et al., 2012).[5] While the details are beyond the scope of this thesis, the work of these authors shows against the background of the investigations presented here that both today and in the past, food prices are of particular concern for societies and an important subject of study in economics.

[3]Next to Engel's law, land is the second important reason why agriculture matters for economic growth (Irz and Roe, 2005, 145, see also Jones and Vollrath, 2013, 202–5). I discuss the role of land for economic growth in more detail in Appendix D.

[4]Note that agriculture has always provided non-food goods (fibre for clothes). More recently, agricultural commodities have been used for bioful production. Related to the latter phenomenon, a relatively new term is the 'bioeconomy', which captures the idea that many products in an economy could potentially be produced based on biomass (Lewandowski et al., 2018, 14).

[5]Poverty can be measured in different ways. The simplest one would be the headcount ratio, a more elaborate way is the Foster-Greer-Thorbecke index; Taylor and Lybbert (2015, ch. 4) provide an overview on these different measures.

One key result of paper 2 is that a more stable foodgrain price led to improved food security at the aggregate level in pre-industrial Germany. Mortality declines in Germany in the 18th century (Pfister and Fertig, 2010, 31, figure 4; see also Figure S3.9 in SA3.6.5). Similar developments are reported for other European pre-industrial economies (Persson, 1999, 33–4 and references therein). For a negative effect of lower price volatility on mortality, two assumptions are necessary. First, if survival probability is a concave function of food consumption, more stable food *consumption* translates into higher survival probability and hence, lower mortality (Persson, 1999, 31–3 based on Ravallion, 1987a, ch. 2; see particularly pp. 24–31). However, arguing that a more stable *price* of food increases survival probability (lowers mortality) at a given income requires the elasticity of the slope of the survival function to be sufficiently large (Ravallion, 1987a, 35–6; see also Ravallion, 1997, 1214–5).[6] This leads to new empirical questions such as: Do the assumptions of Ravallion's model apply in the German case or more generally in pre-industrial Europe? While it is plausible that the grain price volatility moderation is relevant in explaining the downward trend of mortality, other explanations such as less epidemics are potentially important and thus, empirical tests could help to determine the quantitative contributions of these factors.

In the time period we study in paper 3, more precisely in the 1810s, Germany entered the post-Malthusian growth regime (Fertig et al., 2018; Pfister, 2017; Pfister and Fertig, 2010). While the beginning of the 19th century saw other important changes such as the intensified diffusion of the potato, paper 3 adds a new entry on the list of potential mechanisms that account for the transition to the post-Malthusian era in Germany. It is plausible that market integration, which was induced by state growth led to Smithian growth and thereby contributed to increased TFP at the aggregate level.

Smithian growth is the relevant mechanism for effects of market integration on aggregate output in both papers 2 and 3. In a standard Malthusian textbook model, higher TFP allows for a larger population size (Jones and Vollrath, 2013, ch. 8, see also Appendix B.1). Population size is related to the so-called scale effect in growth models that generates higher rates of technological change. The scale effect is the argument that more people have more new ideas, because the potential for each individual for having an idea is independent of population size (Kremer, 1993, 681–2).

In unified growth models, a scale effect on the rate of innovation working through population size à la Kremer is often crucial to explain the increase in the levels of technology and population at early stages of development (Galor, 2011, 147; Møller and Sharp, 2014, 118–9; Nunn and Qian, 2011, 607–8). It is notable that this scale effect is a crucial ingredient in models both with and without a role for human capital[7] investments (Galor and Weil, 2000; cf. Strulik and Weisdorf, 2008).[8]

[6]I study this point in more detail in Appendix E.

[7]"Human capital is the stock of skills and productive knowledge embodied in people (Rosen, 2018, 5992)."

[8]In contrast, in models that aim to study only the modern growth regime, several mechanisms have been proposed to avoid increasing rates of economic growth (which are counter-factual for the modern growth regime, see Jones, 1995b) since the seminal contribution of Romer (1990). The approaches include so-called semi-endogenous growth models (Jones, 1995a; Segerstrom, 1998) that are based on diminishing

One might conjecture that market integration is a simple yet realistic way to increase TFP and the scale of an economy at early stages of development. As the division of labor increases, the learning-by-doing mechanism that helps to increase the technology level like in unified growth models gains importance. Once societies are as sufficiently developed that they allow for property rights protection, the mechanisms of the endogenous growth literature (monopolistic competition and market oriented research and development; Romer, 1990) become relevant for technological change and economic growth.

This thesis has left unstudied roughly 160 years of economic development between the 1830s and the 1990s, that is, the Industrial Revolution that brought the transport revolution with the railway (e.g., Hornung, 2015), and further major technological changes, some relevant ones in agriculture. For example, Mokyr (2002) documents several important innovations that account for a change in agriculture specific useful knowledge: the identification of nitrogen and the law of the minimum (1830), the beginning of systematic experimentation (1840), and the understanding of nitrogen fixing bacteria (1880s) (Mokyr, 2002, 93–4; 19, 72; 93).[9] According to Mokyr (2008, 218), "[...] until 19th century organic chemistry widened the epistemic base, basic distinctions between nitrates, phosphorus and potassium were not made [...]." This new knowledge entailed better fertilizer application and improved yields (ibid.). The Haber-Bosch process (1909) provided a new input, that is, synthetically produced nitrogen fertilizer, to the agricultural production process (Grant, 2009, 192, 197; Mokyr, 2002, 109). This list can be extended with mechanical and biological innovations, which followed (e.g., the tractor and improved seeds, see Grant, 2009, 198 and Pardey et al., 2010, 959–65).

The focus of this thesis were effects of weather and inputs on cereal output volatility in an industrialized economy and the foodgrain market at early stages of economic development. The three essays provide answers on these areas of research but also point at new directions for further explorations in the role of agriculture for the economy, in particular with regard to food prices and agricultural technology.

returns to knowledge/technological opportunities, Schumpterian growth models with variety expansion (Aghion and Howitt, 1998, ch. 12) or Schumpeterian models with rent protection activities (e.g., Grieben and Şener, 2009). Ha and Howitt (2007) and Dinopoulos and Şener (2007) provide excellent overviews on the classification of endogenous growth models.

[9]Mokyr (2002) defines two different types of so-called useful knowledge, propositional knowledge (the epistemic base) and prescriptive knowledge (techniques) (Mokyr, 2002, ch. 1, particularly p. 4). Economic growth is understood as a result of broadening the propositional knowledge, which overcomes simple acceptance of the fact that some techniques work (Mokyr, 2002, 2, 26).

Appendix A

Engel's law and consumer preferences

This appendix provides background information on preferences that are used in macroeconomic models and in particular on those that are consistent with Engel's law. In A.1, I review these preferences; in A.2, I explain what the term 'non-homothetic' means, which is often used in conjuction with preferences that incorporate Engel's law.

A.1 Details on CRRA preferences and the Stone-Geary aggregate

Barro and Sala-i-Martin (2004, 87, 91) and Acemoglu (2009, 308–9, 698) discuss constant relative risk aversion (CRRA) preferences which are given by:

$$U = \int_0^\infty e^{-(\rho-n)t} u(c_t) dt, \tag{A.1}$$

with:

$$u_t = \frac{c_t^{1-\theta} - 1}{1-\theta}. \tag{A.2}$$

Overall utility U is derived from total per-capita consumption c_t, which generates utility at each instant of time t according to the utility function $u(c_t)$. Following Barro and Sala-i-Martin (2004, 87, 91), overall utility is the present value of the sum of all the utilities $u_t(c_t)$ of all household members at each t; $u(c_t)$ is also called 'felicity function'.

The parameter ρ is the time preference rate and n denotes population growth. If the coefficient of relative risk aversion $\theta = 1$, these preferences are logarithmic as in Matsuyama (1992, 321) (except for the role of the minimum food requirement).

These preferences are also called constant intertemporal elasticity of Substitution (CES; also abbreviated CIES) preferences (Acemoglu, 2009, 309; Barro and Sala-i-Martin, 2004, 91 see, e.g., Romer, 1990, S88 or Irz and Roe, 2005, 146). The reason is that $1/\theta = \sigma_u$, that is, the inverse of the risk aversion parameter equals the intertemporal elasticity of substitution between consumption today and tomorrow; and the risk aversion parameter θ again equals the elasticity of the marginal utility of consumption ε_u (Acemoglu, 2009, 295, 297, 309). The higher σ_u (the lower risk aversion θ or the elasticity of marginal utility ε_u), the more consumers are willing to give up consumption today in exchange for more consumption tomorrow.

The elasticity of the marginal utility of consumption is defined as the percentage change in the *slope* of the utility function (where the slope is the first derivative $u_t'(c_t)$) divided by the percentage change of consumption: $\varepsilon_u = -\frac{du_t'(c_t)}{u_t'(c_t)} / \frac{dc_t}{c_t}$. This term can be rewritten in a first step to: $-\frac{du_t'(c_t)}{d(c_t)} \cdot \frac{c_t}{u_t'(c_t)}$ and in a second step to: $-u_t''(c_t) \cdot \frac{c_t}{u_t'(c_t)}$, which equals eq. (8.23) in Acemoglu (2009, 295). Because utility is concave in consumption (increasing but at a decreasing rate), the second derivative of the utility function is negative. To achieve a positive definition of the elasticity, the minus sign is added.

The CRRA preferences above become non-homothetic, if aggregate per-capita consumption c_t is given by the Stone-Geary aggregate (Acemoglu, 2009, 628; Irz and Roe, 2005, 146; Matsuyama, 1992, 321):

$$c_t = (c_{a,t} - z_a)^{\varphi} c_{m,t}^{1-\varphi}. \tag{A.3}$$

Herein, $c_{a,t}$ and $c_{m,t}$ denote consumption of the agricultural and manufacturing good, respectively; φ and $1-\varphi$ are the corresponding shares in aggregate consumption. The minimum food requirement z_a makes the utility function non-homothetic, which we discuss below in A.2.

Engel's law has important effects on the transitional behavior of an economy towards a steady state depending on the model type (e.g., Kongsamut et al., 2001; Irz and Roe, 2005). Exercise 8.31 from Acemoglu (2009, 323–4) analyses how Stone-Geary preferences affect the neoclassical growth model; the corresponding solution is available in Peters and Simsek (2009, 76–81). The important result is that these preferences prevent the economy from the evolution on a balanced growth path (BGP); the latter can only be reached asymptotically, that is, at very high income levels (ibid.).[1] This is because the importance of the minimum food requirement in the capital accumulation equation declines as the level of technology (and hence income) increases (ibid.). Asymptotically, the capital accumulation equation converges to the one with standard preferences (ibid.). Furthermore, while Stone-Geary preferences allow for an income elasticity of food demand lower than one (in line with Engel's law), this elasticity is converging to one at higher consumption levels, which is counterfactual (Irz and Roe, 2005, 146, note 3).

A.2 Homogeneous and homothetic functions

To understand what 'non-homothetic' means, it is useful to review what the terms homogeneous and homothetic mean. A function $f(x_1, x_2)$ is homogeneous of degree k, if increasing x_1 and x_2 by factor t leads to $t^k \cdot f(x_1, x_2)$ (Simon and Blume, 1994, 484). E.g., in case of a production functions that is homogeneous of degree 1, multiplying both inputs x_1 and x_2 by factor $t = 2$

[1] The definition of a BGP goes back to the Kaldor facts (Acemoglu, 2009, 57, 65). "...balanced growth refers to an allocation where output grows at a constant rate and capital-output ratio, the interest rate, and factor shares remain constant (Acemoglu, 2009, 57)." In models with technological progress, the terms steady state and BGP can be used as synonyms (Acemoglu, 2009, 65). "We define a *steady state* as a situation in which the various quantities grow at constant (perhaps zero) rates (Barro and Sala-i-Martin, 2004, 33–4)." Kaldor described six empirical facts about macroeconomic variables. By now, there are 'new Kaldor facts' discussed by Jones and Romer (2010).

exactly doubles the output to $2^1 \cdot f(x_1, x_2)$ (Simon and Blume, 1994, 484–6). If this is the case, the production function exhibits constant returns to scale (CRS) (ibid.). Demand functions of consumers are functions of the goods' prices and income and homogeneous of degree 0 (Simon and Blume, 1994, 486). That is, scaling-up the prices of (in this case) goods x_1 and x_2 and income by the same factor does not affect which bundle of goods the consumer chooses (ibid.). Demand functions are derived from the maximization of the utility function subject to the budget constraint (consisting of available income and goods' prices), so that their properties stem from the utility function (Simon and Blume, 1994, 486, 489–90). Most relevant here is that the utility maximizing bundle of goods is where the marginal rate of substitution (MRS; the ratio of the marginal utilities derived from x_1 and x_2) equals the ratio of the goods' prices (ibid.). If income is scaled up while prices are unchanged, this utility maximizing point has the same MRS; the two utility maximizing points for different levels of income lie on a ray from the origin (ibid.). "...[D]oubling income doubles consumption of every good (Simon and Blume, 1994, 490)."

When studying utility, the exact numbers a utility function attributes the consumption of a particular good is irrelevant, that is, we are not interested in whether utility is exactly doubled (a cardinal property) when consumption is doubled but (just) that utility is larger than before (an ordinal property) (Simon and Blume, 1994, 496–8). In this regard, homogeneity is 'stricter' than homotheticity: "A function [...] is called homothetic if it is a monotone transformation of a homogeneous function [...] (Simon and Blume, 1994, 500)." That is, a homothetic function does not need to be homogeneous (ibid.). A monotonic transformation is loosely speaking the conversion of a function $u(x)$ to another function $g(u(x))$, which preserves the property that a higher x corresponds to higher $u(x)$ and higher $g(u(x))$ (see Simon and Blume, 1994, 497 for a precise mathematical definition and examples). The property of a homogeneous utility function that the utility maximizing points at different income levels have the same MRS is preserved in a homothetic utility function; this is an ordinal property (and a sufficient condition for a homothetic function, see Simon and Blume, 1994, 498, 500, 503). It directly follows that a homothetic utility function has an income elasticity of one (Simon and Blume, 1994, 500). That is, raising income by a constant factor increases consumption by the same factor (ibid.). A homogeneous utility function has the same property. Hence, a homothetic utility function $g(u(x))$ might assign different values of utility than the homogeneous utility function $u(x)$ but for both functions the utility optimizing goods bundle is the same and the income elasticity of demand is unity for varying levels of income.

From this follows that for a non-homothetic utility function, the utility maximizing goods bundle is *not* independent of the income level and the income elasticity of demand is *not* unity. This is the main characteristic of Engel's law as pointed out above: the income elasticity of food demand is lower than one.

Appendix B

Total factor productivity and technical change in pre-industrial Germany

This appendix first supports the conjecture that total factor productivity (TFP) likely increased in Germany in 1650–1790 (B.1). I then review potential sources of technical change in pre-industrial German agriculture (B.2).

B.1 Why did TFP likely increase in Germany 1650–1790?

Based on economic theory, one might conjecture that TFP likely increased 1650–1790. This period features a remarkable increase in the population size by about 40% from 13.5 million in 1618 to 19 million in 1800 (Pfister and Fertig, 2010). I refer to the population size for 1618, because the Thirty Years' War (1618–1648) led to a large loss in population that would lead to a overly optimistic picture of TFP-driven population growth after 1650. The reason is that in a Malthusian equilibrium population size tends back to its equilibrium value. For a population increase in a closed economy Malthusian equilibrium either aggregate factor input of land or TFP in agriculture must have increased (Andersen et al., 2016). Jones and Vollrath (2013, 192–195) show that in a Malthusian economy, a higher TFP results initially in higher per capita income; however, population growth is positively related to per capita income. Thus, population grows and due to decreasing returns in production, per-capita income declines back to subsistence consumption.

Empirically, the evidence from the German real wage (as a proxy for per-capita income) is mixed. Malthusian adjustment was relatively weak, because the adjustment to the stationary textbook equilibrium worked only slowly but at the same time unambiguous evidence of a unit root, that is, persistent changes in the German real wage emerges only if data for the first half of the 19th century are included (Pfister, 2017, 713, 716–8).

In addition, empirical evidence from the earliest available agricultural production function estimates (1830–1880) support the view that the explanatory power of increasing factor input is limited because of diminishing returns (Kopsidis and Hockmann, 2010).

B.2 Technical change in pre-industrial German agriculture

Here, I discuss four factors of technical change in pre-industrial German agriculture 1650–1790: science based innovations, the potato, legumes, and other reasons where our knowledge is limited because research has not yet investigated them intensively.

First, science-based innovations relevant to agriculture such as systematic breeding, mineral fertilizer etc. were mainly made later in the 19th century. Thus, these innovations can be ruled out as explanations for increasing TFP-levels in German agriculture in the years from 1650 to 1790 (Mokyr, 2002; 2009; Kopsidis and Bromley, 2016).

Second, the introduction of the potato is considered a major cause for population growth in the Old World for the period 1700–1900 (Nunn and Qian, 2011). The authors report that the potato was introduced in Germany at the beginning of the 17th century but admit that until the end of that century "... potatoes remained a botanical curiosity in Europe (Nunn and Qian, 2011, 602)." Their results indicate that the effect of potato suitability on population becomes larger in size after 1800 compared to 1750 (Nunn and Qian, 2011, 624–6). Consistent with their research, the cropping share of potatoes was only roughly 2% in Germany around 1800 (vs. 38% rye, 23% oats, 17% barley and 7% wheat; van Zanden, 1999, 368, table 16.6). Roughly 8% of consumed calories were provided by potatoes in Germany (Pfister, 2017, S2, p. 3; similarly Uebele and Grünebaum, 2014 for Saxony in 1792). The potato gains a large share in caloric consumption after 1800 (Pfister, 2017). In short, the potato's importance was clearly limited for the German population increase prior to 1800, although it was known and consumed.

Third, technical change in agriculture prior to 1800 might have been driven in an important way by improved crop rotations, where the key point was to make use of nitrogen-fixing legumes. However, Chorley (1981, 73, 82) estimates the role of nitrogen fixation for Germany around 1770 as minor; the beginning of regular usage of legumes starts around this date (cf. Pfister and Kopsidis, 2015, 14–5 on Saxony ca. 1800). The cropping share of peas and beans (important legumes), for example, was ca. 6% in Germany in 1800 (van Zanden, 1999, 368, table 16.6). Furthermore, Allen (2008, 183, 188–9) emphasizes that increasing nitrogen supply through legumes worked at a slow pace (here on England).

Fourth, other factors, that is, liming, drainage, improvements of agricultural machinery, seed selection (possibly facilitated by trade), improved pest control through improved crop rotations cannot be deemed minor without sufficient knowledge. These points might have played a role and their relative importance prior to 1800 is unknown (Chorley, 1981, 86; Allen, 2008, 201; Slicher van Bath, 1963a, 264). Wide-spread mechanization is regarded as a 19th century phenomenon in Western Europe, however (Slicher van Bath, 1963a, 303). Recently, Andersen et al. (2016) have established a positive effect of the heavy plow—one of the most important mechanical innovations—on urbanization in Denmark and Europe. However, this phenomenon is largely credited to the Middle Ages 1000–1300.

Appendix C

The moving average filter in the frequency domain

Kelly and Ó Gráda (2014a) argue that the Little Ice Age *in Europe* is not real but rather a statistical artifact from smoothing temperature data. These authors discuss two versions of the so-called Slutsky effect.

1. "[...] applying a filter to a white noise series will generate *regular* cycles corresponding to the peaks in the transfer function of the filter (emphasis in original, Kelly and Ó Gráda, 2014a, 1387)."
2. "[...] applying a moving average to a white noise series will generate the appearance of *irregular* oscillations, as the filter is distorted by runs of high or low observations (emphasis in original, Kelly and Ó Gráda, 2014a, 1388)."

In what follows, I refer to the first version and explain why Kelly and Ó Gráda state that the value for the relevant peak in the transfer function of a 25-year moving average (MA) is ca. 17.5 years. For this purpose, I refer to textbooks on time series analysis and mathematical handbooks to derive the equation for the transfer function of the moving average filter given by Kelly and Ó Gráda (2014a, 1387).[1]

A linear filter can be regarded as the sequence (German: *Folge*; Merz and Wüthrich, 2013, 268) of weights (a_u) which are applied to transform an input, a time series y_t, to an output, the filtered time series z_t (Schlittgen, 2012, 46):

$$z_t = \sum_{u=-r}^{s} a_u y_{t-u}. \qquad \text{(Schlittgen 2012, eq. 2.36)}$$

Herein t denotes the period of the time series, e.g., years, and u indexes the periods which are included when the filter transforms the initial time series into the filtered time series. The periods u runs from $-r$ to s: $u = -r, ..., s$. The total number of periods included in the filter to calculate one filtered observation is given by $m = r + s + 1$. For example, $m = 25$ means that 25 years are used to calculate the filtered value z_t according to the equation above. In this case

[1] The two textbooks on time series analysis, which I use predominantly in this part of the appendix, Schlittgen (2012, 134)/Schlittgen and Streitberg (1994, 166), pursue on a different way and provide/derive the gain function for the particular example of a three-period MA filter but not the squared gain function (the terminology is explained in detail below). Thus their representation differs. The closest representation is Sargent (1994, 257) but I add several intermediate steps, generalize to m periods, and adopt a notation that is consistent with the basic definitions taken from Schlittgen (2012) and Schlittgen and Streitberg (1994). Neusser (2011, 104) does not provide a derivation for the MA filter.

$u = -12, -11, ..., -1, 0, 1, ..., 12$. Thus, s denotes the weight for the last lag of y_t in the above equation. E.g., $s = 12$ denotes the weight a_{12} for the lag y_{t-12}. In case of a simple moving average, the m weights a_u are all the same $a_u = 1/(r+s+1) = 1/m$ (Schlittgen and Streitberg, 1994, 36). This is the case which we consider here.

Time series analysis distinguishes the time domain and the frequency domain. The definition of the MA filter given above applies to the time domain. 'Frequency domain' means that a time series can be understood as harmonic waves, that is, overlapping cosine and sine functions of different frequencies (Schlittgen and Streitberg, 1994, 51). The frequency is $\lambda = 1/P$, where P is the period (ibid.). The period is the number of time units of a cylclical/periodic behaviour of a time series (Schlittgen and Streitberg, 1994, 50–1).

Schlittgen and Streitberg (1994, 54) provide an introduction to the frequency domain by explaining the periodigram (or empirical spectrum; German: *Periodigramm* or *Stichprobenspektrum*): *"Das Periodigramm ist eine Funktion $I(\lambda)$ der Frequenz und gibt für jede Frequenz λ an, mit welcher 'Stärke' (oder 'Intensität') harmonische Wellen dieser Frequenz in der Ausgangsreihe 'auftauchen'."*[2] The formal definition involves the covariance of a considered time series with the cosine function and the covariance with the sine function (Schlittgen and Streitberg, 1994, 59). The authors then use the Euler equation to illustrate that the cosine function corresponds to the real part of a complex expression and the sine function corresponds to its imaginary part (Schlittgen and Streitberg, 1994, 68). To see this, we first review complex numbers.

A complex number is defined as $z = a + i \cdot b$ with the real part a and the imaginary part b; i is the imaginary unit (Schlittgen and Streitberg, 1994, 484–5). The trigonometric form of any complex number can be stated as: $z = r \cdot (\cos\varphi + i \cdot \sin\varphi)$, where $r = |z|$ and φ denotes the radian of the corresponding angle on the unit circle (Schlittgen and Streitberg, 1994, 486–7, particularly figure B.2.1). The exponential form is: $z = r \cdot \mathrm{e}^{i\varphi}$ (Schlittgen and Streitberg, 1994, 488). Euler's equation allows to switch from the exponential form of a complex number to its trigonometric form: $\mathrm{e}^{i\varphi} = \cos\varphi + i \cdot \sin\varphi$ (Schlittgen and Streitberg, 1994, 488).

Applying these relationships, the periodigram can be rewritten as a complex exponential (Schlittgen and Streitberg, 1994, 68). The authors then first prove that the periodigram includes the autocovariance function (Schlittgen and Streitberg, 1994, 76–7). Second, they show that the periodigram, written as a complex exponential, is the Fourier transform of the autocovariance function of the corresponding time series (ibid.). The Fourier transform allows to switch from the representation in the time domain to the frequency domain. For a sequence (x_t) the Fourier transform is defined as: $F(\lambda) = \sum_{-\infty}^{\infty} x_t \mathrm{e}^{i2\pi\lambda t}$ (Schlittgen and Streitberg, 1994, 69–70).

The frequency domain is particularly useful for the analysis of filters (Schlittgen and Streitberg, 1994, 164). Compared to the time domain, it easier to analyze in the frequency domain how a filter alters the behavior of the input, that is, the initial time series (ibid.). For example, a filter might attenuate certain frequencies (Schlittgen and Streitberg, 1994, 165). This will be clarified further below.

[2]Translation: The periodigram is a function $I(\lambda)$ of the frequency and for each frequency λ it provides the 'power' (or 'intensity') with which harmonic waves of that frequency 'appear' in the initial series.

The transfer function of a filter is the Fourier transform $F_a(\lambda)$ of the sequence (a_u), the linear filter (Schlittgen, 2012, 132):

$$F_a(\lambda) = \sum_u a_u \mathrm{e}^{i2\pi\lambda u}. \qquad \text{(Schlittgen 2012, eq. 6.16)}$$

The absolute value of the transfer function is called the gain function of the filter: $G_a(\lambda) = |F_a(\lambda)|$ (Schlittgen, 2012, 132). The right-hand-side of the equation is the exponential form of a complex number.

Furthermore, $a_u \sum_u \mathrm{e}^{i2\pi\lambda u}$ is a geometric series (German: *Reihe*; Merz and Wüthrich, 2013, 298). A geometric series $\sum_{k=0}^{n} q^k$ can be rewritten as: $\sum_{k=0}^{n} q^k = \frac{q^0 - q^{n+1}}{1-q}$ (Merz and Wüthrich, 2013, 301). Application to the adapted eq. (6.16), setting $q = \mathrm{e}^{i2\pi\lambda}$, and using $a_u = 1/m$ for the weights (see above) yields:[3]

$$F_a(\lambda) = \frac{1}{m} \frac{\mathrm{e}^{i2\pi\lambda(-r)} - \mathrm{e}^{i2\pi\lambda(s+1)}}{1 - \mathrm{e}^{i2\pi\lambda}}. \tag{C.1}$$

Kelly and Ó Gráda (2014a, 1387) report the square of the absolute value of the transfer function: $|F_a(\lambda)|^2$, which is derived as follows. The absolute value of a complex number equals the absolute value of its complex conjugate: $|z| = |\bar{z}| = \sqrt{a^2 + b^2}$ (Bosch, 1991, 62). The complex conjugate of the complex number $z = a + b \cdot i$ is $\bar{z} = a - b \cdot i$ (ibid.). The exponential form of a complex number is $z = r \cdot \mathrm{e}^{i\varphi}$ (Schlittgen and Streitberg, 1994, 488) and its complex conjugate is: $\bar{z} = r \cdot \mathrm{e}^{-i\varphi}$. Furthermore, $\overline{z_1 \pm z_2} = \overline{z_1} \pm \overline{z_2}$ and $\overline{z_1 / z_2} = \overline{z_1} / \overline{z_2}$ (Schlittgen and Streitberg, 1994, 486). Thus, one can write: $|F_a(\lambda)|^2 = |F_a(\lambda)| \cdot |\overline{F_a(\lambda)}|$ (Schlittgen and Streitberg, 1994, 164). Applying these relationships to eq. (C.1) yields:

$$\begin{aligned} |F_a(\lambda)| \cdot |\overline{F_a(\lambda)}| &= \left|\frac{1}{m} \frac{\mathrm{e}^{i2\pi\lambda(-r)} - \mathrm{e}^{i2\pi\lambda(s+1)}}{1 - \mathrm{e}^{i2\pi\lambda}}\right| \cdot \left|\overline{\frac{1}{m} \frac{\mathrm{e}^{i2\pi\lambda(-r)} - \mathrm{e}^{i2\pi\lambda(s+1)}}{1 - \mathrm{e}^{i2\pi\lambda}}}\right| \\ F_a(\lambda) \cdot \overline{F_a(\lambda)} &= \frac{1}{m^2} \frac{\mathrm{e}^{i2\pi\lambda(-r)} - \mathrm{e}^{i2\pi\lambda(s+1)}}{1 - \mathrm{e}^{i2\pi\lambda}} \cdot \frac{\mathrm{e}^{-i2\pi\lambda(-r)} - \mathrm{e}^{-i2\pi\lambda(s+1)}}{1 - \mathrm{e}^{-i2\pi\lambda}}. \end{aligned} \tag{C.2}$$

This expression is similar to the one given by Sargent (1994, 257), who discusses the special case of a five-year MA filter. The expression can be simplified as follows:

$$\begin{aligned} F_a(\lambda) \cdot \overline{F_a(\lambda)} &= \frac{1}{m^2} \frac{\mathrm{e}^{i2\pi\lambda(-r+r)} - \mathrm{e}^{i2\pi\lambda(s+1+r)} - \mathrm{e}^{i2\pi\lambda(-r-s-1)} + \mathrm{e}^{i2\pi\lambda(s+1-s-1)}}{1 - \mathrm{e}^{i2\pi\lambda} - \mathrm{e}^{-i2\pi\lambda} + \mathrm{e}^{i2\pi\lambda(1-1)}} \\ &= \frac{1}{m^2} \frac{2 - \mathrm{e}^{i2\pi\lambda m} - \mathrm{e}^{i2\pi\lambda(-m)}}{2 - \mathrm{e}^{i2\pi\lambda} - \mathrm{e}^{-i2\pi\lambda}} \\ &= \frac{1}{m^2} \frac{2 - \left(\mathrm{e}^{i2\pi\lambda m} + \mathrm{e}^{-i2\pi\lambda m}\right)}{2 - \left(\mathrm{e}^{i2\pi\lambda} + \mathrm{e}^{-i2\pi\lambda}\right)}. \end{aligned} \tag{C.3}$$

Euler's equation allows to switch from the exponential form of a complex number to its trigonometric form: $\mathrm{e}^{i\varphi} = \cos\varphi + i \cdot \sin\varphi$ (Schlittgen and Streitberg, 1994, 488). Thus, $\mathrm{e}^{i\varphi} + \mathrm{e}^{-i\varphi} = \cos\varphi + i \cdot \sin\varphi + \cos\varphi - i \cdot \sin\varphi = 2\cos\varphi$. Applied to this case, we have:

[3] It is possible to have negative indices in a series (Merz and Wüthrich, 2013, 298). We can apply the formula to a geometric series with a complex exponential function (Schlittgen and Streitberg, 1994, 489).

$$F_a(\lambda) \cdot \overline{F_a(\lambda)} = \frac{1}{m^2} \frac{2 - 2\cos 2\pi\lambda m}{2 - 2\cos 2\pi\lambda}$$
$$= \frac{1}{m^2} \frac{1 - \cos 2\pi\lambda m}{1 - \cos 2\pi\lambda}. \quad \text{(C.4)}$$

Once we use $\omega = 2\pi\lambda$, that is, the angular frequency (German: *Kreisfrequenz*, Schlittgen and Streitberg, 1994, 482) in eq. (C.4), we obtain the expression given by Kelly and Ó Gráda (2014a, 1387) (cf. Sargent, 1994, 257 for the special case of a five-year MA filter).

Why does this expression simplify the analysis of the MA filter? We started with the MA filter in the time domain. The initial time series y_t, the input to the filter has a spectrum $f_y(\lambda)$. The theoretical spectrum of a time series is the Fourier transform of the theoretical autocovariance function (Schlittgen, 2012, 11, 129; Schlittgen and Streitberg, 1994, 156). The periodigram, which was introduced above, is the empirical counterpart of the theoretical spectrum, that is, the empirical spectrum of the empircal autocovariance function (ibid.). Multiplication of the spectrum of y_t with eq. (C.4) yields the spectrum $f_z(\lambda)$ of the output of the filter, that is, the spectrum of the filtered time series z_t: $f_z(\lambda) = |F_a(\lambda)|^2 \cdot f_y(\lambda)$ (Schlittgen, 2012, 133; Schlittgen and Streitberg, 1994, 164). E.g., a value of $|F_a(\lambda)|^2 < 1$ for a given frequency λ means that the filter attenuates that particular frequency in the spectrum of the filter's output (Schlittgen and Streitberg, 1994, 165).

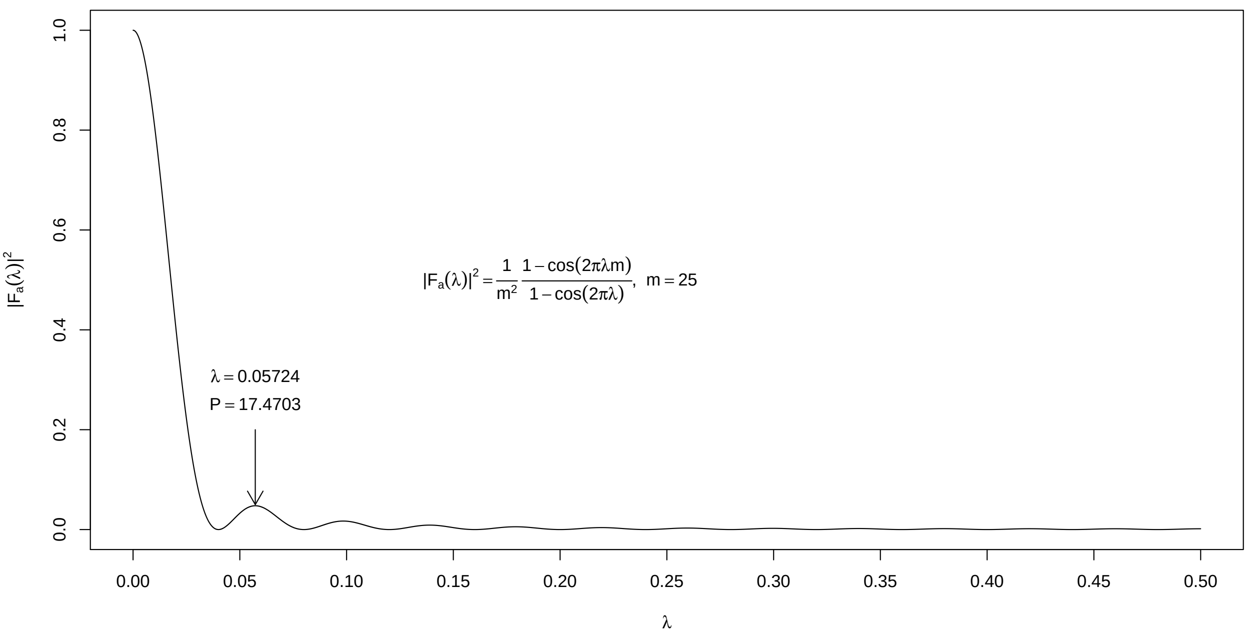

FIGURE C.1: Squared absolute value of the transfer function (=squared gain function) of the moving average filter with m=25 periods. P: Period, λ: frequency. Source: own calculation.

To illustrate, we draw the function $|F_a(\lambda)|^2$ (squared absolute value of the transfer function = squared gain function) of an m-period MA filter in Figure C.1 for $m = 25$, e.g., 25 years.[4] Very low values of λ pass the filter almost unchanged. This is the reason why the MA is called a low-pass filter (Schlittgen, 2012, 134). But even moderately higher frequences are attenuated very much: the curve tends to zero very fast. Additionally, there is a peak in this function which occurs at the frequency $\lambda = 0.05724$ which corresponds to the period $P = 1/\lambda = 17.4703$. The latter value represents "[...] around 17.5 years [...]" which Kelly and Ó Gráda (2014a, 1387) judge as "too short [...] to generate Little Ice Age behavior." In other words, a 25-year MA filtered time series displays cyclical behavior with a period of approximately 17.5 years, simply because the filter allows the corresponding frequencies to pass the filter while attenuating both lower and higher frequencies.

[4]Schlittgen (2012, 134) provides a plot of the gain function (not the squared gain function) of a three-period MA filter. Note that we only plot frequencies λ in the interval $[0, 0.5]$, because the cosine function is periodic with $P = 2\pi$ and $\cos(-x) = \cos(x)$ so that we also also know the behavior of the function for $-0.5 \leq \lambda \leq 0$ and hence, for a full interval of $1 \cdot 2\pi$ (Schlittgen and Streitberg, 1994, 61–2, 480; here on periodigrams).

Appendix D

The production factor land and economic growth

Under the otherwise unchanged assumption of constant returns to scale in the production function, land leads to diminishing returns to the other production factors (capital and labor). To counter the increasing food demand that results from population growth and puts pressure on the fixed production factor land, the rate of technological progress must be sufficiently high; otherwise a steady state (for a definition see note 1 in Appendix A.) with sustained growth does not exist (Irz and Roe, 2005, 149). This result does not only arise in the model by Irz and Roe (2005) but also in the Solow-model, which I discuss briefly below.

In the two-sector-model (agriculture and manufacturing) of Irz and Roe, either technological progress in agriculture and/or manufacturing must be fast enough. The latter point, namely, that technological progress *only* in manufacturing can potentially counter population growth is somewhat surprising. Irz and Roe (2005, 147) assume (like Kongsamut et al., 2001) that only the manufacturing good can be saved and invested into an increase of the economy's capital stock. That is, the consumer faces the decision between consumption of the manufacturing good or investment in the capital stock. Capital can be used in either the manufacturing or the agricultural sector to produce higher levels of manufacturing and/or agricultural output, respectively. Hence, technological progress in manufacturing allows to reallocate capital to agriculture while keeping manufacturing output at the same level. Via this capital-accumulation-channel, there is a positive 'spillover' of technological progress in manufacturing to the agricultural sector (Irz and Roe, 2005, 149).

The model by Irz and Roe (2005) is a neoclassical growth model with exogenous technological progress but an endogenous savings rate resulting from consumer optimization. This model type is called Ramsey-model or 'Ramsey-Cass-Koopmans-model' (to give credit to all authors who developed it; Ramsey worked in the 1920s, Cass and Koopmans in the 1960s, see Barro and Sala-i-Martin, 2004, 16, 18).

A simpler framework, that is, without consumer optimization, would be the Solow-model. Exercise 2.11. from Acemoglu (2009, 72) deals with the effect of land in the Solow-model without technological progress. With land *and* population growth, the model does not feature a steady state anymore (Peters and Simsek, 2009, 1–5).

Appendix E

Grain price volatility and survival probability (Ravallion 1987)

Ravallion (1987a, ch. 2) explains how lower grain price volatility leads to higher survival chances. The necessary condition he derives is that the elasticity of the slope of the survival function is larger than two (Ravallion, 1987a, 35–6). His result it given in eqs. (2.9) and (2.10) (Ravallion, 1987a, 36), which are verified below.

According to Ravallion, survival $s(.)$ is an implicit function of consumption x: $s(x), x = \frac{m}{p}$ (time indices dropped). Consumption is determined by exogenous money income m and the price of food p.

Ravallion gives the second derivative of the survival function with regard to the price in a way, where he factored out γ, that is, the elasticity of the slope of the survival function, which is given in his eq. (2.10). We now plug Ravallion's (2.10) into (2.9) and simplify:

$$\begin{aligned}\frac{\partial^2 s(\frac{m}{p})}{\partial p} &= x \cdot s'(x)\left\{2 - \overbrace{\left[(-x) \cdot \frac{s''(x)}{s'(x)}\right]}^{\gamma}\right\} \cdot p^{-2} \\ &= \left[2x \cdot s'(x) - (-x^2) \cdot s'(x) \cdot \frac{s''(x)}{s'(x)}\right] \cdot p^{-2} \\ &= \left[2xs'(x) + x^2 s''(x)\right] \cdot p^{-2} \\ &= \left[2s'(x) + xs''(x)\right] \cdot x \cdot p^{-2}.\end{aligned} \tag{E.1}$$

To verify (E.1), we form the second derivative of the survival function w.r.t. the price from scratch. Using the chain rule, the first derivative is:

$$\frac{\partial s(\frac{m}{p})}{\partial p} = \overbrace{s'(\frac{m}{p}) \cdot (-1)}^{u} \cdot \overbrace{\frac{m}{p^2}}^{v}. \tag{E.2}$$

Applying the product rule, the second derivative is:

$$
\begin{aligned}
\frac{\partial^2 s(\frac{m}{p})}{\partial p} &= \overbrace{(-1)s''(\frac{m}{p})\cdot(-1)\cdot\frac{m}{p^2}}^{u'}\cdot\overbrace{\frac{m}{p^2}}^{v}+\overbrace{s'(\frac{m}{p})\cdot(-1)}^{u}\cdot\overbrace{\frac{m}{p^3}\cdot(-2)}^{v'} \\
&= s''(\frac{m}{p})\cdot\frac{m^2}{p^4}+2\cdot s'(\frac{m}{p})\cdot\frac{m}{p^3} \\
&= s''(\frac{m}{p})\cdot\frac{m}{p}\cdot\frac{m}{p^3}+2\cdot s'(x)\cdot\frac{m}{p^3} \\
&= [s''(x)\cdot x+2s'(x)]\cdot\frac{m}{p^3} \\
&= [s''(x)\cdot x+2s'(x)]\cdot x\cdot p^{-2}.
\end{aligned}
\tag{E.3}
$$

Eqs. (E.3) and (E.1) are equivalent.

Why is γ the elasticity of the slope of the survival function? Simon and Blume (1994, 304–5) define the concept of elasticity. Adapted to the considered case, the *consumption* elasticity of the slope of the survival function is:

$$
\begin{aligned}
\gamma &\equiv -\frac{\text{\% change in slope of survival function}}{\text{\% change of consumption}} \\
&= -\frac{\frac{\Delta s'(x)}{s'(x)}}{\frac{\Delta x}{x}} \approx -\frac{\frac{\partial s'(x)}{s'(x)}}{\frac{\partial x}{x}} = -\frac{\frac{\partial s'(x)}{\partial x}}{\frac{s'(x)}{x}} = -x\cdot\frac{s''(x)}{s'(x)},
\end{aligned}
\tag{E.4}
$$

where the last term is the one given by Ravallion (1987a, 36). The elasticity γ is negative, because the slope of the survival function becomes smaller as consumption increases.

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