# What can price volatility tell us about market related institutions? Conditional heteroscedasticity in historical commodity price series.[1]

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**Abstract**

Researchers have always been aware of the working of markets. Especially the volatility of markets is seen as an indicator of market efficiency in the broadest sense. Yet, the way in which volatility is estimated makes it difficult to compare the price volatility across regions or over time for two empirical reasons. First, in case of non-stationary price series or even if the series are trend stationary, the variance is inflated by the trend. Second, the variance of the log prices contains information on a lot of factors that are not related to market efficiency, but rather to region or time specific factors. Hence both the standard Coefficient of Variance CV) and the detrended version will bias the estimates of market efficiency.

In order to remove the effect of the trend and of the time- and region-specific effects we apply a conditional heteroscedasticity model to several European and Asian datasets which are chosen to get as widely divergent datasets as possible. We find that although the CV results in all cases in a higher level of volatility, the amount of overestimation depends on the presence of factors like inflation, the demand structure, and agricultural structure which, although influencing price volatility, are not directly related to the risk management strategies or efficient working of markets. Using this approach we find that on average in Europe price volatility had declined (and market efficiency increased) following the early 16$^{th}$ century.

**Keywords:** market efficiency, coefficient of variance, historical prices, Europe, Asia.

**JEL codes:** N14 (economic history Europe), N15 (Economic History Asia), G14 (market efficiency)

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## 1 Introduction

Researchers have always been aware of classifying the working of markets. Ancient or Early Modern markets were often classified as "conspicuously conventional, irrational and status-ridden" when an author wanted to stress the economic backwardness of a society or as "a sphere of activity with its own profit-maximizing, want-satisfying logic and rationality" when the opposite point of view was defended by Andreau (2002, 15).

Since quantitative ways to measure market efficiency are in short supply, this discussion has been allowed to survive until relatively recent times. From an economic-theoretical (i.e Temin 2002) perspective, studies have shown that even in ancient times markets showed a remarkable degree of efficiency, a fact that is also widely acknowledged among scholars of medieval economies (i.e. Britnell and Campbell 1995). This view, however, is in contrast with the evidence from those studies making use of price volatility of (largely) agricultural products. The demand for such products is generally price inelastic and, hence, deviations in the price are largely caused by supply shocks. Since market integration leads to arbitrage between different markets which consequently smoothes price fluctuations, price volatility is thought to be indicative for well functioning markets (i.e. Persson 1999). These studies often find an increase in market efficiency from the late medieval period onwards.

Indeed, this is exactly what we find if we look at Table 1. With the exception of the 16th century,

Table 1

there is a clear downward trend in the coefficient of variation from 385 BC until 1900. Hence, according to this measure, the further back in time one goes, the less efficient markets are.

How can the evidence based on historical analyses, which indicates a relatively efficient market already in the medieval period, be squared with the statistical behavior of prices which indicates a declining volatility over time? It has been widely acknowledged that the volatility of prices is also caused by other factors than purely the market mechanism (i.e. Persson 1999: 107-8). A clear effect can be found in the estimates for the 16th century in Table 1. We can see that price volatility in the sixteenth century, a century that is commonly described as "the price revolution" (Munro 2003), seems extraordinary high. In fact, according to above table, price volatility, measured by Coefficient of Variation (CV), in England between 1500 and 1650 AD was as high as in Babylonia between 250 and 150 BC. Yet, very few people would argue that market efficiency actually declined during that period.[2]

There are also several other factors that, although they do play a role in broadly defined market efficiency, are profoundly country (or time) specific. For example, the agricultural structure may have a strong influence on seasonality and interannual volatility. A clear example can be found in the China where for example Anhui province has a single rice harvest per year while Guangxi and Guangdong have two rice harvests a year and Hainan even has three. Consequently, even if one harvest failed, prices would still be lowered at the end of the year when the second (or third) harvests become available. This is completely different from, for example, Java, that not only has one rice harvest a year, but also rice made up roughly 50 per cent of total value added in agriculture during the twentieth century (Van der Eng 1996, Table A.1.2). Although the presence of multiple harvests clearly reduces price fluctuations, it does not say much about the efficiency of the institutions, that we are really interested in.

Because of these problems in calculating market efficiency from price series, in this paper we argue that the standard volatility measures tell us little about actual development of market related institutions over time and are not comparable across countries with different

[2] An exception is Bateman (2007) who argues that market efficiency follows a U-shaped pattern with a through during the sixteenth century.

product structures, consumer preferences and weather conditions.[3] Hence, one needs to have a careful look on the existing volatility measures, and find one which can be used to proxy institutional efficiency and is comparable across regions (or at least have less incomparable factors). Building on many related studies (i.e. Persson 1999; Söderberg 2004), we argue that the standard deviation of the log of the detrended prices series, or preferably that of regression residuals from some structural model of prices (or in the absence of necessary data an ARMA model) can both be used for such purpose, even though the latter is preferable. It is important to note that this paper focuses on the measurement of volatility within a single region over time (a time-series perspective) and not on measuring the spatial variability of prices (a cross-sectional perspective). In the next section we briefly explain the problem with traditional measures and argue that both the trend and volatility of these measures are problematical. This is elaborated upon in Sections 3 and 4 where we discuss the effect of the trend and the volatility respectively. Section 5 discusses an alternative measure of dispersion that is comparable over time and across countries. Section 6 applies these theories to empirical models while Section 7 concludes the paper.

## 2 Problems with the standard measure of volatility (CV)

The Coefficient of Variation is often used as a measure of market efficiency (Persson 1999). The problem, however, is that the CV, as a measure of price volatility, captures many external and internal market factors ranging from agricultural structure and consumption, to the effect of plagues, trade, and monetary shocks. The result is that this measure captures factors that may differ among countries, but that do not directly influence market efficiency as such. In the next sections, we will look at the traditional measures and show that statistics reflecting the spread of a variable around its arithmetical average (unconditional mean) contains a

[3] Alternatively, we could say that standard volatility measure tells us a lot about market integration but does that in such a noisy and incomparable way that we cannot rely on them.

number of region/culture specific factors unrelated to the defined efficiency. As such, traditional measures based on the unconditional mean like the CV or the standard deviation of log prices (being comparable to the CV[4]), are not comparable over time and space. Instead we suggest using a conditional heteroscedasticity methodology to draw conclusion on the market integration of different societies in different periods and to compare them from this perspective. This methodology allows us to see if markets managed to improve on their handling of the effects of unexpected price shocks.

As a first step it is important to precisely define what is meant by efficiently working markets. We adopt a very general definition of risk in commodity markets. By risk we mean the degree of uncertainty regarding future prices. This is often approximated by the degree of volatility of price series, which is deductable from the expectation that if risk management techniques are efficient, they will reduce uncertainty and thereby result in smoother prices.[5] Generally speaking, one may expect such reduction in price volatility to arise from four major types of risk management techniques.

1. Intertemporal risk reduction (storage for example), which can reduce seasonal or even cyclical movement of prices. This oldest form of risk management in the markets of basic foodstuffs possibly saw the earliest large scale involvement of the state in the

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[4] That the two measures of dispersion are almost equivalent can be shown easily. Let d be equal the deviation from the mean: $d_t = y_t - \bar{y}$. Now the Coefficient of Variation can be expressed as follows: $CV = \sqrt{\frac{\sum_{t=1}^{T}(y_t - \bar{y})^2}{n}}\frac{1}{\bar{y}} = \frac{\sqrt{\sum_{t=1}^{T} d_t^2}}{\sqrt{n}\bar{y}}$ while the standard deviation of the log series is the following: $\sigma_{\ln y} = \sqrt{\frac{\sum_{t=1}^{T}\left(\ln y_t - \overline{\ln y}\right)^2}{n}} = \frac{\sqrt{\sum_{t=1}^{T}\ln\left(\frac{y_t}{\bar{y}}\right)^2}}{\sqrt{n}} = \frac{\sqrt{\sum_{t=1}^{T}\ln\left(1+\frac{d_t}{\bar{y}}\right)^2}}{\sqrt{n}} \approx \frac{\sqrt{\sum_{t=1}^{T}\left(\frac{d_t}{\bar{y}}\right)^2}}{\sqrt{n}} = \frac{\sqrt{\frac{1}{\bar{y}^2}\sum_{t=1}^{T} d_t^2}}{\sqrt{n}} = \frac{\sqrt{\sum_{t=1}^{T} d_t^2}}{\sqrt{n}\bar{y}}$. In the last derivation we made use of the approximation that for small values of x, $\ln(1+x) \approx x$.

[5] We also use the expression 'market related institutions' in this text, by which we mean those institutions that foster a more efficient flow of information among agents, secure property rights, and reduce transaction costs in a particular market (North 1991). Obviously, it is not essential to link institutions with price volatility in this paper, but one may expect that unless a major technological innovation or some profound change in the natural environment took place, a long-run reduction in price volatility should be a result of improving market-related institutions.

Ancient East. This practice continued during the medieval and early modern periods although it has been questioned how important it was (Fenoaltea, 1976; McCloskey and Nash, 1984; Komlos and Landes, 1991; Will *et al.*, 1991; Poynder, 1999).

2. Spatial market integration (or spatial diversification) that reduces volatility through linking different coexisting markets by means of trade. Obviously, if different regions trade with each other, price equalization may reduce the price effect of idiosyncratic shocks. Improving international relations (trade liberalization, longer periods of peace) or technological development of transportation (for example large cargo vessels, refrigeration). This type of market integration has been subject of several studies (i.e. Jacks 2005; Ó Gráda 2005; Keller and Shuie 2007; Özmucur and Pamuk, 2007; Federico and Persson, 2007; Studer 2008).
3. Increased diversification of the consumption structure. If consumers have a wide range of substitutes to choose from, product specific shocks will have lower effect on prices. We can find several historical examples for diversifying consumption through discoveries and long-distance trade, like the introduction of maize, rice and potato in Europe, or by technological development (margarine as a substitute for butter) (Ó Gráda 1992; Reis 2005).
4. Finally, innovations through product development may also contribute to more stable yields and less volatility. Even though one would associate this immediately to genetic engineering, selective breeding can be seen as an early version of product development in historical times (i.e. Overton 1996, 113-114).

The above mentioned techniques altogether contributed to the reduction of uncertainty of the prices of basic foodstuff that were the main commodities prior to the Industrial Revolution. The challenge is to find a way to quantify the efficiency of such techniques in different societies in different periods of time in a comparable manner.

As pointed out, more efficient risk management techniques should lead to less variance in prices. Hence, measures of dispersion of prices are often used to capture market efficiency. The problem with the CV and related measures, however, is that they include more than just the effects of risk reducing techniques listed above. They also include a volatility component arising from different agricultural structure, inflation, and different demand structure. All these factors as such, although clearly influencing volatility, do not have a direct relation with the market efficiency as market related institutions cannot influence them directly. They affect the unconditional variance though; hence they have to be filtered out if one wants to compare market efficiency over time and across regions/countries.

These factors may influence the estimated market efficiency either directly, or via the trend. In the next section, we will show that if there is a trend in the series, or if the prices series has unit root (not unexpected in case of prices), the longer period we choose the higher the calculated dispersion is going to grow. As such, without detrending the series first, this exercise is faulty. Yet, even if we remove the trend, the variance of the series will have components that are region and period specific and thus makes it very difficult if not impossible to compare dispersion measures across different regions. The effect of the trend will be discussed in Section 3 while the region- and time specific components of the variance will be dealt with in Section 4.

## 3 The effect of trend on the variance of time series

Above we pointed out that the timing of CV estimates may seriously distort the picture in the face of trends in the series, for example caused by inflation. In the followings we examine two possibilities: a stationary time series with linear trend (deterministic trend model), and a unit root process with or without trend (that is with or without a drift parameter).

Let us assume that our series is stationary and can be modeled as follows:

$$y_t = \beta_0 + \beta_1 X_t + \gamma t + u_t \quad (1)$$

where $t = 1, 2..., T$ and $u_t \sim IID(0, \sigma_u)$, X is some exogenous variable (Cov(X,u)=0).

The variance of y around its arithmetical mean (unconditional variance) is:

$$Var(y) = \beta_1^2 Var(X) + \gamma^2 Var(t) + 2\beta_1 \gamma Cov(X,t) + Var(u) \quad (2)$$

For simplicity, let us assume that the variable X has no trend, so Cov(X,t)=0. In that case the variance of y will depend on time in a monotonous way only through the variance of the time trend. The variance of the trend can be expressed as follows (Hamilton 1994: 456):

$$Var(t) = \frac{1}{T}\sum_{t=1}^{T} t^2 - \frac{1}{T^2}\left(\sum_{t=1}^{T} t\right)^2 \text{ where } \sum_{t=1}^{T} t^2 = T(T+1)(2T+1)/6 \text{ and } \sum_{t=1}^{T} t = T(T+1)/2 \quad (3)$$

, which results in the following: $Var(t) = \frac{1}{12}\left(T^2 - 1\right)$. Clearly, if we choose a longer period to estimate the variance of our series (T grows), it will inflate the variance of y. Note that this effect is independent of the sign of the time trend.

Yet, above example solely focuses on a deterministic trend model. Let us now take a more likely case, when our series are not stationary. This is expected in case of price series, especially when the conditions of a weak form of efficient markets are fulfilled. This means that no one can "outsmart" the market using public (or past) information to make a profit. If such information is available, in a well functioning market agents will immediately use the information and eliminate the extraordinary profit. In short, price changes are not predictable from past information, and our best guess for the next period's price is the current price. Such a series is called random walk (with no trend) or random walk with drift (if there is a trend) and is found even in ancient economies such as Babylon ca. 200BC (Temin 2002).

Let's start with the following equation:

$$y_t = \gamma + y_{t-1} + u_t \quad (4)$$

where $\gamma$ is the drift parameter ($\gamma$=0 means no trend) and $u_t \sim IID(0,\sigma_u)$. By repeated substitution we arrive at the following expression: $y_T = \gamma T + \sum_{t=1}^{T} u_t$. Since the effect of past innovations will not wear off in this model, their effect on the variance will accumulate:

$Var(y_T) = \sum_{t=1}^{T} Var(u_t) = T \cdot Var(u)$. Again we find that with the presence of unit root, the longer period we choose to calculate variance and the derived dispersion measures, the more "volatility" we will find. As such, these measures are seriously misleading. For example, the 16th century witnessed massive inflation (Munro 2003). As can be seen in Table 1, this means that the CV is inflated by the underlying trend in price level. However, this is by no means an indication of deteriorating institutions or market disintegration in this period. As we show in the following section, simple de-trending techniques like using a deterministic trend or using the CV of the first difference of prices are not necessarily ideal solutions either.[6]

## 4 The effect of structural differences on the variance of price series

The trend clearly has an important effect on volatility measures, either arising from inflation, technology, or other factors. Yet, even if one removes the trend form the series, the lack of comparability still remains an issue. In order to see the reason, one needs to think in terms of a structural model of prices, which ultimately is determined both by those factors reducing risk (as outlined in section 2) and by those factors that are instrumental to the economies, but not intended to reduce risk as such like the agricultural structure, the demand structure and inflation (see section 1). Here, the last variables, of course, can be predicted (and hence

[6] Let us assume that the series is close to having a unit root but still a stationary process, for example, $y_t = \lambda_0 + \lambda_1 y_{t-1} + \lambda_2 x_t + e_t, 0 < |\lambda_1| < 1$ then first differencing will neither remove the effect of the previous period prices nor that of the exogenous variables: $\Delta y_t = (\lambda_1 - 1) y_{t-1} + \lambda_2 x_t + e_t$. That is, simply taking the CV or the variance of a first differenced price series may still contain a lot of different factors besides the residual variance and the problem outlined in section 4 are still valid. The problem can be solved through first differencing only if one is sure (by means of testing) that the DGP is like in (4.).

captured by a structural model), while the first ones manifest themselves by the reduced effect of external shocks.

Let us assume that the price and quantity of commodity *i* is determined by the following system of supply and demand equations:

$$q_{i,t}^{S} = \sum_{j=1}^{k} \beta_j X_{j,t} + u_t \text{ supply, } u_t \sim IID(0,\sigma_u) \quad (5)$$

$$q_{i,t}^{d} = \gamma p_{i,t} + \sum_{m=1}^{r} \delta Z_{m,t} + v_t \text{ demand, } v_t \sim IID(0,\sigma_v) \quad (6)$$

where $q^S$ and $q^d$ denotes the quantity supplied and demanded, *X* and *Z* are different factors that affect supply and demand respectively. *X* may contain different exogenous factors like wars, weather conditions or lagged price of commodity i (capturing expectations regarding present period prices), while *Z* denotes variables like the price of other products that are either substitutes or complements of commodity *i*, or exogenous factors affecting demand, like income or different political variables.

In equilibrium $q^S = q^d$, hence we can express the reduced form for the price of commodity i as follows:

$$p_{i,t} = \sum_{j=1}^{k} \frac{\beta_j}{\gamma} X_{j,t} - \sum_{m=1}^{r} \frac{\delta_m}{\gamma} Z_{m,t} + \frac{u_t - v_t}{\gamma} = \sum_{j=1}^{k} \Pi_j X_{j,t} - \sum_{m=1}^{r} \Theta_m Z_{m,t} + e_t ,\ e_t \sim IID(0,\sigma_e) \quad (7)$$

Assuming that the variables *X* and *Z* are indeed exogenous (or at least predetermined), *Cov(X,e)=Cov(Z,e)=0*. The unconditional variance of $p_i$ is:

$$Var(p_{i,t}) = \sum_{j=1}^{k} \Pi_j^2 Var(X_j) + \sum_{m=1}^{r} \Theta_m^2 Var(Z_m) + \sum_{j<l} 2\Pi_j \Pi_l Cov(X_j, X_l) + \sum_{m<n} 2\Theta_j \Theta_n Cov(Z_m, Z_n) +$$
$$+ \sum_{j=1}^{k} \sum_{j<m}^{r} 2\Pi_j \Theta_m Cov(X_j, Z_m) + Var(e_t)$$

(8)

That is, the variance of the prices, and so all derived measures, like the Coefficient of Variation will contain not only the effect of shocks on prices (*Var(e)*), but also that of the price volatility of related products, exogenous factors and their relations measured by the covariances. The latter is important, since these reflect region specific conditions and differences in main crops (relationship between weather conditions and the yield of different crops), tastes (relationship among the prices of different products) and the diversification of the consumption structure (the number and sign of price covariances that enter the equation). If, for example, we have a society where 80% of the income is spent on two main products with different environmental requirements/sensitivity, we can expect that in that particular region the price covariance of the two main staples will be lower than in a country where the two main products are similar (see wheat-rice in the North of China vs. wheat-barley in Europe).

Undeniably, the variance of prices contains basically all information we need to know to judge the efficiency of market related institutions, but it also contains a lot of region and period specific factors, that are not possible to separate. So, if one compares the CV of different regions, or within one region over time, even if the trend and the non-stationarity has been taken care of, one still cannot know exactly what is measured. One could of course estimate a structural equation (at least the reduced form price equation) and use the coefficients to have some estimates of the effect of different factors, but in most cases we do not have such detailed set of exogenous variables available for historical analysis. For this reason, we suggest not using CV and related measures on time series to draw conclusions on market efficiency or market integration in a particular region or country over time.

## 5 Alternative methodology: conditional heteroscedasticity

Our suggestion to get around the problem is to apply a technique widely used in financial econometrics and macroeconometrics, a conditional heteroscedasticity model, which is based on the standard autoregressive models that are occasionally used in market efficiency studies (i.e. Soderberg 2004). The unconditional variance of price series, as we saw above, consists basically of two parts: the variance of the residual term, and the variance of the conditional mean (fitted value) around the arithmetical average. The latter has a lot of information, but the lack of data will prevent us from separating what we need from the region/period specific factors within a structural model. With conditional heteroscedasticity methods, we concentrate on the variance around the conditional mean, that is, the residual variance only.

The residual variance will reflect the share of shocks in total variance, that is the effect ofunexpected events (unexpected since the residual has no significant autocorrelation in a correctly specified model) on price volatility. If markets are more integrated or the institutional background is more efficient then these unexpected events should have a lower effect on prices. The error is of course a random variable, so its magnitude may change. For our purposes the residual can be modeled as follows:

$$e_t = \chi(\Omega_t)\cdot\varepsilon_t \quad (9)$$

where

$$\varepsilon_t \sim IID(0,\sigma_\varepsilon) \quad (10)$$

, where ε is a random variable representing the size of the shocks, and χ is a multiplier showing us the effect of the shock on prices. The factor χ may obviously depend on the degree of market integration or other factors (denoted by $\Omega_t$) and in case of a tendency of improvement or deterioration it should be dependent on time. As for the size of shocks, we can assume that they are on average zero, and homoscedastic, in other words, the magnitude of the shocks does not depend on time or the order of the observations. Of course, there may be periods when shocks are larger (wars, disasters, sudden changes in climate) but if we take a reasonable long sample, this condition should hold. This also means that in order to have meaningful results on the efficiency of market and related institutions to cope with risks, one should not analyze short periods. If we assume that the magnitude of the shocks and the parameter $\chi$ are independent, we obtain for the residual variance:

$$Var(e) = Var(\chi_t) \cdot Var(\varepsilon_t) + \overline{\chi} Var(\varepsilon_t) = (Var(\chi_t) + \overline{\chi}) \cdot \sigma_\varepsilon^2 \quad (11)$$

If we find a trend-like behavior in the residual variance (or any similar measures, like the standard deviation or the mean absolute deviation of the residual), we can interpret it as a sign of the market's ability to cope with the effect of shocks. The cornerstone of the suggested method is our assumption regarding the random shocks. While our assumption about the homoscedasticity of ε within the same country or region sounds feasible, it is much less likely that the shocks have the same variance over all regions as well, which makes this method not ideal for cross-country comparisons. Still, acknowledging this weakness of the proposed method, it should have much less incomparable factors than the existing methodology. A possible solution might be to apply a model which allows of country specific differences in the error-term, combining the conditional heteroscedasticity models with a panel analysis. This direction is however not pursued in this paper but is preferable when one has enough observations to make a panel dataset.

There are several different approaches to model conditional heteroscedasticity. The early methodology was suggested by Engle (1982), where the estimation is carried out in two steps. In the first step one models the conditional mean of the time series (mean equation) as an AR(p) model (or a structural model) so that there remains no serial correlation in the residuals (there is a heteroscedasticity of course, so robust standard errors are required for model selection), and the squared residuals are modeled in the second step as an AR process (variance equation). This leads to the well-known ARCH specification.

$$y_t = \gamma_0 + \sum_{i=1}^{p} \gamma_i y_{t-i} + u_t \tag{12}$$

and

$$\hat{u}_t^2 = \phi_0 + \sum_{j=1}^{q} \phi_j \hat{u}_{t-q}^2 + \sum_{l=1}^{r} \Omega_{l,t} \tag{13}$$

where Ω denotes the effect of r exogenous factors that may explain the heteroscedastic nature of the residual.

Bollerslev (1986) suggested a different specification that may avoid the problems with the ARCH specification and lead to more efficient estimates, called the Generalized Autoregressive Heteroscedasticity (GARCH).

$$\sigma_{u,t}^2 = \phi_0 + \sum_{k=1}^{n} \eta_k \sigma_{u,t-k}^2 + \sum_{j=1}^{q} \phi_j \hat{u}_{t-q}^2 + \sum_{l=1}^{r} \Omega_{l,t} \tag{14}$$

Obviously the second specification cannot be estimated with OLS, so one needs to rely on a Maximum Likelihood estimator and estimate the two equations simultaneously. Unfortunately, the second approach may be unsuitable for historical analysis, since we often have gaps in the data and the ARCH/GARCH procedures in most econometric packages cannot handle missing data. In those cases, it might be more useful to return to the two-step method using ARCH specification. Since in this paper we are looking for some long-run tendencies of the residual variance to decrease, we use a time trend, and test for structural breaks. If one has no gaps in the series, it is preferable to use the ML estimator. In practice other ARCH/GARCH specifications might also be explored, like the TARCH specification allowing for different effect of shocks with different signs, or GARCH-M (Glosten et al 1993) specification allowing for an effect of volatility on price level. In this paper, however, in order to make all estimations comparable, we use uniformly an ARCH(1) specification.

## 6 Empirical application

In this section we will illustrate above method by appying a conditional heteroscedasticity model to four historical price series: two monthly (wheat prices in Pisa 1549-1716 [Malanima 1976][7] and Paris 1548-1698 [Baulant and Meuvret 1962; Poynder 1999]) and two annual series (rice prices in Hiroshima, Japan, 1620-1857 [Iwahashi 1981] and wheat prices in Vienna 1439-1800 [Pribram 1938]). Since the volatility measures are not comparable if the series have different frequency, the monthly series are transformed to annual series by taking average of the monthly observations, when compared with the annual series. Since it is possible that using annual series would distort our results, the same exercise is done with the monthly series as well, and we find that the underlying trend in the residual variance does not change when different frequency data is used. The Vienna and the two monthly series have gaps in the data. In the case of the Vienna and the Paris prices where to many gaps were present, we use the two-step procedure suggested by Engle (1982). The monthly wheat prices from Pisa, have a single gap in 1631, making it possible to estimate the ARCH model on the half of the series separately and use an ML estimator. In order to make the results from the different series comparable, as noted in the previous section, in all cases we use an ARCH(1) specification. We would like to stress that decision reflects only our preference for comparability, but in actual empirical usage one should rather select models for their fit or because of some theoretical considerations. We experimented with other specifications as well, and it turned out that using a more complex structure to model residual variance improved the efficiency of the estimation (lower standard errors), but did not fundamentally change the results.

[7] The original series run until 1816, but since we have found a structural break around 1716, and so that the results are more comparable with the Paris series, we decided to limit the sample to 1550-1716.

We start with reporting the annualized prices in below graphs. All series exhibit large

Figure 1

Figure 2

Figure 3

Figure 4

price increases in the 16th century which is consistent with the influx of silver from the Americas and, in the case of Japan, a strong increase in minting of silver coins (Miyamoto 1999, 59). Besides this common rise in prices, it is also clear that inter-annual volatility in all four series is strongly different.

Now, we start by estimating the standard measures of dispersion for the annual series (Table 2). If one based her or his judgment about the degree of market

Table 2

integration/risk management on the traditional measure, one would find that the wheat market of Pisa was especially efficient, resulting in a small volatility of prices, while Paris and Hiroshima seem to have roughly equal volatility, with Vienna being the last.

Indeed, although volatility is Paris and Hiroshima was of similar magnitude, the underlying reason for this volatility is different. It has often been argued that in Paris markets were extremely volatile at that period. Food shortages, an over-emphasizing of arable agriculture at the cost of pasture, and inflation all contributed to increasing price volatility. In the period 1550-1600 the real price of wheat even trebled (Knecht 2001, 260). In Japan, on the other hand, it has been argued that during the seventeenth century rising population pressure caused the shogunate to create a policy to keep the rice price down in the form of demand suppressing policies such as a limit on sake brewing (Miyamoto 1999a, 57). From the late 17th century these pressures eased because of greater agricultural production. Hence, except for the Tenmei crop failures (1781-1789), prices remained relatively stable.

Viennese price fluctuations, however, were much higher. On the one hand, Vienna had relatively little problems with a supply side crisis because of its status as a main city in the Habsburg Empire (Weigl, 2000b, pp. 162-163). On the other hand, Vienna experienced an extremely strong population growth in combination with the fact that wheat was the most important staple food (Sandgruber (1982, p. 141). Not only did this increase the share of the poor (Weigl 2000, pp. 198-199; Weigl 2001, p. 51), but also made it more susceptible for supply side shocks. The most important of these supply side shocks in the presence of Turkish in Austria. In 1683 they even sieged Vienna, which is clearly to be seen in the variance of the price series, after which variance is lower. Finally, the Pisa series reflect the lowest volatility. Partly that is caused by little change in population and agricultural structure, and partly by the relative constancy of consumption.

Hence, in all countries we find that exogenous factors like inflation, agricultural structure, and demand played a role in price volatility but these are not be necessarily directly related to risk management techniques. In Paris and Japan this seems to have been almost to the same degree. Both were subject to inflation and government intervention driving demand patterns but in essence no serious disturbances took place. In Vienna, however, the main source of inflation consisted of external shocks and government policy that had a profound impact on volatility. This suggest that, after removing that country specific effects, volatility in Vienna will decline more than in Hiroshima and Paris. In Pisa, however, only small effects of region-specific variability took place, causing a correction to have only a small effect.

Besides the annual volatility, for above estimates it is also clear that the length of the period differ by sample which implies higher CVs for longer series. As suggested above, one could argue that we should look at smaller periods, thereby reducing the effect of the trend on variance. But as we showed in the previous sections, the CV has so many fallacies when used of measuring market integration that we do not pursue this direction. Instead we turn to the conditional heteroscedasticity models. The first step of modeling such a conditional heteroscedasticity is to establish if the log of the price series are stationary. In Table 3 we report the results form three unit root tests, all of them are designed to improve the low power of the traditional Dickey-Fuller type unit root tests when rejecting the null hypothesis of non-stationarity. The test statistics are in most cases consistent with each other implying that the price series are stationary. The only case when there is contradiction among the test results is the annual log prices of wheat in

Table 3

Pisa: the Philips-Perron test reject the null hypothesis of non-stationarity at 1 , while the DF-GLS test suggest that the series are not stationary. The Ng-Perron test is again contradictory. Hence we choose to treat this series as non-stationary and take its first-difference before estimation of the ARCH model.

For the annual series for Paris, Pisa, Hiroshima we could use the standard ML procedure to estimate the mean and the variance equations simultaneously. In case of the Vienna series, the gaps do not allow us to do so, so there we first estimate the expected value of the log of wheat prices and then we model the square of the residuals in a second step.

Table 4

As Table 4 suggests, the series in most cases have a positive trend (except for Pisa), so our critique in section 3 is likely to be applicable to these series. With the exception of the Pisa series the residuals does not turn out to be normally distributed but with the large number of observations this should not pose any problems. The residuals does not have any significant autocorrelation either.

In all cases In case of Vienna we found the square of the residuals to have no significant serial correlation, while a linear trend yielding a significant, negative coefficient. That is the reason why the conditional standard deviation in Figure 8 is so smooth.

In Figure 5 to 8 we plot the conditional standard deviation of the annual price series for the four cities. In Paris, Pisa and Vienna we find a clear downward trend. Even though the standard deviations start out at different levels (the highest in Vienna with 0.4, and the lowest in Hiroshima with a bit above 0.16) the reduction is quite conspicuous. In case of Pisa and Vienna the standard deviation is halved, while in Paris it is reduced by roughly 60%. Hiroshima seems to be the exception: the time trend in the variance equation was not significant and had a positive coefficient. This is not surprising, however, given that the early volatility as discussed before was caused by the demand structure and inflation, which were both factors that need to be removed as they have no direct relation with the risk factors in

Figure 5

Figure 6

Figure 7

Figure 8

markets (Miyamoto 1999b, 120).

So far, we discussed the annual series. As a second step, we use the monthly prices series to see if using a different frequency data alters the results. As we noted at the beginning of this section, the Pisa series has just a single gap, making it possible to use ML estimator at the cost of cutting the sample in two. Incrase of the Paris prices however, we had 3 gaps, lasting only for a few months but estimating an ARCH model on 4 different, shorter periods would strongly contradict our own suggestions in section 5 on the length of the sample period. With shorter periods chosen the probability that what we really observe is not some changes in the working of markets but rather some temporary fluctuations in the random shocks would be much too high. The regression results are reported in Table 5.

Table 5

Even though some of the statistics are necessarily changed with the different frequency of data used, the picture has not changed much. We find a significant negative trend in the residual variance (or standard deviation) in all three cases. The degree of improvement is slight in Paris, while in Pisa we find a significant improvement in the 16$^{th}$ century and the first decades of the 17$^{th}$ century, followed by just some minor improvement throughout the rest of the sample period (Figuere 9, 10a and b). Using annual data therefore does not seem to be misleading.

Figure 9

Figure 10a and 10b

**Conclusion**

In this paper we follow part of the literature which argues that using traditional dispersion measures, like the Coefficient of Variation as indicators for the degree of market integration is misleading because it includes country and time-specific factors that not directly influence market institutions. This has two empirical consequences. First, in case of non-stationary price series or even if the series are trend stationary, the variance is inflated by the trend. This leads to erroneous conclusion regarding the degree of market integration in the presence of inflation, like the Price Revolution during the 16th century. Secondly, the variance of the log prices contains information on a lot of factors that are either not related to market integration, but rather region specific, or are not comparable or separable. Hence both the standard CV and the detrended version will bias the estimates of market efficiency. As an alternative, we suggest using the standard deviation of the residual (or the standard error of the regression) to compare the degree of market integration among different regions, together with a conditional heteroscedasticity approach to test for any improvement over time.

Applied on four heterogeneous price series (Pisa, Hiroshima, Paris, and Vienna), we find that the CV results in all cases in a higher level of volatility since it measures unconditional volatility. Hence, using the CV or some related measure of unconditional volatility of commodity price series is likely to lead biased, noisy estimates of market efficiency. Unfortunately, the amount of overestimation depends on the presence of factors like inflation, the demand structure, and agricultural structure which, although influencing price volatility, are not directly related to the price reducing effects on risk strategies within markets. These factors turn out to be much higher in Vienna than in other cities, hence biasing a cross regional comparison. Equally, we find that on average in Europe price volatility declines (and market efficiency increases) since the early 16th century, which compares well with the established view of increasing market efficiency from ca. 1600 onwards (Persson 1999; Jacks 2004).

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**Table 1** Coefficient of variation (CV) of barley per period and country

| country | Period | Product | Mean | std deviation | CV | Unit | Source |
|---|---|---|---|---|---|---|---|
| Babylon | 385- 250BC | barley | 19.807 | 19.032 | 0.961 | shekel/100 l | Slotksy 1997; Vargyas 2001 |
| Babylon | 250 -150BC | barley | 10.900 | 7.088 | 0.650 | shekel/100 l | Slotksy 1997; Vargyas 2001 |
| Babylon | 150 - 61BC | barley | 21.263 | 17.518 | 0.824 | shekel/100 l | Slotksy 1997; Vargyas 2001 |
| England | 1209-1347AD | barley | 0.408 | 0.154 | 0.378 | shilling/bu | Clark (2004) |
| England | 1350-1500AD | barley | 0.423 | 0.147 | 0.347 | shilling/bu | Clark (2004) |
| **England** | **1500-1650AD** | **barley** | **1.264** | **0.800** | **0.633** | **shilling/bu** | **Clark (2004)** |
| England | 1650-1800AD | barley | 2.393 | 0.696 | 0.291 | shilling/bu | Clark (2004) |
| England | 1800-1900AD | barley | 4.300 | 1.069 | 0.249 | shilling/bu | Clark (2004) |
| Florence | 1325-1347AD | barley | 10.183 | 3.278 | 0.322 | denier/setier | de La Roncière (1982) |
| **Modena** | **1554-1650AD** | **barley** | **98.753** | **49.927** | **0.506** | **soldi/staro** | **Basini (1974)** |
| Modena | 1650-1700AD | barley | 120.000 | 31.503 | 0.263 | soldi/staro | Basini (1974) |
| **Vienna** | **1500-1650AD** | **barley** | **28.834** | **24.558** | **0.852** | **kreuzen/metzen** | **Pribram (1938)** |
| Vienna | 1650-1800AD | barley | 61.299 | 28.419 | 0.464 | kreuzen/metzen | Pribram (1938) |

Table 2

Coefficient of Variation and the standard deviation of the log prices

| | Paris wheat prices | | Hiroshima rice prices, annual 1620-1857 | Pisa white wheat prices | | Vienna wheat prices, annual, 1439-1800 (with gaps) |
|---|---|---|---|---|---|---|
| | annual 1549-1697 | monthly Sept 1549 – Aug 1698 | | annual 1550-1817 | monthly Oct 1548-Jul 1818 | |
| CV | 0.497 | 0.543 | 0.409 | 0.392 | 0.412 | 0.683 |
| std dev of the log series | 0.503 | 0.522 | 0.435 | 0.350 | 0.373 | 0.753 |

Table 3

Unit root tests (with linear trend)

| | Philips Perron | DF-GLS | Ng-Perron test |
|---|---|---|---|
| Paris log of wheat prices 1549-1697 (annual) | -4.388 (p<0.01) | -3.389 (p<0.05) | MZa=-20.23 (p<0.05)<br>MZt=-3.179 (p<0.05)<br>MSB=0.157(p<0.01)<br>MPT=4.514 (p<0.01) |
| Paris log of wheat prices 1549M09-1698M09 (monthly) | -6.552 (p<0.01) | -3.710 (p<0.01) | MZa=-27.22 (p<0.01)<br>MZt=-3.67 (p<0.01)<br>MSB=0.135(p>0.01)<br>MPT=3.45 (p>0.01) |
| Pisa log of wheat prices 1550-1817(annual) | -4.856 (p<0.01) | -2.578 (p>0.1) | MZa=-16.78 (p<0.1)<br>MZt=-2.889 (p<0.1)<br>MSB=0.172 (p<0.05)<br>MPT=5.479 (p<0.1) |
| Pisa log of wheat prices(monthly), 1548M10-1631M07 | -6.240 (p<0.01) | -6.141 (p<0.01) | MZa=-74.14 (p<0.01)<br>MZt=-6.07 (p<0.01)<br>MSB=0.082(p>0.01)<br>MPT=1.308 (p>0.01) |
| Pisa log of wheat prices(monthly), 1631M10-1818M07 | -3.539 (p<0.01) | -3.471 (p<0.05) | MZa=-27.58 (p<0.01)<br>MZt=-3.70 (p<0.01)<br>MSB=0.134(p>0.01)<br>MPT=3.383 (p>0.01) |
| Hiroshima log of rice prices 1620-1857 | -5.423 (p<0.01) | -3.626 (p<0.01) | MZa=-24.65 (p<0.01)<br>MZt=-3.510 (p<0.01)<br>MSB=0.142 (p>0.01)<br>MPT=3.697 (p>0.01) |
| Vienna log of wheat prices, 1439-1800 | -4.593 (p<0.01) | -3.928 (p<0.01) | MZa=-31.38 (p<0.01)<br>MZt=-3.956(p<0.01)<br>MSB=0.126(p>0.01)<br>MPT=2.931 (p>0.01) |

Table 4

Conditional heteroscedasticity estimates annual series

| | Paris 1549-1697 | Pisa 1550-1817 (differenced series) | Hiroshima 1620-1857 | Vienna 1439-1800 |
|---|---|---|---|---|
| mean equation | | | | |
| Constant | 1.228<br>(5.44) | 0.006<br>(1.12) | -1.953<br>(-16.4) | 2.691<br>(8.85) |
| AR(1) | 0.824<br>(8.28) | 0.057<br>(0.86) | 0.679<br>(7.99) | 0.738<br>(8.64) |
| AR(2) | -0.368<br>(-2.80) | -0.403<br>(-6.63) | - | - |
| AR(3) | 0.200<br>(2.33) | -0.184<br>(-2.80) | - | - |
| AR(4) | - | -0.111<br>(-1.96) | - | - |
| trend | 0.028<br>(4.70) | - | 0.011<br>(4.03) | 0.0066<br>(6.24) |
| $trend^2$ | -0.000138<br>(-3.73) | - | -0.0000295<br>(-2.35) | - |
| variance equation | | | | |
| constant | 0.046<br>(2.69) | 0.032<br>(6.20) | 0.026<br>(3.92) | 0.135<br>(3.95) |
| ARCH(1) | 0.292<br>(1.86) | 0.095<br>(1.30) | 0.201<br>(1.83) | - |
| trend | -0.0000777<br>(-0.43) | -0.0000742<br>(-3.04) | 0.0000306<br>(0.58) | -0.000245<br>(-2.19) |
| diagnostic tests | | | | |
| $R^2$ | 0.769 | 0.189 | 0.800 | 0.858 |
| Q(5) | 2.816<br>(p=0.245) | 0.840<br>(p=0.359) | 6.615<br>(p=0.158) | 3.495<br>(p=0.479) |
| Q(12) | 6.462<br>(p=0.693) | 14.359<br>(p=0.073) | 9.668<br>(p=0.560) | 6.579<br>(p=0.832) |
| Jarque-Bera test of the normality of the residual | 9.316<br>(p=0.009) | 1.027<br>(p=0.598) | 49.98<br>(p=0.000) | 30.62<br>(p=0.000) |
| std. error of the regression | 0.241 | 0.159 | 0.194 | 0.278 |

Note: z-statistics are reported in parentheses, with the exception serial correlation and normality tests, where p values are reported.

Table 5
Conditional heteroscedasticity estimates motnhly series

| | Paris log of wheat price[a] | Pisa log of wheat price | |
|---|---|---|---|
| | 1549M09-1698M09 | 1548M12-1631M07 | 1631M12-1818M07 |
| | | | |
| Constant | 1.507<br>(8.50) | 2.849<br>(43.2) | 2.634<br>(25.0) |
| AR(1) | 0.930<br>(30.8) | 1.138<br>(31.3) | 1.059<br>(41.3) |
| AR(2) | -0.039<br>(-0.68) | -0.158<br>(-4.40) | -0.051<br>(-1.89) |
| AR(3) | 0.101<br>(1.82) | - | -0.091<br>(-5.88) |
| AR(11) | -0.043<br>(-4.23) | - | 0.069<br>(4.90) |
| trend | 0.002<br>(4.64) | - | - |
| $trend^2$ | -0.000000681<br>(-3.39) | - | - |
| | | | |
| constant | 0.072<br>(13.82) | 0.0104<br>(21.27) | 0.0028<br>(18.3) |
| ARCH(1) | 0.238<br>(4.24) | 0.176<br>(5.46) | 0.279<br>(9.09) |
| trend | -0.0000000811<br>(-1.71) | -0.00000096<br>(-17.6) | -0.000000127<br>(-1.89) |
| | | | |
| $R^2$ | 0.960 | 0.951 | 0.972 |
| Q(5) | 9.771<br>(p=0.002) | 2.288<br>(p=0.130) | 6.849<br>(p=0.009) |
| Q(12) | 18.51<br>(p=0.018) | 10.71<br>(p=0.219) | 9.599<br>(p=0.294) |
| Jarque-Bera test of the normality of the residual | 34004<br>(p<0.1) | 1.716<br>(p=0.424) | 1826<br>(p<.01) |
| std. error of the regression | 0.100 | 0.174 | 0.059 |

Note: seasonal dummies are included but not reported in the mean equation
[a] Due to outliers the dependent variable in the second stage (variance equation) is the absolute value of the residual form the first step (mean equation)

Figure 1
log of annual wheat prices in Paris, 1548-1698

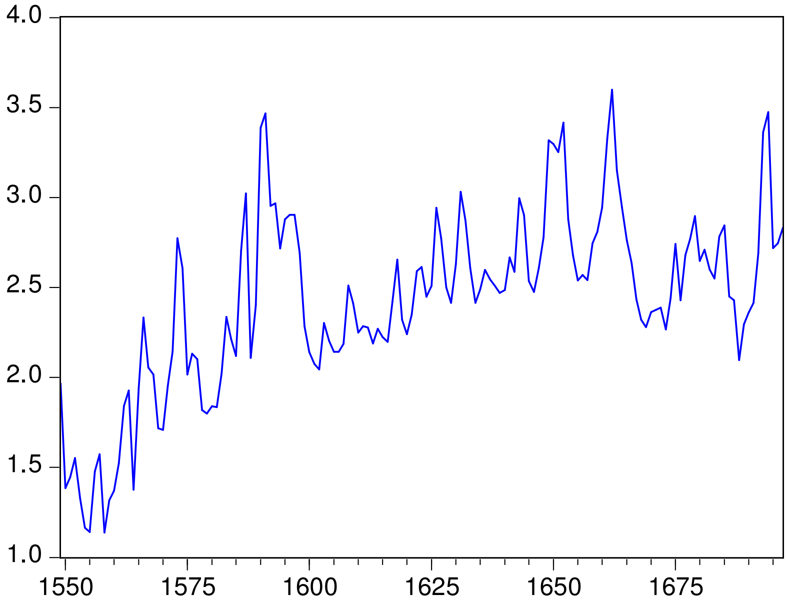


Figure 2
log of annual wheat prices in Pisa, 1549-1716

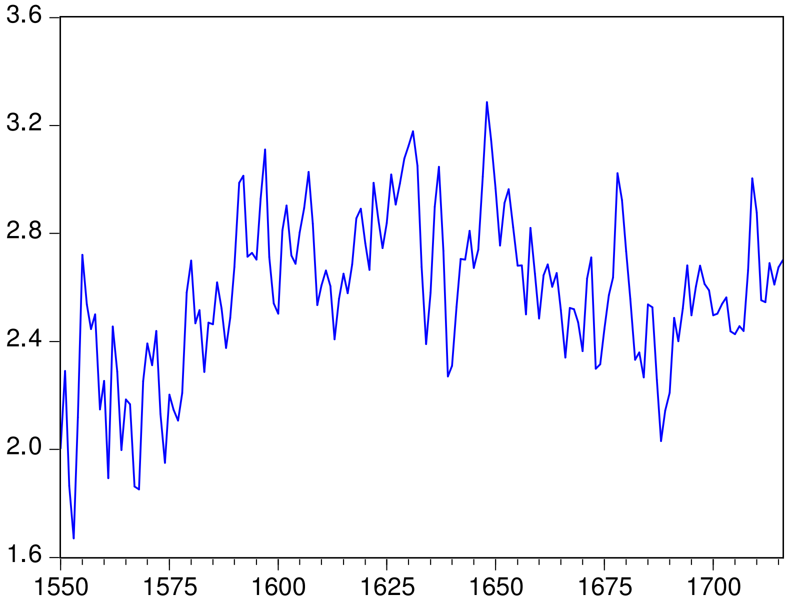

Figure 3
Log of annual rice prices in Hiroshima, 1620-1857

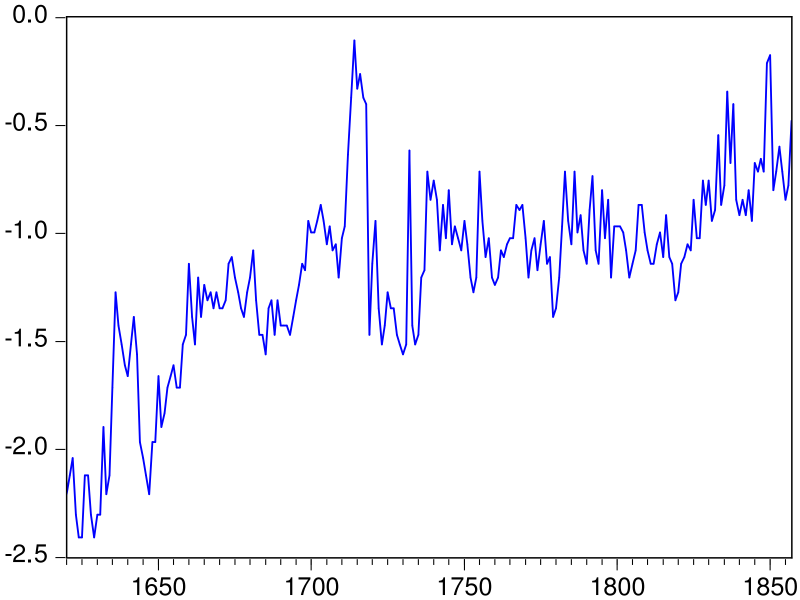


Figure 4
log of annual wheat prices in Vienna 1439-1800

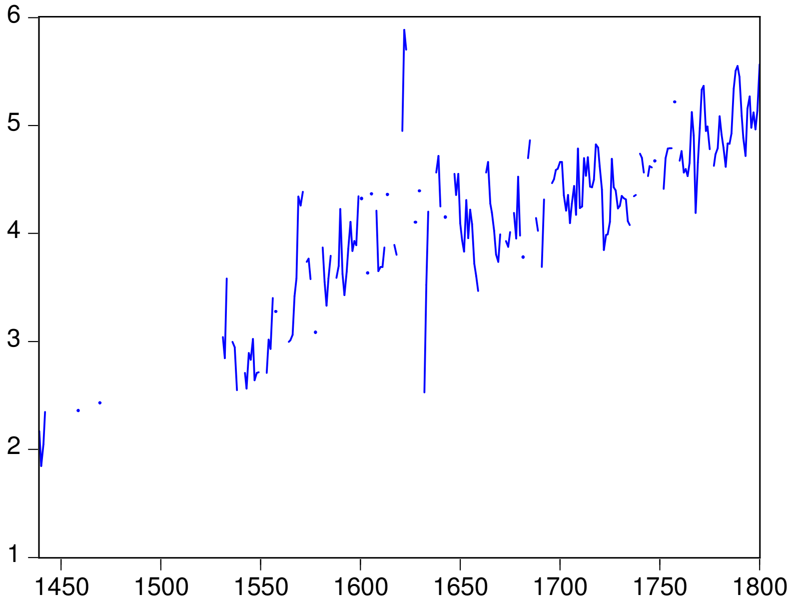

Figure 5
Conditional standard deviation of the residuals, Wheat prices Paris 1549-1697

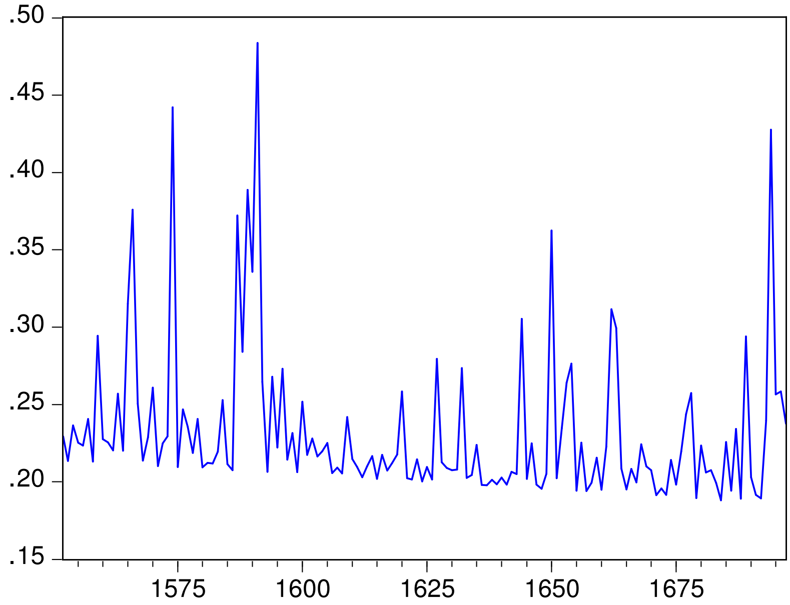


Figure 6
Conditional standard deviation of the residuals, wheat prices in Pisa 1550-1716

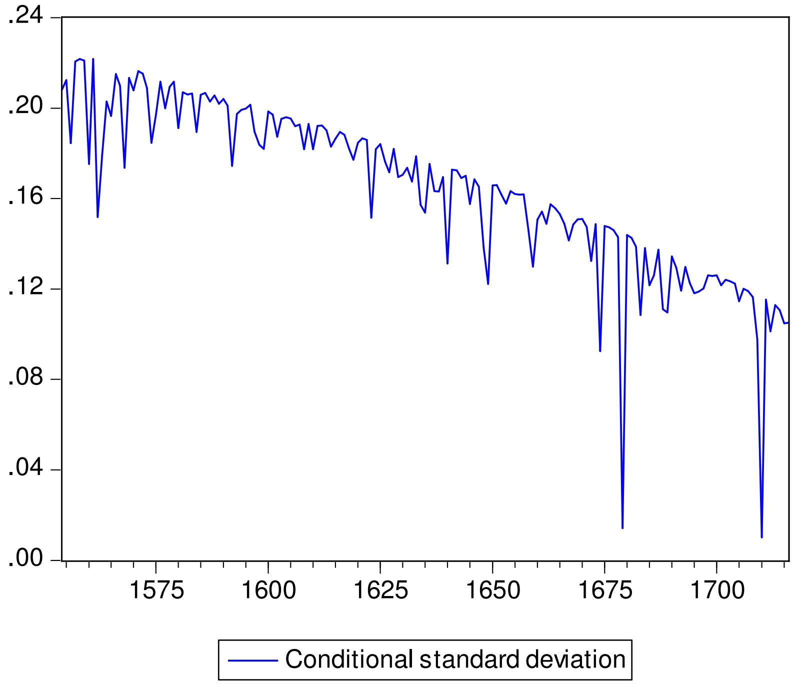

Figure 7
Conditional standard deviation of the residuals, rice prices in Hiroshima 1620-1857

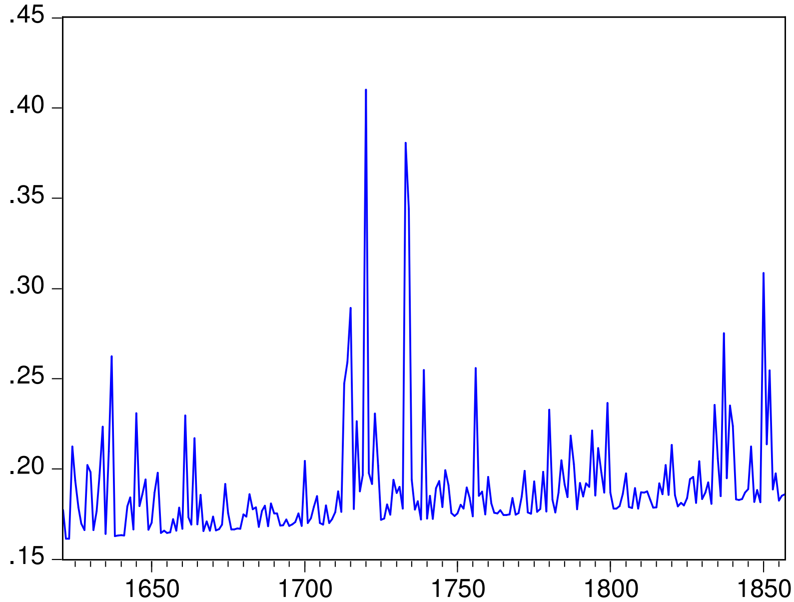


Figure 8
Conditional standard deviation of the residual, Vienna wheat prices 1439-1800

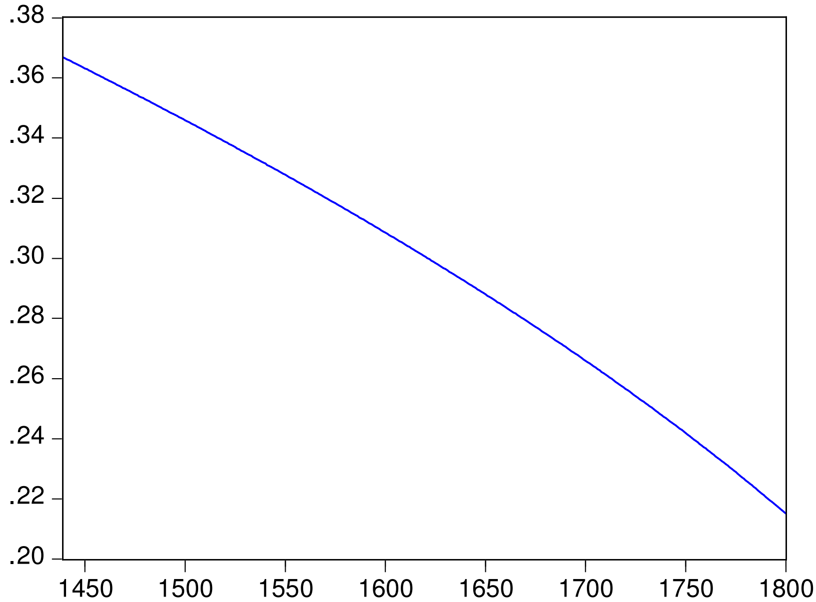

Figure 9

Conditional standard deviation of the residuals, wheat prices in Paris 1548M09-1698M08

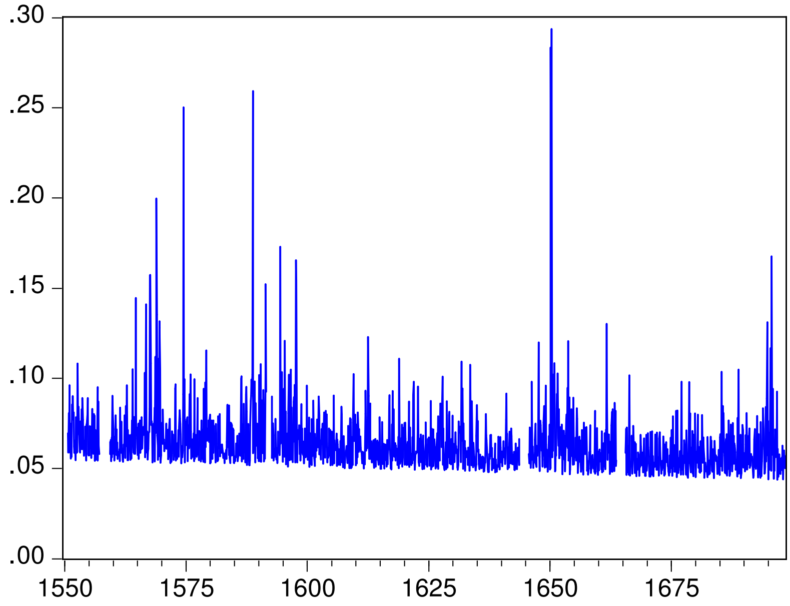


Figure 10a

Conditional standard deviation of the residuals, wheat prices in Pisa 1548M10-1631M07

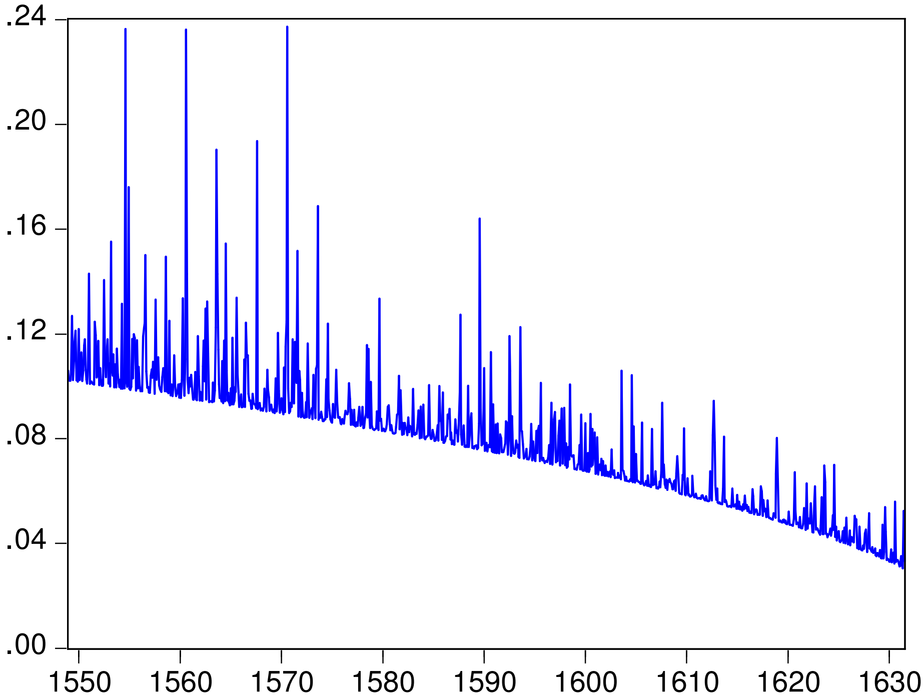

Figure 10b

Conditional standard deviation of the residuals, wheat prices in Pisa 1631M09-1818M07