

Do Banks Improve Financial Market Integration?

Liam Brunt¹ & Edmund Cannon²

17 January 2002

Abstract³

Using a large panel of weekly wheat prices, we infer the annual rate of return on capital in each county in England and Wales in the period 1770-1820. Throughout this period markets were efficient in the sense that weekly returns were serially uncorrelated. We show that the interest rate differential between London and each county can be explained by the density of bank coverage in that county. The explosion in provincial banking in England and Wales during the industrial revolution significantly reduced regional differentials in interest rates. This is direct evidence that financial intermediation determines the degree of market integration.

Keywords

Banks, Financial Integration, Industrial Revolution

JEL Classification

O16, N13, G21

¹ St. John's College, Oxford OX1 3JP. United Kingdom. liam.brunt@st-johns.ox.ac.uk

² University of Bristol, Department of Economics, 8 Woodland Road, Bristol BS8 1TN. United Kingdom. edmund.cannon@bristol.ac.uk

³ The authors gratefully acknowledge funding by ESRC Research Grant No. R000220371 and a Royal Economic Society Small Grant to help buy some additional software. The authors would also like to thank Ludivine Jeandupeux for research assistance; Rob Brewer, Anna Chernova, Becca Fell, Saranna Fordyce, Dave Lyne, Olivia Milburn, Hannah Shaw, Derick Shore, Liz Washbrook and Alun Williams for helping with obtaining, entering or checking the data set; Colin Knowles for computing assistance; the Bristol City Library for allowing us to borrow copies of the London Gazette not normally available for loan; and Cliff Attfield, Giam Pietro Cipriani, Suro Sahay, Lucy White for other assistance and helpful comments.

Introduction. Britain was distinguished from other economies during the Industrial Revolution by the sophistication of her financial markets (Neal, 1990) and this is often invoked as an important cause of her industrial success in the eighteenth and nineteenth centuries (Cameron, 1967). The importance of efficient capital markets - and the difficulty of achieving them - has also received considerable attention from development economists in recent years (Ikhide, 1996; Sial & Carter, 1996; Rosenzweig & Wolpin, 1993). And there has also been widespread discussion of their role in generating economic growth (MacKinnon, 1973; King & Levine, 1993).

The growth of commercial banking in Britain was truly spectacular, rising from 120 banks in 1784 to 660 banks in 1814 (Pressnell, 1956). If banking does indeed have a positive effect on the real economy then we should certainly be able to observe it in action in the British Industrial Revolution. There is, however, very little direct evidence on financial market integration during the Industrial Revolution. Since the link between the cost of capital and the rate of investment during the Industrial Revolution has never been established empirically, it is difficult to estimate the effect of financial integration on British economic growth.

The purpose of this paper is to quantify the effect of commercial banking on financial market integration in Britain during the Industrial Revolution. The fundamental problem in analysing financial markets in this period is that we have very little information on the cost of capital. Hence it is difficult to estimate the effect of any changes that might take place on the demand or supply side of the market, or any change in the nature of financial intermediation. The only recent analysis by Buchinsky and Polak (1993) was based on two series of deeds on property transactions, one for Yorkshire and the other for Middlesex, from which they inferred regional building, presumed to be determined partly by the cost of capital. Whilst their results are suggestive, one would not want to place too much weight on them alone.

Our departure point is to estimate the rate of return on capital in each county. We can do this by looking at the appreciation of a real asset through the year. The only asset which is sufficiently well documented through the period is grain. Bearing in mind that agriculture was the largest sector in the British economy until 1840 - and grain was the single most important output - the rate of return on holding grain is probably the single best indicator of the cost of capital in the economy. McCloskey & Nash (1984) and Taub (1987) show how the seasonal variation in grain prices is

related to the interest rate. Grain is an asset, and in equilibrium the holders of grain must be compensated for storage and interest costs. We use the appreciation of grain prices through the harvest year to estimate the rate of interest prevailing in each county from 1770 onwards. We have previously used this technique successfully on data from earlier periods (Brunt & Cannon, 1999). Our estimates are based on a large panel of weekly grain prices collected by the British Government between 1770 and 1820; this enables us to estimate year-specific county-specific rates of return on capital. We compare the movement of our interest rate series over time to that of the Consol rate, and find a positive relationship.

We then explain the geographical and temporal variation in interest rates with reference to the spread of banks outside London (the so-called ‘country banks’). At this time the Bank of England enjoyed great privileges in the commercial banking market. As a result, other banks were restricted to being partnerships (as opposed to joint stock companies) with a maximum of six partners all of whom had unlimited liability. Hence the size of individual banks was greatly circumscribed and the geographical reach of each bank was inevitably very limited. We take advantage of this fact by using the number of banks as a measure of the availability of banking services within each geographical area. There is county-level data available for country banks from 1800 onwards. We show that the density of country banks has a significant effect on narrowing the differential between the rate of return on Consols traded in London and the rate of return on grain traded in each county outside London.

The remainder of the paper is organised as follows. Section 1 discusses the data in more detail and illustrates the seasonal pattern of prices more closely. Section 2 begins with a discussion of market efficiency in the sense used in the finance literature - namely that returns should be serially uncorrelated and prices follow a martingale process (weak market efficiency). This is important because it influences our interpretation of the price data. We then estimate the gross rate of return on grain for each county-year. Section 3 uses these estimates to analyse the level of financial market integration at this time and how this was influenced by banks. Section 4 concludes.

1. Discussion of the data. Our analysis is based on weekly price data for wheat collected by the Receiver of Corn Returns after 1770, and published in the *London*

Gazette for each county in England and two regions in Wales. For a full discussion of the data set that we used, see Brunt & Cannon (2001).

Weekly price data is available for the 40 English counties, London, North Wales and South Wales - resulting in 43 price series for each grain. For simplicity, we refer to these 43 geographical areas as “counties”. We analyse each of these series from 10 November 1770 (when the data begin) to 30 September 1820 (with the exception of London, which ends on 8 July 1793). The Welsh data published in the *London Gazette* is disaggregated further after 10 November 1790, at which point the data is reported for each of the twelve Welsh counties individually. We have simply averaged the six Northern and six Southern Welsh counties respectively to extend the two series through to 1820. Unfortunately there is no price data for London for the period 1795-1820 and therefore we have no rates of return on wheat for the period for which we have data on banks.

Data for the period after 30 September 1820 is also available, but there are several changes in the list of towns for which data is collected in the period 1820-1823, which might result in potential structural breaks in any extended series. More importantly, some of the inland counties are poorly represented until a further change in the data set in 1828. For this reason 1820 forms a natural break: analysis of data after this period will be considered in later research.

The county prices are weighted averages of prices in several market towns in each county (where the volume of trade was used as weights). Details of how the data was constructed can be found in Brunt & Cannon (2001). Suffice it to say here that the number of towns monitored in each English county was almost always three. London was obviously a singleton, and the tiny county of Rutland had only two monitored towns; some of the counties have more towns (for example, Norfolk has twelve). By law, the average price for a county could be reported in the *London Gazette* if and only if returns had been received from at least two thirds of the towns in that county.

Some of the price series have missing observations, but the scale of this problem is minimal. For the years 1770 to 1820 we do not know why the prices are missing. For the period after 1820 we know that most missing price observations occur because there was no trade in that particular product in that particular week (so there was simply no market price). Over the entire period, the number of missing observations for wheat is very small, and is zero for some counties. The worst

offender is Herefordshire where there are 62 missing observations out of 2604 (so data is available for 97.6% of the observations).

Since we do not have quantity data to correspond to these price data, we cannot draw strong conclusions about the regional pattern of trade or production, but the pattern of missing price observations suggests that wheat was widely grown in all areas, even if the climate were unsuitable. This is what we would expect, given the high transport cost of moving wheat.

Figure 1 illustrates the time series price of wheat over the period 1770-1820. The solid middle line represents the mean average price of England and Wales: the two thinner lines represent the maximum and minimum price of the 43 counties in each week. This graph illustrates that prices in different regions moved closely together. This fact alone does not allow us to infer immediately that markets were highly integrated. It would be quite possible for the random shocks affecting prices to be highly correlated across counties (notably the weather), resulting in similar prices in markets that were only very weakly linked.

A further feature of the data is the tendency for the underlying levels and year-on-year variability of prices to be constant until about 1794. After this time there is both a secular increase in prices and much greater variability, due to the problems of the Napoleonic Wars. Ideally we should approach this problem by deflating the price series to obtain the real prices of the grains, but the lack of a suitable price deflator makes this impossible. Our analysis of rates of return will thus be confined to nominal rates of return, but since we are comparing the rates of return with nominal returns on consols, this will not influence our results.

We now turn to the seasonal pattern of the grain prices. The simplest way to describe the seasonal pattern is to conduct a regression of the form

$$(1) \quad \ln P_t = a_0 + \sum_{i=1}^{52} a_i + e_t$$

where a_0 is a constant and the a_i are 52 dummies for each week of the year with the constraint imposed that

$$(2) \quad \sum_{i=1}^{52} a_i = 0.$$

Regression (1) can be applied to each of the 43 county series for each grain. The mean values of the a are shown in Figure 2. Thus it appears that there is a strong seasonal pattern: the precise pattern varies between counties (particularly in amplitude), but is basically the same. However, care must be exercised in interpreting the graph, since in most counties the regression is insignificant using conventional standard errors. Chambers and Bailey (1995) argue on this basis that there is no seasonal effect at all (although inconsistently they maintain that there is a significant price fall in September). However, the use of conventional tests in this context is inappropriate: from Figure 1, it can be seen that the most of the temporal variation in prices is not the seasonal effect within years but the combination of the year-on-year effect throughout the sample and the large inflationary component after 1794: any attempt to conduct tests on the basis of equation (1) would need to account for the immense autocorrelation and heteroskedasticity present in the resulting residuals e_t . For this reason we only present Figure 2 as illustrative.

The graph reveals that the average seasonal pattern of prices is broadly consistent with the McCloskey-Nash characterisation: prices rise for most of the year and then fall at the point of the harvest. The gross rate of returns for the period when prices are rising is about 6 per cent, but this is for a period of about 30 weeks: expressed at an annual rate the returns would be 10.5 per cent.

In one aspect, however, the seasonal patterns for wheat depart from the McCloskey-Nash pattern: the period of the year when prices are falling is much longer. In the simple McCloskey-Nash hypothesis, price falls are only possible if stocks of grain are relatively low, or if they are unanticipated. Neither of these cases can be relevant on a consistent basis for the period of three-four months immediately after the harvest.

It is necessary to have a more sophisticated description of the pattern of production to understand pricing behaviour in this part of the year. Part of the reason for the pattern is that the timing of the harvest was itself highly variable and thus new stocks of grain could become available at any time between late July and early September. More importantly, grain had to be threshed before it was brought to market. Labour costs were highly seasonal, with the highest wages in the period July-September. This is because labour was needed for the urgent tasks of harvesting and then ploughing before the weather deteriorated in the autumn. After this period, the demand for labour fell considerably and only then did farmers allow much of their

labour force to be diverted by threshing, most of which took place in November and December. This remained true into the late nineteenth century (Young, 1770; Morton, 1868). Indeed, the structure of the Poor Laws meant that the opportunity cost of labour to farmers fell even more than the wage during this period. If farmers laid off workers in the winter then they had to subsidise them through the local Poor Rate, so the true cost of employing labourers in the winter was actually less than the market wage (Boyer, 1990). A further complication is that the grain had to be threshed in a particular order, since the straw was used partly as animal feed (Young, 1770). For this reason the falls in the price of wheat during this period do not fully reflect price rises in the underlying asset and we prefer to concentrate on the period from late December, by which time threshing was usually completed.

This section has provided a brief description of the wheat price data. We have shown that the seasonal pattern of wheat broadly follows the pattern suggested by McCloskey-Nash and hence we shall be able to use this data to estimate the rate of return on capital.

2. Econometric Estimates of the Gross Rate of Return. The simplest way to estimate rates of return is to calculate the average change in log-prices using the regression

$$(3) \quad \Delta \ln P_t = \mathbf{a} + u_t.$$

In fact the estimate of \mathbf{a} from this regression is just the difference between the first and the last price divided by the number of weeks. However, additional information can be obtained from intervening prices to determine the consistency of price behaviour with the theory. For this reason we start with a more general regression, namely,

$$(4) \quad \ln P_t = \mathbf{a} + \mathbf{x}_1 \ln P_{t-1} + \mathbf{x}_2 \ln P_{t-2} + \mathbf{x}_3 \ln P_{t-3} + u_t,$$

For the McCloskey-Nash hypothesis to make sense we should expect these prices to follow a random walk and also require some form of weak market efficiency, represented by the conditions that $\xi_1 = 1$, $\xi_2 = \xi_3 = 0$ and u_t be serially uncorrelated.

The estimate of α would then be the rate of return. In financial time series it is common to model the variance of the residual to be evolving over time (in the simplest case using just an ARCH specification), but our analysis of prices within a year, with at most 40 observations, precludes this approach. Confining ourselves to OLS means that our estimates may be inefficient and conventional standard errors unreliable. In these circumstances, we use Heteroskedastic Consistent Standard Errors and test explicitly for evidence of ARCH in the residuals (although with so few observations the power of this test will be weak). In fact we find little evidence of heteroskedasticity within years.

A more important problem is that although we are allowing the rate of return to vary between years, we are not considering the possibility that it varies within year: correspondingly we are assuming that agents are able to estimate the different values of α for each year, which we may refer to as α_y , where the y subscript denotes the year to which we are referring (reserving the t subscript for the week of the year). This makes it less straightforward to interpret the tests for $\xi_2 = \xi_3 = 0$ under the strict assumption of weak market efficiency. Weak market efficiency assumes that returns are uncorrelated given information on past prices alone, i.e., not including α_y . If this parameter is not known, then $\Delta \ln P_{t-1}$ provides useful information about α_y and we should not expect returns to be uncorrelated. For this reason our tests of market efficiency are not, strictly speaking, tests of weak-market efficiency, since they condition upon α_y .

We start our tests by using Augmented Dickey-Fuller (ADF) tests using the re-parameterised and extended equation

$$(5) \quad \Delta \ln P_t = \mathbf{a} + bt + (\mathbf{r} - 1) \ln P_{t-1} + \mathbf{I}_1 \Delta \ln P_{t-1} + \mathbf{I}_2 \Delta \ln P_{t-2} + u_t :$$

under the alternative hypothesis that there is no unit root the data would have to follow a deterministic trend and hence omitting the bt term would bias the test towards failure to reject the null, even if the null were false. We conduct this test on data for each county for each harvest year from 24 December to 8 August, resulting in 50 tests of a unit root for each county (harvest years 1770-71 through to 1819-20). To check the robustness of these tests we varied the start and end dates of the period and also varied the number of lags in the ADF test.

Given the number of tests, we should expect to reject the null hypothesis about 5% of the time, so it is desirable to summarise the information from all of these tests. Since within each county the different years are independent, it is possible to combine the ADF tests using a technique suggested by Fisher (1932) and advocated by Maddala and Wu (1999) in the similar context of evaluating test statistics within panel data. If the independent probability values of the ADF tests are denoted p_i , then the statistic

$$(6) \quad -2 \sum_{i=1}^N \ln p_i$$

has a χ^2 distribution with $2N$ degrees of freedom. Maddala & Wu's (1999) Monte Carlo tests show that the test statistic has almost exactly the right size when $N = 50$ and $T = 25$, which corresponds very closely to the test that we are conducting here, although the power is not particularly high (26%). We calculate these statistics for each of the 43 counties, calculating the probability values of the Dickey-Fuller statistic from a Monte Carlo experiment with 60,000 observations. Only one Fisher test is significant at the 5% level out of 43, so we cannot reject the null hypothesis of a unit root overall.

Proceeding on this basis, we imposed this restriction and estimated the further regressions

$$(7) \quad \Delta \ln P_t = \mathbf{a} + \mathbf{I}_1 \Delta \ln P_{t-1} + \mathbf{I}_2 \Delta \ln P_{t-2} + u_t$$

In the first case we can test for market efficiency by considering either the individual for $\lambda_1 = 0$ and $\lambda_2 = 0$ or the joint test $\lambda_1 = \lambda_2 = 0$: in all three cases we use the Heteroskedastic Consistent Moment Estimator of White using a small sample correction suggested by Davidson and McKinnon (1993) of multiplying the elements by $T/(T - 2)$. In fact tests for general heteroskedasticity using the White test, for ARCH and for autocorrelation suggested that the residuals were well behaved and F and t tests for market efficiency using conventional moment estimators gave qualitatively the same results in all but a few cases. Summaries of these results using the Fisher method outline above confirm that all of the regressions are well-specified.

Thus we have cannot reject the hypothesis of market efficiency once we have allowed for changes in the rate of return over time.

It would be possible to estimate the gross rate of return using $\alpha/(1 - \lambda_1 - \lambda_2)$ from equation (7). However, since we cannot reject the null hypothesis that $\lambda_1 = \lambda_2 = 0$, this method is unnecessary and will actually make the estimates less precise due to estimation error in λ_1 or λ_2 . For this reason we in fact we estimated the gross rates of return from the simple equation (3), which various specification tests suggested was a valid reduction of the more general equation.

3. Gross Rates of Return and Financial Markets. The previous section estimated annual series for the gross rates of return for each county. In this section we consider the relationship between these rates of return and rates of return on financial assets in London and the effect of financial institutions. As has been noted above, we cannot make a direct comparison of wheat rates of return in the counties with their counterpart in London because the London Gazette does not publish the London prices for the relevant period. To measure financial asset returns we use data on government 3% Consols, which is readily available from Neal (1990). If we look at frequencies of data greater than a year, there is no relationship between the price level of wheat, which depended mostly upon the harvest and hence the weather, and the price level of consols, which depended largely upon whether the UK was at war. We considered various time periods over which to obtain the holding period, partly to ensure that our results were not affected by the timing of coupon payments (made in early January and early July), but the results are very similar. The results reported here are for end-December to end-June and end-January to end-July.

Our country bank data is taken from the Shannon MSS. Unfortunately disaggregated data (for 41 counties) is available only from 1800, so our analysis will be confined to the period 1800 to 1820. An alternative source of data is the British Parliamentary Papers (1819), but data in this source are only available from 1808. However, we compared the two sources and found them to be very similar. The total number of banks grew considerably in this period, from 370 in 1800 to 656 in 1811: thereafter there was a modest fall and the number of banks fluctuated. We also know that the number of banks for a few years before 1800, but only on an aggregate basis: suffice it to say that there was considerable growth in country banking from 1784.

Unsurprisingly, the raw data are not very informative about the availability of financial institutions, since counties vary considerably in size: the county with the most banks is the biggest, Yorkshire, and the one with the least banks is the smallest, Rutland. Accordingly we scaled the data by alternative measures of the “size” of each county. Our two measures were the surface area and the annual population: the latter was obtained by linearly interpolating between the census years of 1801, 1811 and 1821. Obviously the area measure is constant over time, whereas the population is growing: the latter measure might be preferable since the demand for financial institutions is clearly related to the number of potential customers. In either case the resulting panel of data shows considerable differences in the time-series pattern within each county, which suggests that any relationship we find will not be a statistical artefact of the data trending etc. All of our estimation was using OLS, since there are no available instrument variables.

We now consider the relationship between rates of return in the counties, the London interest rate and the number of country banks. Assuming that banks should arbitrage differences in interest rates, we should expect the number of banks to reduce the differences between interest rates. The rates of return series are highly volatile over time, but the county series are all very similar to each other. In fact the standard deviation of these rates of return for wheat does not change over the period: since the national total of banks grew considerably from 1784 to 1820, this suggests that any arbitrage that was taking place was not equalising rates of return in different regions of the country with each other. However, there is still the possibility that country banks were able to arbitrage rates of return with financial markets in London. This possibility seems especially interesting since it is known that each country bank maintained links primarily with a correspondent bank in London (Bagehot, 1873), and can thus be interpreted as channelling sources of funds to or from the main financial centre.

To test the hypothesis that banks reduce the gap between interest rates in London and we considered the following panel regression

$$(9) \quad |(\mathbf{a}_{it} - \mathbf{d}) - h_t| = \mathbf{h}_i + \mathbf{q} B_{it} + \mathbf{g}t + v_{it},$$

where B_{it} is a measure of Banks/Size of county. The trend term is included to allow for the effects of economic growth, which might increase the demand for financial services.

Table 1 presents results from both within-groups and between-groups estimation of equation (9). The parameter estimates are considerably different, with the latter showing a much lower coefficient. Given the crudity of our measurement of county size, this is unsurprising: the banks measure is poorly measured in a cross-sectional sense. However, the individual time series for each county are probably well measured, since the variation in banks is relatively large over time, even compared with the average growth rate of the county economies. Thus the within groups estimator is much to be preferred.

The two main results from Table 1 are clear: the gap between the London consol rate and the rate of return on grain is negatively related to the extent of country banking, both across time and counties. The strongest effect is represented by the within group estimator, suggesting that the large increase in banking over this period played a substantial part in reducing differences between regional and London rates of return. The average measure of Banks/Area rose from 0.59 to 0.97 over the period, so an estimated coefficient of between 0.22 and 0.23 suggests that the overall effect of banks was to reduce the effect by about 8% points, which is much too high, even allowing for the countervailing effect from the trend variable. The Banks/Population variable rose on average from 0.045 to 0.059 and the average coefficient from regressions using slightly different measures of the consol return is 2.73, suggesting a fall of 3.8% in the differences between rates of return, which is very plausible.

4. Conclusion and Discussion. In this paper we have analysed county wheat prices in the period 1770-1820. The seasonal pattern of wheat prices is consistent with that suggested by McCloskey and Nash and hence we can attempt to estimate rates of return from that part of the year when prices are rising.

Our estimates of the county rates of return on wheat differ considerably from the holding return on 3% consols traded in financial markets over the period 1800-1820. However, the discrepancy between the regional wheat returns and the London financial returns is negatively correlated with the a measure of the number of county banks. As country banks reduced the difference between rates of return in London and rates of return in the provinces, we can infer that they were providing conduits for

excess funds to be invested in London and enabling areas where credit was short to benefit from the London market. This is strong evidence that county banks played an important rôle in developing financial integration in the industrial revolution.

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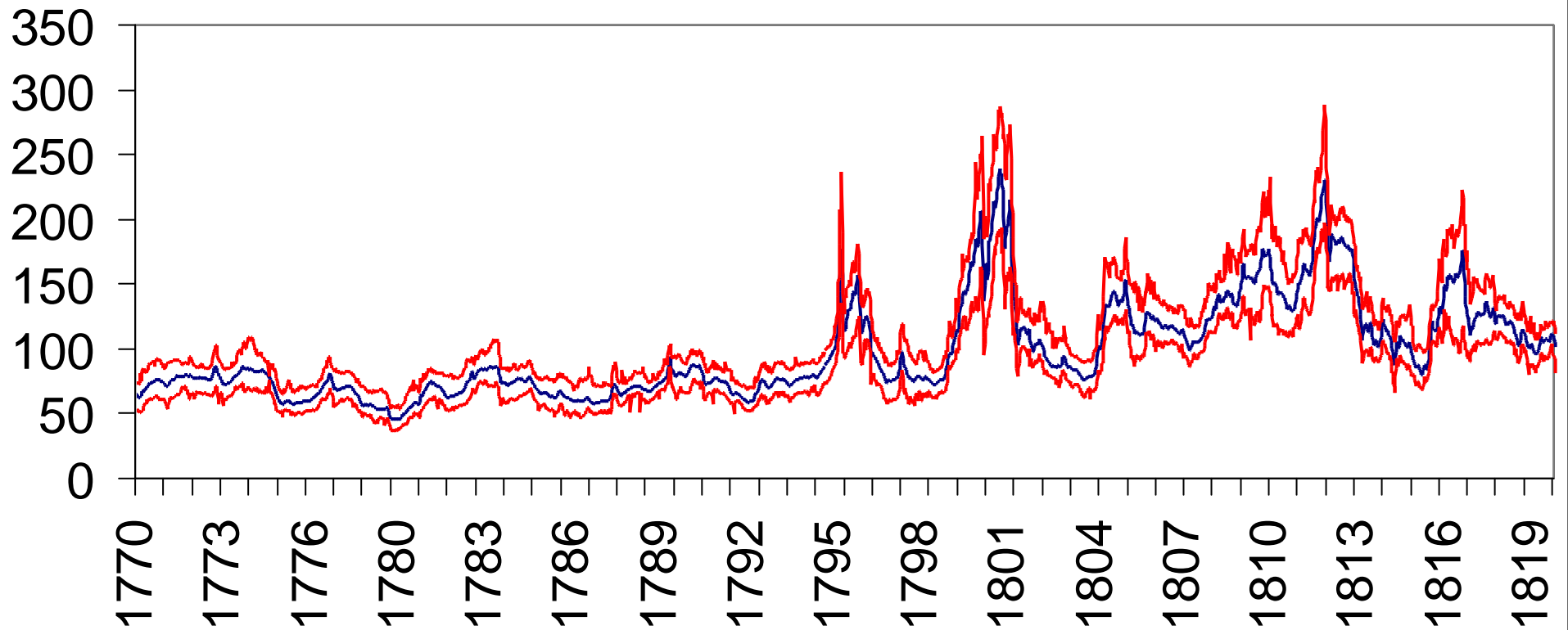
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Table 1

Regression of Absolute Difference between Wheat Returns and Consol Holding Return on Measures of Banks				
Consol holding return used				
Jan-July		*		*
Dec-June	*		*	
Within Groups Estimation				
Banks/Area			-0.22	-0.23
<i>hcse</i>			<i>0.035</i>	<i>0.042</i>
Banks/Pop	-2.630	-2.830		
<i>hcse</i>	<i>0.430</i>	<i>0.499</i>		
Trend	0.0036	0.003	0.006	0.0055
<i>hcse</i>	<i>0.0013</i>	<i>0.0016</i>	<i>0.0016</i>	<i>0.0018</i>
N	861	861	861	861
R-squared	0.055	0.042	0.056	0.042
σ	0.22	0.26	0.22	0.26
Between Groups Estimation				
Banks/Area			-0.021	-0.023
<i>hcse</i>			<i>0.0082</i>	<i>0.0087</i>
Banks/Pop	-0.306	-0.299		
<i>hcse</i>	<i>0.167</i>	<i>0.172</i>		
N	41	41	41	41
R-squared	0.044	0.036	0.042	0.11
σ	0.034	0.037	0.033	0.036

**Figure 1: Weekly Wheat Prices
Mean, Min and Max (pence per Qtr)**



**Figure 2 Wheat Mean County Seasonal Effect
Nov 1770 - Aug 1820**

