Are Statistical Reporting Agencies Getting It Right? Data Rationality and Business Cycle Asymmetry*

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Abstract

This paper provides new evidence on the rationality of early releases of industrial production (IP) and producer price index (PPI) data. Rather than following the usual practice of examining only first available and fully revised data, we examine the entire revision history for each variable. Thus, we are able to assess, for example, whether earlier releases of data are in any sense "less" rational than later releases, and when data become rational. Our findings suggest that seasonally unadjusted IP and PPI become rational after approximately 3-4 months, while seasonally adjusted versions of these series remain irrational for at least 6-12 months after initial release. For all variables examined, we find evidence that the remaining revision is predictable from its own past or from publicly available information in other economic and financial variables. Additionally, we find that there is a clear increase in the volatility of revisions during recessions, suggesting that early data releases are less reliable in tougher economic times. Finally, we explore whether nonlinearities in economic behavior manifest themselves in the form of nonlinearities in the rationality of early releases of economic data, by separately analyzing expansionary and recessionary economic phases and by allowing for structural breaks. These types of nonlinearities are shown to be prevalent, and in some cases lead to incorrect inferences concerning data rationality when they are not taken account of.

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1 Introduction

The construction of accurate preliminary announcements of macroeconomic variables remains an area of key interest to policymakers and researchers alike. The reasons for this are many. For example, policymakers have to rely upon preliminary estimates of key macroeconomic variables when making their decisions. Optimal policy is dependent on accurate assessments of the state of the economy, which implies that the policymakers are interested in whether early releases of data, when viewed as predictions of final or "true" data, may be "rational", using the terminology of Muth (1961). Similarly, researchers constructing empirical models for studying policy decisions are faced with the task of ensuring that the data used in their analysis correspond as closely as possible to those data policymakers actually had available in real-time. This issue is often ignored, as in most cases historical data are used as available at the time the research is undertaken. Only the most recent observations in these data are preliminary releases, corresponding with the data available to policymakers. More distant observations, though, are "final" releases, which possibly have undergone substantial revisions over time. Hence, the data used *ex post* by the modeler often are not the same as those used *ex ante* by the policymaker.

The above notions have led to a huge literature on examining the rationality of late predictions and early releases of macroeconomic variables, and the properties of the associated revision processes.[1] Three of the main features that tie the papers in this research area together are the following: First, many of them are concerned with either GDP or money data. Exceptions include Diebold and Rudebusch (1991) and Hamilton and Perez-Quiros (1996), who examine the predictive content of the composite leading index for output growth in real-time; Keane and Runkle (1990), who evaluate the rationality of price forecasts; and Kennedy (1993), who considers data on the index of industrial production. Second, the focus in many of these papers is on comparison of first available or "preliminary" data with fully revised or "final" data. One reason for this narrow focus is that data on the entire revision process for macroeconomic variables have been largely unavailable until recently. From the above list of references, only Amato and Swanson (2001), Bernanke and Boivin (2003), and Croushore and Stark (2001, 2003) consider complete revision histories for the variables that they examine. A further noteable exception to the failure to use revision histories is

[1] A partial list of the many publications in the area include: Morgenstern (1963), Stekler (1967), Howrey (1978), Zarnowitz (1978), Pierce (1981), Boschen and Grossman (1982), Mankiw, Runkle and Shapiro (1984), Mankiw and Shapiro (1986), Mork (1987), Milbourne and Smith (1989), Keane and Runkle (1989, 1990), Diebold and Rudebusch (1991), Neftçi and Theodossiou (1991), Kennedy (1993), Kavajecz and Collins (1995), Mariano and Tanizaki (1995), Rathjens and Robins (1995), Hamilton and Perez-Quiros (1996), Runkle (1998), Gallo and Marcellino (1999), Faust, Rogers and Wright (2003, 2004), Amato and Swanson (2001), Bernanke and Boivin (2003), Croushore and Stark (2001), and the references contained therein.

the seminal paper of Keane and Runkle (1990). In their paper, Keane and Runkle were careful to collect information on what releases of the GDP deflator were available at the dates when forecasts were made, and so revision information was actually in forecasters' information sets. Third, a common theme in these papers is that the rationality (or lack thereof) of preliminary data generally is assumed to be constant with respect to the business cycle and constant over time.

In this paper, we add to the literature on assessing the rationality of preliminary data by examining seasonally adjusted and unadjusted data for industrial production (IP) and the producer price index for finished goods (PPI). A number of features of our analysis differentiate our work from previous research. First, we have constructed monthly "real-time" data sets which include the entire revision history of the variables that we examine. This means that for each calendar date, we have a complete historical record of the actual values of each variable that were available at different release dates. Thus, we can inspect the entire revision process of the variables in detail, rather than just looking at the properties of first versus final releases of data, for example. One reason why this is useful is that we are now able to assess whether earlier releases are in any sense "less rational" than later releases. Put another way, we can measure how long it takes before the observed data become rational. In addition, we can include revision histories in the information sets used to examine the rationality of a particular release of data. This allows us to assess whether the remaining revision is predictable from its own past, that is whether revision histories can be used to construct "better" early releases of data.

Second, we recognize that business cycle asymmetry is a stylized characteristic of economic activity, and argue that there is no reason to preclude the possibility that nonlinearities in economic behavior manifest themselves in the form of nonlinearities in the revision process or in the rationality of early releases of macroeconomic data.[2] A number of papers recognize that nonlinearities may be present in the rationality of preliminary GDP data, including Brodsky and Newbold (1994) and Rathjens and Robins (1995), although they do not examine the entire revision process, and do not consider any explicit form of nonlinearity. Our approach is to directly test for the presence of nonlinearities in the revision process or in the rationality of early releases based on separate analysis of expansionary and recessionary economic episodes. The distinction between expansionary and recessionary episodes is useful because it allows us to determine the extent to which preliminary announcements are accurate in different phases of the business cycle. For example, a particular data release may be rational during expansions while it is irrational during recessions, or *vice versa.*

[2] See e.g. Burns and Mitchell (1946), Shapiro and Watson (1988), Diebold and Rudebusch (1996), King and Watson (1996), Ramsey and Rothman (1996), Baxter and King (1999), Stock and Watson (1999), Granger (2001) and the references contained therein, for discussions of business cycle asymmetry.

Third, there is a growing body of evidence showing that the statistical (business cycle) properties of US macroeconomic variables, output and inflation in particular, have changed during the post World War II period.[3] The explanations for these changes range from technological change, such as improvements in inventory management and information technology, to improved monetary policy. One of our goals in this paper is to investigate whether the revision processes of industrial production and inflation have also been subject to structural breaks. Put differently, we argue that changes in the rationality of early data releases that arise over time may be caused by changes in the data collection and processing techniques used by the statistical agencies. In summary, we wish to shed light on the question of whether government statistics can be made better, and to discuss the "closeness" of preliminary data to final data, an issue of relevance to agents and decision makers who use preliminary data.

We find that seasonally unadjusted IP and PPI releases become rational after approximately 3-4 months. Subsequent releases do not contain any new information. Seasonally adjusted IP and PPI data, on the other hand, remain irrational for at least 6-12 months. For most variables, the past of the revision process appears useful for ex ante prediction of the remaining revision, suggesting that rules might be constructed for the improvement of early data releases. Furthermore, we find evidence of both structural breaks and business cycle asymmetry in the revision process. One noteworthy feature of the revision process is that volatility of early data revisions increases during recessions, suggesting that early releases are less reliable in tougher economic times. Not surprisingly, this increase in revision volatility is associated with a general increase in the volatility of the growth rates of our series during recessions, and so is in part due to a general and overall increase in economic uncertainty during contractionary phases of the business cycle. The presence of structural change and nonlinearity in the revision process implies that failure to account for these features may lead to incorrect conclusions concerning data rationality based upon linear models. Indeed we find that rationality of early data releases frequently depends on the stage of the business cycle, and has changed over time.

The rest of the paper is organized as follows. Section 2 contains a summary of the methodology used, as well as a brief discussion of previous research. In Section 3, we introduce our real-time data sets and discuss the results of an exploratory data analysis describing the main features of the revision processes of our variables. Section 4 contains our main empirical findings, and conclusions are gathered in Section 5.

[3]See Watson (1994), Stock and Watson (1996), McConnell and Perez-Quiros (2000), Blanchard and Simon (2001), Chauvet and Potter (2001), and Sensier and van Dijk (2004), among others.

2 Testing Data Rationality: Methodology

In the sequel, the following notation is used. Let ${}_{t+k}X_t$ denote the value of the (annualized) monthly growth rate of a variable of interest which pertains to calendar date t as it is available at time $t+k$. In this setup, if we assume a one month reporting lag, then first release or "preliminary" data are denoted by ${}_{t+1}X_t$. In addition, we denote fully revised or "final" data, which is obtained as $k \to \infty$, by ${}_fX_t$.

Research in the area of testing rationality of preliminary announcements is based almost exclusively on the framework put forward by Mankiw and Shapiro (1986), linking the first and final releases of data. Their set-up aims to determine whether the first release ${}_{t+1}X_t$ is a noisy estimate of the fully revised data, or a rational forecast of ${}_fX_t$, or neither of the two. Note that in the first case, the revision is uncorrelated with the fully revised data, while in the second case it is uncorrelated with the first release data. Similarly, in case the preliminary announcement is equal to the final data plus measurement error, the variance of ${}_fX_t$ should be smaller than the variance of ${}_{t+1}X_t$, while the reverse should hold if ${}_{t+1}X_t$ is a rational forecast of ${}_fX_t$.

Assuming that the value of X measured at time t by the reporting agency is the value of X reported at time t, the errors-in-variables hypothesis can be tested by means of the regression model:

$$ {}_{t+1}X_t = \alpha + {}_fX_t\,\beta + \varepsilon_{t+1}, \tag{1} $$

where ε_{t+1} is an error term that is assumed to be uncorrelated with ${}_fX_t$. In particular, the null hypothesis that ${}_{t+1}X_t$ is equal to ${}_fX_t$ plus measurement error is given by $\alpha = 0$ and $\beta = 1$.

Using Muth's (1961) notion of rational expectations, the preliminary release ${}_{t+1}X_t$ is a rational forecast of the final data ${}_fX_t$ if and only if

$$ {}_{t+1}X_t = \mathsf{E}[{}_fX_t|\Omega_{t+1}], \tag{2} $$

where Ω_{t+1} the information set available at time $t+1$. This possibility can be examined by a second regression model, which takes the form:

$$ {}_fX_t = \alpha + {}_{t+1}X_t\,\beta + W'_{t+1}\gamma + \varepsilon_{t+1}, \tag{3} $$

where W_{t+1} is an $m \times 1$ vector of variables representing the conditioning information set available at time period $t+1$ and ε_{t+1} is an error term assumed to be uncorrelated with ${}_{t+1}X_t$ and W_{t+1}. The null hypothesis of interest in this model is that $\alpha = 0$, $\beta = 1$, and $\gamma = 0$, based on the notion of testing for rationality of ${}_{t+1}X_t$ for ${}_fX_t$ by finding out whether the conditioning information in W_{t+1}, available in real-time to the data issuing agency, could have been used to construct better conditional predictions of final data. Notice that this hypothesis, if rejected, is consistent with the errors-in-variables hypothesis.[4] Following Keane and Runkle (1990), the test of rationality of ${}_{t+1}X_t$ in the context of model (3) can be broken down into two sub-hypotheses, namely (i) unbiasedness and (ii) efficiency. The hypothesis of unbiasedness can be tested by imposing the restriction that $\gamma = 0$ and testing $\alpha = 0$, $\beta = 1$, while efficiency requires that $\alpha = 0$, $\beta = 1$, and $\gamma = 0$.

Based on an examination of preliminary and final money stock data, Mankiw *et al.* (1984) fail to reject the null hypothesis of unconditional unbiasedness in (1) and find evidence against the null that $\alpha = 0$, $\beta = 1$, and $\gamma = 0$ in (3), suggesting that preliminary money stock announcements are not rational and are an example of the classical errors-in-variables problem. In subsequent literature, attention has focused primarily on the second type of regression model, given as (3) above. For example, Kavajecz and Collins (1995) find that seasonally unadjusted money announcements are rational while adjusted ones are not. For GDP data, Mankiw and Shapiro (1986) find little evidence against the null hypothesis of rationality, while Mork (1987) and Rathjens and Robins (1995) find evidence of irrationality, particularly in the form of prediction bias (i.e. $\alpha \neq 0$ in (3)). Keane and Runkle (1990) examine the rationality of survey price forecasts rather than preliminary (or real-time) data, using the novel approach of constructing panels of real-time survey predictions. This allows them to avoid aggregation bias, for example, and may be one of the reasons why they find evidence supporting rationality, even though previous studies focusing on price forecasts had found evidence to the contrary.

One feature of our approach that differentiates it from previous research is that we have the entire revision history for each variable at our disposal, so that we can determine the "timing" of data rationality by generalizing (3) as follows:

$$ {}_fX_t - {}_{t+k}X_t = \alpha + {}_{t+k}X_t\,\beta + W'_{t+k}\gamma + \varepsilon_{t+k}, \tag{4} $$

where $k = 1, 2, \ldots$ defines the release of data (that is, for $k = 1$ we are looking at preliminary data, for $k = 2$ the data have been revised once, etc.). Notice that in (4), the null hypotheses of interest are now that $\alpha = \beta = 0$, assuming that $\gamma = 0$ (unbiasedness), and $\alpha = \beta = \gamma = 0$ (efficiency). Notice also that β in (4) has a different interpretation from β in (3), because ${}_{t+k}X_t$ is subtracted from both sides of (4) (such that the dependent variable ${}_fX_t - {}_{t+k}X_t$ represents the revision remaining after the k-th data release), and in (3), $k = 1$. Irrationality of preliminary data releases may arise simply because they are constructed using incomplete information sets.

[4]For further discussion on the relationship between errors-in-variables hypotheses and rationality hypotheses, the reader is referred to Croushore and Stark (2003) and Faust, Rogers, and Wright (2004), where the errors-in-variables and rational forecast models are associated with the notions of "noise" and "news", respectively.

For example, releases of aggregate industrial production are based on reported firm production levels. If, say, some firms are "late" in reporting, predictions of missing production levels may be used when constructing preliminary data releases, and these predictions may be inefficient. Over time, however, as the missing production data become available, newer releases may be expected to be "more" efficient. In this scenario, it follows that after some reasonable amount of time, all subsequent data releases are rational. Knowledge of the point in time after which releases of data are efficient has implications for policymakers, for example, particularly if they are interested in equating early data releases with efficient predictions of final data. Finally, notice that in (4) for $k > 1$, we may define W_{t+k} to include characteristics of the revision history, such as the revision between the first and k^{th} release ${}_{t+k}X_t - {}_{t+1}X_t$. In this way, we are able to examine whether inefficiency arises via information available in the revision history for a given release of data as well as through other sources.[5]

Obviously, inference based on fitting linear regression models of the form given by (4) may be affected by the presence of some form of nonlinearity. In the context of macroeconomic variables, two important types of nonlinearity that also may influence the revision process and rationality of early releases are business cycle asymmetry and structural change. In the remainder of this section, we describe how we have investigated the relevance of these nonlinearities.

2.1 Data Rationality and the Business Cycle

Our real-time data sets are useful for examining a number of business cycle features of macroeconomic data for which little is known, including asymmetry in the properties of the revision process, in data release rationality and in the time needed before early releases to become efficient.

Asymmetry in the revision process or in data release rationality may arise, for example, if the population of firms changes over the business cycle, due to the creation and destruction of firms during expansions and recessions. If early releases of aggregate production levels are based on the same sample of firms irrespective of the stage of the business cycle, this sample does not accurately represent the underlying population of firms, as young and newly-created firms are likely to be under-represented during expansions and over-represented during recessions. This may lead to biased early estimates of aggregate production levels in both recessions and expansions, where the sign and magnitude of the bias can be different, implying asymmetry in the revision process and/or rationality of preliminary releases.[6]

[5] A generalization of (4) is given by ${}_{t+l}X_t - {}_{t+k}X_t = \alpha + {}_{t+k}X_t\,\beta + W'_{t+k}\gamma + \varepsilon_{t+k}$, where $k < l$. By fitting models of this form, we may examine the rationality of a particular release of data relative to later releases of data. In the sequel, however, we focus on the model given in (4).

[6] Business cycle asymmetry in data release efficiency may also arise if government reporting agencies are conserva-

Our approach to this issue is to test for asymmetric unbiasedness and efficiency by fitting models of the form:

$$
\begin{aligned}
{}_f X_t - {}_{t+k}X_t = {} & \left(\alpha_1 + {}_{t+k}X_t\,\beta_1 + W'_{t+k}\gamma_1\right) \mathtt{I}[s_t = 0] \\
& + \left(\alpha_2 + {}_{t+k}X_t\,\beta_2 + W'_{t+k}\gamma_2\right) \mathtt{I}[s_t = 1] + \varepsilon_{t+k}, \quad (5)
\end{aligned}
$$

where $s_t = 0\,(1)$ if calendar month t is part of an expansion (recession), which is defined using the NBER-dated business cycle peaks and troughs, and where $\mathtt{I}[\cdot]$ is an indicator variable, taking the value 1 if its argument is true and 0 otherwise. Results based on this approach, however, should be viewed only as a rough initial guide to assessing the importance of asymmetry in our data, as the recession indicator variable is not in agents' information sets until (usually) after, or close to, the end of the recession. Tests for this type of nonlinearity are all based on checking the equality of coefficients in the above regression model. For example, consider the case where we are only interested in testing unbiasedness in expansions and recessions, so that $\gamma_1 = \gamma_2 = 0$ is assumed to hold. Upon rejecting the hypothesis of linear unbiasedness ($\alpha = \beta = 0$ in (4) with $\gamma = 0$ imposed), we test for asymmetry in the (un)biasedness properties by testing the null hypothesis $\alpha_1 = \alpha_2$ and $\beta_1 = \beta_2$ in (5). In cases where we find such asymmetry, we re-run all of our rationality tests by splitting the data into recessionary and expansionary phases. This allows us to ascertain whether absence of rationality in the entire sample is due primarily to a lack thereof during recessionary periods, for example.

2.2 Data Rationality and Structural Change

Structural changes in the revision process or data rationality may be caused, for example, by improvements in data collection and processing methods used by statistical reporting agencies during our sample period (see e.g. Rathjens and Robins (1995) for further discussion). To explore this possibility, we check for structural changes in the unbiasedness and efficiency test regressions. In particular, we use the sup-Wald test as developed by Andrews (1993):

$$
\text{SupW} = \sup_{\tau_1 \leq \tau \leq \tau_2} W_T(\tau), \quad (6)
$$

tive during expansionary periods (e.g. they tend to under-report economic growth estimates so as not to "over-heat" expectations and hence growth), and are liberal during contractionary periods, thereby leading to self-fulfilling cycles of economic decline (see e.g. Chauvet and Guo (2003), among others). This would lead to differing levels of efficiency for different observations in the same release of data, depending on whether they pertain to calendar months during expansionary or contractionary periods. The validity of this argument may be questioned given the independence of most statistical offices, however.

where $W_T(\tau)$ denotes a Wald statistic of the hypothesis of constancy of the parameters α, β (and γ) in (4) against the alternative of a one-time change at a fixed break date τ, given by

$$
\begin{aligned}
{}_fX_t - {}_{t+k}X_t &= \left(\alpha_1 + {}_{t+k}X_t\,\beta_1 + W'_{t+k}\gamma_1\right) \mathtt{I}[t < \tau] \\
&\quad + \left(\alpha_2 + {}_{t+k}X_t\,\beta_2 + W'_{t+k}\gamma_2\right) \mathtt{I}[t \geq \tau] + \varepsilon_{t+k}. \quad (7)
\end{aligned}
$$

The structural change tests are computed by imposing 15% symmetric trimming (i.e. we set $\tau_1 = [\pi T]$ and $\tau_2 = [(1-\pi)T] + 1$, with $\pi = 0.15$, where $[\cdot]$ denotes integer part and T is the sample size). The value of τ that minimizes the sum of squared residuals corresponding to (7) is taken to be the estimate of the break date, denoted as τ_B. We use the method of Hansen (1997) to obtain approximate asymptotic p-values for the sup-Wald test.[7] Given appropriate estimates of possible break dates, we also construct unbiasedness and efficiency tests on pre- and post-break samples, in order to assess whether our findings are driven by non-robustness of standard efficiency tests to structural change.

Estimation of all models in the sequel is carried out by least squares, with reported test statistics all based on heteroskedasticity and autocorrelation consistent standard error estimators.

3 Real-Time Data: Overview and Statistical Properties

We have collected seasonally adjusted (SA) and unadjusted (NSA) real-time monthly data for US industrial production (IP) and the producer price index for finished goods (PPI).[8] Although all data are available in levels, we examine only (annualized) monthly growth rates in this paper. This allows us to ignore issues relating to unit roots and cointegration, and to avoid the problem of accounting for pure base year changes when comparing multiple revisions of data for a particular calendar date.[9,10] In addition, the use of growth rates allows for comparison of our findings with those of previous studies.[11]

[7]Given that we find evidence for structural change in the revision process, we should in principle construct p-values for our unbiasedness and efficiency regressions using the methodology of Hansen (2000). However, in our case the distortions to relevant p-values are small, and so we report only the standard p-values.

[8]It should be stressed that SA data are constructed using seasonal factor that are generally changed only once each year. A feature of the data construction process which may in part account for some of the inefficiency in SA data that we will subsequently discuss.

[9]By a “pure base year change” we mean that data is revised only because of a base year change, without regular or definitional revisions occurring at the same time.

[10]As pointed out by an anonymous referee, it is important to note that monthly data are often much more noisy than their quarterly counterparts (e.g. IP versus GDP). Our results reflect this, as can be seen by comparing the results of various previous quareterly studies, including Keane and Runkle (1990) as well as the papers cited herein by Croushore et al., Chauvet et al., and Fuast et al., for example. Additionally, it is important to stress that earlier papers such as Kennedy (1993) already showed some inefficiencies in the data production process.

[11]The revision series examined in the sequel were all tested for a unit root using the augmented Dickey-Fuller test, and all series were found to be covariance stationary.

The number of release dates, or "vintages", for which we have historical real-time data available varies by series. In particular, for NSA IP, SA IP, and NSA PPI, the first vintage is 1963:1, and the last vintage is 2004:1, with historical data for each vintage going back to 1962:12. For SA PPI, the corresponding dates are 1978:2-2004:1 and 1978:1. To facilitate comparison of the results of NSA and SA PPI, we use the NSA data from the vintage of 1978:2 and calendar date 1978:1 onwards only. In the sequel we examine data for calendar periods up until 2001:12, while we use the vintage of 2004:1 as our "fully revised" data. Even though we can never claim to have a final record of historical data which is immune from potential future revision, we feel that the difference of 2 years between the last calendar date in our sample period and the date of this vintage is sufficient to consider all observations in this vintage as "fully revised". This is particularly true because we remove the effect of all benchmark revisions from our data prior to carrying out unbiasedness and efficiency tests, as discussed below.[12] Note that although benchmark revisions may include more than just "base year" and "weighting" changes, they are not generally forecastable, justifying their removal in our test regressions reported below.

The real-time industrial production data sets have been compiled from historical issues of the *Federal Reserve Bulletin* and the *Survey of Current Business*. Recent IP releases also are available on the Federal Reserve Board's web pages at `http://www.federalreserve.gov/releases/G17/`. In addition, a file containing the first five releases of seasonally adjusted IP from 1972:1 onwards is available on the same site, while the Federal Reserve Bank of Philadelphia recently made available a complete real-time data set on SA IP, see `http://www.phil.frb.org/econ/forecast/reaindex.html`. All of the data for PPI have been gathered from issues of the *Survey of Current Business, National Economic Trends*, and *Business Statistics*. Recent data are available on the web site of the Bureau of Labor Statistics at `http://stats.bls.gov/ppihome.html`.

A typical release of IP data consists of a first release for the previous month and revisions for the preceding one to five months (due to the availability of new source data and the revision of source data). In addition, from time to time more comprehensive re-benchmarking revisions and base-year changes occur, which affect the entire (or at least a large part of the) historical time series. During our sample period, base-year changes occurred in September 1971, July 1985, April 1990 and February 1997. Further, major revisions due to re-benchmarking occurred in July 1976, May 1993, December 1994, February 1997 (only for the seasonally adjusted series), and annually as of December 1997. See Kennedy (1993), Robertson and Tallman (1998) and Swanson, Ghysels

[12]For the NSA and SA PPI data, non-benchmark revisions occur only during the first 7 and 19 releases, respectively. For NSA and SA IP, 8.1% and 14.6% of the observations is still subject to non-benchmark revisions after 24 months, but the absolute magnitude of these revisions is very small.

and Callan (1999) for additional discussion of the revision process of industrial production.

The real-time data sets for the producer price index involve more infrequent revision. In fact, most observations on seasonally unadjusted PPI are revised only once, three months after their initial release. The same applies to seasonally adjusted PPI, although for these data additional "periodic" revisions occur at approximately 12 month intervals (usually February of each year). These periodic revisions involve incorporating "more comprehensive information" and usually affect data for the preceding 12-15 months. Non-benchmark revisions do not occur anymore after the first 7 and 19 releases for the NSA and SA PPI data, respectively. Finally, there has been no benchmark revision for seasonally unadjusted PPI since 1988, and the base-year was changed only in February 1971 (from 1957-9 to 1967) and February 1988 (to 1982).

A rough impression of the magnitude of the revisions in IP and PPI can be obtained from the plots given in Figures 1-4. In each figure, the first plot is of first available and final release data; the second plot shows the complete revision from preliminary to final release; the third plot is of benchmark revision; and the last plot is of non-benchmark revision.[13] While benchmark revisions often dominate non-benchmark revisions, both types of revision are rather large relative to the actual values of the series shown in the first plot. The statistical properties of the revision process are analysed in more detail below.

Tables 1-4 report a variety of summary statistics for each variable. These summary statistics include full-sample means of different transformations of the real-time data (see columns with the header "μ"), and means of sub-samples determined by: (i) application of structural change tests similar to the one discussed in Section 2.2 above (see columns with the header "μ_1 and μ_2" under "Structural Change"); and (ii) partitioning the data into those pertaining to calendar months in expansionary phases and recessionary phases of the business cycle as defined by NBER turning points (see the columns under the heading "B.C. Asymmetry"). The lower panel of each table contains similar results for volatilities (denoted σ, σ_1, and σ_2). Statistics are reported for fully revised data

[13]The following procedure is used to back out benchmark revisions from the data: the revisions occurring in vintages which are known to involve a comprehensive benchmark revision or base-year change are attributed completely to "benchmark revisions" and regular revisions in those vintages are set equal to zero. Of note is that we also used the approach is Kean and Runkle (1990), where benchmark revision was accommodated by using as "final data" those vintages available immediately before benchmark revisions. Interestingly, our findings remain completely unchanged when this approach is adopted, suggesting that the two approaches are in some sense interchangeable. Furthermore, when benchmark revision is ignored, test rejections increase markedly, as do regression $\overline{R}^2$ values, in accord with the findings of Keane and Runkle (1990) that failure to account for benchmark revisions can strongly affect results. Finally, it should be noted that other benchmark dates were also used, in order to assess the robustness of our findings. These include the set of dates provided by an anonymous referee as well as dates provided in the documentation accompanying the Federal Reserve Bank of Philadelphia real-time IP data. Our results were found to be surprisingly robust to the use of these alternative dating strategies, although the use of fewer dates generally resulted in findings of more inefficiency.

($_fX_t$), first available data ($_{t+1}X_t$), the complete revision ($_fX_t - {}_{t+1}X_t$), and the components of the complete revision due to "benchmark revisions" (base-year changes and other major revisions) and non-benchmark or regular revisions. In addition, statistics are reported for: (i) "fixed-width revisions" (i.e. $_{t+k+1}X_t - {}_{t+k}X_t$); (ii) "increasing-width revisions" (i.e. $_{t+k+1}X_t - {}_{t+1}X_t$); and (iii) "remaining revisions" (i.e. $_fX_t - {}_{t+k}X_t$). These last three types of data transformations are computed for regular revisions, which are defined to be the remaining revisions after removing benchmark revisions from the data, as detailed above. Note that "regular revisions" are of particular interest as these are used in our unbiasedness and efficiency regressions, as discussed below. A number of observations can be made based on these tables.

First, the fully revised (NSA and SA) IP growth rate is considerably higher than the preliminary announcement growth rate, on average, while for PPI they are very close. Hence, reporting agencies appear to be conservative when reporting the first release of IP. Note that for IP, the mean non-benchmark revision is about 3 times as large as the mean benchmark revision, and the latter is not significantly different from zero.

Second, the mean fixed-width, increasing width, and remaining revisions for industrial production are often significantly different from zero (as denoted by superscripts a, b, and c, referring to rejections of the null hypothesis that the mean revision is zero at the 1%, 5%, and 10% significance levels, respectively). As might be expected, there are fewer significant entries in the PPI tables. For example, for the NSA PPI, only the 3^{rd} and 4^{th} fixed-width revision means are significantly different from zero, which is due to the fact that most observations are revised only once, three months after initial publication.

Third, both first available and fully revised PPI data are characterized by a structural break in mean, which is dated in 1981 (see the first two rows of Tables 3 and 4). For both NSA and SA data, the post-break mean inflation rate is substantially lower than that for the pre-break period. For IP, evidence in favor of structural breaks in the mean is much weaker, with only the mean of seasonally adjusted fully revised IP data appearing to have possible changed, around 1970 (the p-value of the break test is 0.174). Interestingly, though, non-benchmark *revisions* for both NSA and SA IP data do exhibit evidence of a structural break (see the sup-Wald test rejection probabilities in the 4th column of entries in Tables 1 and 2). In particular, the mean non-benchmark revision is considerably smaller in the latter part of the sample (post 1976 for NSA data and post 1977 for SA data), suggesting that data collection and processing methods have become more efficient over time.

Fourth, with regard to business cycle asymmetry, notice that inflation is higher and industrial production growth is negative and larger in absolute magnitude during recessionary periods than during expansionary periods (see the last three columns of the tables). Thus, the stylized fact that recessions are shorter in duration, but greater in intensity is borne out in our data sets. For both NSA and SA IP, the hypothesis of equality of the mean of the complete non-benchmark revision during expansions and recessions is very close to being rejected. The non-benchmark revision for calendar months in expansionary periods is about 5 (3) times larger than for calendar months in recessionary periods for NSA (SA) data, while in addition, the mean non-benchmark revision for recessionary periods is not significant. Hence, it appears that preliminary IP data is correct on average during recessions, but not during expansions. Furthermore, the Fed is slow in adjusting the growth rate for expansionary periods, as the remaining non-benchmark revision is still significant after the 12^{th} release of data. Finally, note that during recessions the IP growth rate is adjusted downward initially, as on average the first fixed-width revision is negative. This implies that the second release of IP actually is further away from the final data than the first release. This is not the case during expansions.

Fifth, there is rather overwhelming evidence of structural breaks in the volatility of both first available and fully revised data, and in revisions.[14] In particular, for IP data the volatility of both benchmark and non-benchmark revisions has declined substantially over time, suggesting that preliminary announcements have become more precise, and providing further evidence that data collection and reporting methods have improved. Notice though that volatility of non-benchmark revisions in NSA PPI data has increased (slightly), suggesting the opposite.

Sixth, there is evidence in the IP series that there are business cycle asymmetries in the volatility, not only for first available and fully revised data but also for revisions. For example, the differential between volatilities in the complete non-benchmark revision of both NSA and SA IP data during expansionary and recessionary phases is approximately 25%, with volatility being larger during recessions. This finding suggests that uncertainty is different during different phases of the business cycle, and that this difference in uncertainty has an effect on the reliability of preliminary and early releases of IP data. Put another way, while the first release of data may appear to be more accurate on average during recessions, the volatility of revisions shows the opposite pattern.[15] Not surprisingly, this increase in revision volatility is associated with a general increase in the volatility of the growth rates of our series during recessions, and so is in part due to a general and overall

[14]The structural change in volatility of first release and fully revised SA IP is dated in 1984, in agreement with McConnell and Perez-Quiros (2000) and others, who report that the volatility of quarterly GDP has declined since around that time. Similarly, it is not unexpected that the change in volatility of PPI is dated in 1981, after the period of high inflation rates due to the OPEC oil crises in the 1970s.

[15]Recall that the complete non-benchmark revision in IP is substantially larger for calendar months in expansionary periods than for calendar months in recessionary periods.

increase in economic uncertainty during contractionary phases of the business cycle.[16]

Finally, upon inspecting the correlations between fully revised and first available data, we find that seasonally unadjusted first available data are much more highly correlated with their fully revised counterparts than the corresponding seasonally adjusted data.[17] Thus, the seasonal adjustment process itself, which is highly nonlinear (see e.g. Ghysels, Granger and Siklos (1996) for discussion of nonlinear aspects of seasonal adjustment filters currently used by statistical reporting agencies) seems to weaken the linkage between first available and final data. Furthermore, regardless as to whether the data have been seasonally adjusted, the correlations of both first available and fully revised data with the revisions themselves are often far from zero and are both positive and negative (correlations in excess of 25% are not uncommon, for example).

Overall, the main conclusion from this exploratory data analysis that carries through to the rest of our analysis is that there is ample evidence of both structural changes and business cycle asymmetries in the revisions to IP and PPI data. This suggests that these features may need to be accounted for when testing for unbiasedness and efficiency.

4 Testing Data Rationality: Empirical Findings

In this section, results based on regression models of the form given in (4), (5), and (7) are discussed. In these regression models, W_{t+k} includes the revision between the first and k^{th} release of data $({}_{t+k}X_t - {}_{t+1}X_t)$, the 3-month Treasury bill rate, the spread between yields on 10-year Treasury bonds and 3-month T-bills, the spread between Baa and Aaa rated corporate bonds, the first difference of logged crude oil prices (West Texas Intermediate Crude), and the dividend re-invested return on the S&P500. These variables are similar to those used by previous authors (see above), where more detailed motivation for their use can be found. All conditioning variables, except of course ${}_{t+k}X_t - {}_{t+1}X_t$, are measured at the end of month $t+k-1$.[18] Additionally, and because IP revisions generally occur for the previous three observations, we include fixed width revisions for

[16]It is worth noting that non-benchmark revision volatility in IP is larger during recessions until the 2^{nd} or 3^{rd} data release. For later releases, this situation is reversed, and there is more uncertainty regarding the remaining revision during expansions.

[17]For IP, the correlations between first available and fully revised data are approximately 0.90 and 0.75 for NSA and SA data, respectively. For PPI, the corresponding correlations are approximately 0.95 and 0.90.

[18]We also constructed efficiency tests with W_{t+k} defined to contain all variables measured at the end of calendar month t, regardless of the value of k. These types of tests allowed us to determine the length of time needed before all useful information available at the time of first release is incorporated into the revised data. In addition, we tried setting $W_{t+k} = {}_{t+k}X_t - {}_{t+1}X_t$, (i.e. only including the revision between the first and k^{th} release of data, in order to focus on the forecastability of the revision process from its own past). These alternative efficiency tests led to qualitatively similar conclusions to those reported below. Detailed results are therefore not shown here, but are available on request from the authors.

these most recent observations pertaining to months $t+k-2$, $t+k-3$, and $t+k-4$ in the set of control variables. As mentioned before, for the unbiasedness tests, we always impose $\gamma = 0$ or $\gamma_1 = \gamma_2 = 0$ in the appropriate regression models. In the efficiency regressions, we also include a set of centered seasonal dummies.[19]

4.1 Linear Models

The basic test of unbiasedness involves testing the null hypothesis that $\alpha = \beta = 0$ in (4), while imposing the restriction that $\gamma = 0$. Probability values for Wald tests of this null are given in the third last column of entries in Tables 5 and 6. Based on a rejection probability value of 0.10 (which is used in all subsequent discussions), for NSA IP we see that there is bias in the 1^{st} through 3^{rd} releases of data, and none thereafter. Thus, reporting agencies seem to "get it right", on average, after the first three revisions. The bias in SA IP (which mainly is due to the intercept α being non-zero, cf. Mork (1987) and Rathjens and Robbins (1995)) persists much longer (i.e. approximately 12 months). One reason for this finding may be the very nature of the seasonal adjustment process. In particular, seasonal adjustment procedures make use of two-sided moving average filters, with one side using historical data and one side using as yet to be determined future data. If the filters place enough weight on data that are not known with certainty for a full year or more, this could account for the increase in bias. In summary, while it is known that preliminary data are often biased, we now have evidence that the bias remains prevalent for multiple months of new releases, and for a year or more with SA data. This suggests that if one's objective is to use timely unbiased data, unadjusted data is preferable (see Kavajecz and Collins (1995) for an extensive discussion of this topic). Even more interesting, note that unadjusted PPI is essentially unbiased across all releases, except the 4^{th} release (for the reasons explained above - that is, revision usually occurs only 3 months after initial release). However, seasonally adjusted PPI is biased at all releases, up to 12 months (notice that here the coefficient β also is significantly different from 0, in contrast to the results for SA IP). Thus, even a full year of revisions is not sufficient to ensure that SA PPI

[19] In particular, we include $\sum_{s=1}^{11} \delta_s D^*_{s,t}$, where $D^*_{s,t} = D_{s,t} - D_{12,t}$, with $D_{s,t} = 1$ if time period t corresponds to month s and $D_{s,t} = 0$ otherwise. Note that the coefficient δ_s measures the difference between the intercept in month s and the average intercept, α. The seasonal effect for December can be computed as $\delta_{12} = -\sum_{s=1}^{11} \delta_s$, and hence, by construction, it holds that $\sum_{s=1}^{12} \delta_s = 0$. As a measure of the importance of seasonal effects, we report $\delta^* \equiv \sqrt{\sum_{s=1}^{12} \hat{\delta}_s^2}$ in the tables. Including seasonal dummies in the unbiasedness regressions does not yield qualitatively different results from those reported here. Tabulated results are available upon request from the authors.

releases are unbiased.[20,21]

The structural change tests reported in the second last column of the tables suggest that there is a structural break in the revision process for SA PPI in 1981, regardless of data release. On the other hand, there is no evidence of structural breaks in the adjusted IP data, and for unadjusted IP and PPI data all evidence of structural breaks is in early data releases.

Our tests for business cycle asymmetry reported in the last column of the tables provide moderate evidence of such nonlinearity for unadjusted IP, strong evidence for adjusted IP, moderate evidence for NSA PPI, and no evidence for SA PPI.[22]

In Tables 7 and 8, efficiency test results are contained in the 8^{th} column (tests of the hypothesis that $\gamma = 0$ in (4)) and the 10^{th} column (tests of the hypothesis that $\alpha = \beta = \gamma = \delta = 0$).[23] Given that we already have results on the unbiasedness of our data, we now focus on the joint hypothesis of unbiasedness and efficiency (i.e. $\alpha = \beta = \gamma = \delta = 0$). For this hypothesis, early releases of unadjusted IP become efficient after 3 months, while efficiency is realized for adjusted IP data after 3-4 months. Recall, though, that when only biasedness is tested for, the SA data remain biased even after data have been revised 12 times. This is true, even though *no further* irrationality is found to be due to missing information after 3-4 months.[24]

NSA price data also become efficient after 3-4 data releases, while adjusted price data are only efficient after 6 months. Unreported alternative efficiency tests which only include the quantity ${}_{t+k}X_t - {}_{t+1}X_t$ in W_{t+k} lead to the same results, suggesting that the remainder of the revision process can be forecast from its own past. Additionally, for all series, inefficiency remains prevalent

[20]Of note is that many of the unbiasedness regression models have serially correlated and conditionally heteroskedastic errors, according to Breusch-Godfrey serial correlation and autoregressive conditional heteroskedasticity (ARCH) tests, which are reported in the working paper version (see Swanson and van Dijk (2003)). This suggests that regression coefficients may be biased and that the regression models may be misspecified, a problem which persists even when W_{t+k} is included in the regression model for the purpose of testing efficiency of early releases.

[21]It is our opinion that the seasonal adjustment procedure may be so highly nonlinear that a zero mean forecast error for NSA data is essentially transformed into a non-zero mean forecast error for SA data, although further research must be carried out in this area before concrete statements along these lines can be made.

[22]Interestingly, there is no evidence of business cycle asymmetry in the NSA PPI regressions when $k = 1$ (i.e. based on the use of preliminary data in the unbiasedness regressions). Rather, business cycle asymmetry becomes more apparent *after* the preliminary data have been revised once (which from our above discussion we know happens after an interval of approximately 4 months).

[23]When testing for seasonality alone (see the 12^{th} column in the linear efficiency tables) *revisions* from SA data appear to exhibit seasonality.

[24]The reason for this finding may be that W_{t+k} enters into the regression models linearly, while the seasonal adjustment filter applied to the unadjusted data is highly nonlinear. In addition, note that the finding that it takes approximately 3 months before unadjusted IP data are not only unbiased but also efficient suggests that another sort of rationality test could be performed, by checking how many releases of IP data have an impact on returns in the stock market, say. If more than 3 releases have an impact, then that would suggest that agents are irrational, in the sense that they need not have used additional releases of IP when forming their expectations, as earlier releases were already fully rational. An assessment of rationality based on this argument is left for future research.

for a longer period of time when the information set used to check for efficiency includes additional regressors. Clearly, then, the revision processes of our price and IP data share some common features, even though the first available and fully revised PPI data conform to the errors-in-variables model outlined in Mankiw *et al.* (1984), while the IP data do not, as discussed above.

Notice also from Tables 7 and 8 that there is substantial evidence of both structural breaks and asymmetric business cycle effects in all series. For seasonally unadjusted IP, these features appear to be prevalent only for the first few releases, but for the remaining variables, test rejections also occur for later releases. This suggests that it may be of interest to re-fit our bias and efficiency models with imposed structural breaks and business cycle asymmetry. This is done in the remaining tables.[25] In particular, whenever there is a value of k, say k^*, for which structural change or business cycle asymmetry is present in the linear unbiasedness and/or efficiency regressions reported in Tables 5-8, all regressions with $k \leq k^*$ are re-estimated. This allows us to ascertain whether any of our linear unbiasedness and efficiency findings are dependent on the fact that nonlinearities present in the data are not accounted for in linear regression models.

4.2 Structural Change Models

In Table 9, notice that for $k = 1, 2$ (i.e. for those values of k for which there is evidence of structural breaks in the linear unbiasedness regressions), the data remain biased when separately testing unbiasedness of early NSA IP releases before and after the structural break, which for most releases is estimated to have occurred in the second half of the 1970s, or later. However, an interesting feature of the data arises when we examine the results for $k = 3, 6$, and 12.[26] In particular, there is a clear improvement in the quality of the data at later releases during the post-break periods. This finding stems from the observation that the unbiasedness null hypothesis is rejected pre-break, while the data are clearly unbiased post-break. Thus, unbiasedness findings for $k = 3, 6$, and 12 in the linear model are driven by *strong* unbiasedness in the post-break period. This sort of picture emerges for all of our series, and whenever either unbiasedness or efficiency regressions are considered, pointing to the dangers involved with simply fitting linear models without proper testing for nonlinearity. Additionally, this feature of the data is consistent

[25]Following the suggestion of the Associate Editor, we note that it is worth examining the sensitivity of our results to alternative business cycle dating methodologies. Along these lines, we also carried out our analysis using recessions defined as periods for which IP growth for a first three month period and a subsequent prior 3 month period (so that "quarterly" growth rather than monthly was used) was negative. In this case, all 6 months involved were assigned to "recession". Of note is that the results reported remain largely unchanged when this alternative business cycle dating methodology is applied.

[26]Even though no evidence of structural change was found for these values of k, we still re-estimated these regressions models with structural change dates picked using the methods discussed above in order to illustrate an interesting feature of the data.

with our earlier finding that early releases of data have become more accurate over time.

Interestingly, upon examination of both SA and NSA PPI data (see Table 10), many releases of data that are unbiased prior to the break date, are biased thereafter (the break dates for the series are in the early to mid 1980s). This is in agreement with the observation that the volatility of non-benchmark PPI revisions has increased over time.

A further direction of analysis in our context, which is left to future research, is the examination of the robustness of our findings to the presence of structural breaks and or temporal dependence in the volatility processes of our series. For further discussion of time dependent volatility in business cycle data, the reader is referred to Chauvet and Popli (2004).

For those releases of data where structural breaks were found, imposing these structural breaks does little to change the efficiency test results obtained with linear models. In addition, the same sort of asymmetry noted above for unbiasedness pre- and post-break also holds when efficiency regressions are re-run allowing for structural breaks. Detailed results are not shown here, but are available on request.

4.3 Business Cycle Asymmetry Models

Tables 11 and 12 contain unbiasedness test results for IP and PPI based on models with imposed business cycle asymmetry of the variety discussed in Section 2.1. For IP, there appears to be more bias during expansionary than during recessionary episodes (see the last two columns of each table, where probabilities that there is no bias are given, with the last column corresponding to recessionary periods and the second last corresponding to expansionary periods), especially for SA data. Note that in addition to prediction bias ($\alpha_1 \neq 0$) we also find that β_1 differs significantly from zero for all SA releases considered (but not for NSA releases). This corroborates our previous finding that the mean non-benchmark revision is significantly larger during expansionary periods. Also note that for NSA IP data, we find evidence against unbiasedness during recessionary periods for the 2^{nd} and 3^{rd} releases but not for the first. This is in agreement with our earlier finding that the first revision during recessionary months actually pushes the second release of NSA IP further away from the final data than the first release.

The results of nonlinear efficiency tests are broadly supportive of the results based on our tests based on the linear regression model, and these are therefore not shown in detail here. However, one observation worth noting is that for SA IP, we find that the data remain inefficient during recessions for 12 months, while they become efficient after at most 6 months during expansions.

5 Concluding Remarks

In this paper we have examined the entire revision process for IP and PPI data. This allowed us to construct tests of rationality not only for preliminary data or "first releases" of these important macroeconomic variables, but also for later releases of data. In addition, various features of the revision process itself, which hitherto have not been discussed, are examined, allowing us to address various questions about revision accuracy, volatility, and timing.

Our findings suggest that unadjusted IP and PPI data releases become rational after approximately 3 months. Seasonally adjusted data, on the other hand, remain irrational for at least 6-12 months. In addition, bias and inefficiency are usually removed from the data after around the same number of releases. We find evidence of predictability of the revision process, either from its own past or from other publicly available information, suggesting a possible route for improving the reporting of preliminary data. We further find evidence of both structural breaks and business cycle nonlinearities, and find that failure to account for these features of the data in some cases leads to incorrect conclusions concerning unbiasedness and efficiency. Finally, there is a clear increase in revision volatility during recessions, suggesting that early data are less reliable in tougher economic times, a finding consistent with our observation that early releases of data growth rates are also more volatile during recessions.

A number of issues remain for future research. For example, note that the explanatory power of the efficiency regressions is quite small in general. Hence it remains to assess, in real-time, whether the revision history of a variable (or information from other macroeconomic and financial variables) can be used to sharpen future preliminary releases of that variable. Faust, Rogers and Wright (2004) have already made important progress in this area by examining preliminary and final GDP data for the G-7 countries, and find some evidence that it can indeed be done, albeit not for US data. Additionally, it should prove of interest to ascertain whether the revision history of one economic variable is useful for predicting other variables, in real-time.

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Table 1: Structural Change and Business Cycle Asymmetry in Mean and Volatility: Real-Time Seasonally Unadjusted Industrial Production

		Structural Change				B.C. Asymmetry		
	μ	μ_1	μ_2	$\mu_1=\mu_2$	τ_B	μ_1	μ_2	$\mu_1=\mu_2$
Fully revised	3.156^a	7.034^a	2.380^a	0.174	1969.6	4.849^a	-7.345^a	0.000
First available	1.776^b	4.200^a	1.269	0.644	1969.9	3.526^a	-9.073^a	0.000
Complete revision	1.380^a	3.142^a	0.772^a	0.105	1972.12	1.323^a	1.728	0.788
Non-benchmark revision	1.041^a	2.107^a	0.472^b	0.003	1976.7	1.167^a	0.261	0.146
Benchmark revision	0.338	0.066	1.720^b	0.571	1995.7	0.156	1.467	0.334
Fixed width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+k}X_t$								
k=1	0.295^c	1.341^a	-0.008	0.010	1971.9	0.410^b	-0.421	0.153
k=2	0.460^a	0.637^a	0.113	0.166	1988.10	0.448^a	0.535^c	0.798
k=3	0.257^a	0.475^a	0.102^b	0.112	1979.2	0.265^a	0.203	0.776
Increasing width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+1}X_t$								
k=1	0.295^c	1.341^a	-0.008	0.010	1971.9	0.410^b	-0.421	0.153
k=2	0.755^a	1.767^a	0.462^b	0.050	1971.9	0.858^a	0.114	0.200
k=3	1.011^a	1.898^a	0.560^b	0.007	1976.2	1.124^a	0.317	0.182
k=6	0.991^a	1.977^a	0.489^b	0.006	1976.2	1.104^a	0.291	0.181
k=12	0.996^a	2.005^a	0.482^b	0.006	1976.2	1.110^a	0.291	0.178
Remaining (non-benchmark) revisions ${}_fX_t - {}_{t+k}X_t$								
k=1	1.041^a	2.107^a	0.472^b	0.003	1976.7	1.167^a	0.261	0.146
k=2	0.746^a	1.160^a	0.272^c	0.014	1983.10	0.757^a	0.682^c	0.854
k=3	0.286^a	0.671^a	0.011	0.008	1979.3	0.309^a	0.147	0.511
k=6	0.061^c	0.235^a	-0.037	0.133	1977.1	0.074^c	-0.019	0.173
k=12	0.061^c	0.151^a	-0.045	0.104	1984.1	0.076^b	-0.030	0.082
k=24	0.010	0.055^c	-0.042	0.590	1984.1	0.012	0.000	0.600

		Structural Change				B.C. Asymmetry		
	σ	σ_1	σ_2	$\sigma_1=\sigma_2$	τ_B	σ_1	σ_2	$\sigma_1=\sigma_2$
Fully revised	26.351	32.057	23.579	0.002	1975.9	25.804	29.742	0.150
First available	27.901	30.189	25.186	0.146	1984.2	27.056	33.137	0.047
Complete revision	11.081	12.351	8.165	0.000	1990.2	11.011	11.516	0.769
Non-benchmark revision	5.442	6.611	3.395	0.000	1987.10	5.264	6.549	0.103
Benchmark revision	10.079	10.994	7.867	0.001	1990.7	10.028	10.393	0.797
Fixed width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+k}X_t$								
k=1	3.276	4.436	2.717	0.000	1975.8	3.136	4.142	0.015
k=2	2.660	3.089	1.888	0.000	1988.1	2.536	3.433	0.023
k=3	1.414	0.867	1.516	0.004	1969.1	1.334	1.911	0.004
Increasing width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+1}X_t$								
k=1	3.276	4.436	2.717	0.000	1975.8	3.136	4.142	0.015
k=2	4.727	5.651	3.109	0.000	1987.10	4.542	5.876	0.048
k=3	5.100	6.082	3.379	0.000	1987.10	4.911	6.272	0.064
k=6	5.158	6.173	3.377	0.000	1987.10	4.969	6.331	0.080
k=12	5.240	6.297	3.388	0.000	1987.10	5.064	6.331	0.103
Remaining (non-benchmark) revisions ${}_fX_t - {}_{t+k}X_t$								
k=1	5.442	6.611	3.395	0.000	1987.10	5.264	6.549	0.103
k=2	3.625	4.342	2.368	0.000	1987.10	3.448	4.723	0.025
k=3	1.911	2.219	1.447	0.000	1986.5	1.868	2.175	0.224
k=6	0.767	1.203	0.130	0.000	1986.2	0.836	0.341	0.008
k=12	0.518	0.831	0.083	0.000	1985.8	0.544	0.355	0.287
k=24	0.160	0.277	0.013	0.001	1984.8	0.183	0.013	0.005

Notes: The table contains summary statistics for the mean and variance of real-time data on annualized monthly growth rates of seasonally unadjusted Industrial Production over the period 1963.1-2001.12, based on data vintages for 1963.1-2004.1. In the upper block, the column headed μ contains the unconditional mean, the columns headed μ_1 and μ_2 under "Structural Change" contain the means before and after the break-point τ_B, which is determined by maximizing the point-wise heteroskedasticity- and autocorrelation-consistent Wald test for testing the null hypothesis $\mu_1 = \mu_2$. The p-value corresponding to the null hypothesis that there was no structural break in the mean of the process is reported in the column headed $\mu_1 = \mu_2$. The columns headed μ_1 and μ_2 under "B.C. Asymmetry" contain the means during expansions and recessions, respectively, which are defined according to NBER business cycle turning points. The column headed $\mu_1 = \mu_2$ contains the p-value for the Wald test of equality of these two means. Entries marked with ^{a}, ^{b} and ^{c} are significantly different from zero at the 1, 5 and 10% level, respectively, using HAC standard errors. The lower block of the table contains similar statistics for the standard deviations of the time series (computed under the assumption of a constant mean).

Table 2: Structural Change and Business Cycle Asymmetry in Mean and Volatility: Real-Time Seasonally Adjusted Industrial Production

		Structural Change				B.C. Asymmetry		
	μ	μ_1	μ_2	$\mu_1=\mu_2$	τ_B	μ_1	μ_2	$\mu_1=\mu_2$
Fully revised	3.137^a	6.537^a	2.488^a	0.014	1969.3	4.909^a	-7.853^a	0.000
First available	2.015^a	4.477^a	1.515^b	0.197	1969.7	3.941^a	-9.924^a	0.000
Complete revision	1.122^a	2.213^a	0.737^b	0.154	1973.2	0.969^a	2.072^b	0.241
Non-benchmark revision	0.851^a	1.434^a	0.522^a	0.009	1977.10	0.927^a	0.380	0.242
Benchmark revision	0.270	0.033	1.296^a	0.180	1994.8	0.041	1.691^b	0.064
Fixed width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+k}X_t$								
k=1	0.165^c	0.239^a	-0.173	0.423	1994.11	0.219^b	-0.176	0.185
k=2	0.343^a	0.531^a	0.130	0.099	1983.9	0.332^a	0.411^b	0.676
k=3	0.275^a	0.327^a	0.015^a	0.450	1975.6	0.288^a	0.194	0.523
Increasing width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+1}X_t$								
k=1	0.165^c	0.239^a	-0.173	0.423	1994.11	0.219^b	-0.176	0.185
k=2	0.508^a	0.645^a	-0.119	0.052	1994.12	0.551^a	0.235	0.396
k=3	0.782^a	0.964^a	-0.047	0.008	1994.12	0.839^a	0.429	0.318
k=6	0.803^a	0.984^a	-0.022	0.011	1994.12	0.864^a	0.430	0.293
k=12	0.788^a	0.965^a	-0.022	0.014	1994.12	0.846^a	0.430	0.313
Remaining (non-benchmark) revisions ${}_fX_t - {}_{t+k}X_t$								
k=1	0.851^a	1.434^a	0.522^a	0.009	1977.10	0.927^a	0.380	0.242
k=2	0.687^a	1.238^a	0.406^a	0.042	1976.2	0.708^a	0.556^b	0.633
k=3	0.344^a	0.773^a	0.194^b	0.339	1973.1	0.376^a	0.145	0.289
k=6	0.084	0.267^b	-0.019	0.517	1977.1	0.102	-0.025	0.296
k=12	0.062	0.228^c	-0.033	0.232	1977.1	0.080	-0.049	0.294
k=24	0.001	0.100	-0.054	0.794	1977.1	-0.001	0.015	0.849

		Structural Change				B.C. Asymmetry		
	σ	σ_1	σ_2	$\sigma_1=\sigma_2$	τ_B	σ_1	σ_2	$\sigma_1=\sigma_2$
Fully revised	8.207	9.920	6.192	0.000	1984.1	7.163	14.681	0.000
First available	7.907	9.929	5.465	0.000	1984.4	6.599	16.019	0.000
Complete revision	5.532	6.179	4.105	0.000	1989.1	5.435	6.133	0.470
Non-benchmark revision	4.004	4.480	3.170	0.000	1987.10	3.888	4.722	0.125
Benchmark revision	4.991	5.860	4.019	0.000	1983.7	4.924	5.404	0.512
Fixed width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+k}X_t$								
k=1	2.029	1.572	2.159	0.032	1971.7	1.954	2.495	0.061
k=2	1.679	1.313	1.798	0.014	1972.7	1.611	2.103	0.032
k=3	1.013	0.930	1.336	0.002	1993.12	1.014	1.010	0.976
Increasing width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+1}X_t$								
k=1	2.029	1.572	2.159	0.032	1971.7	1.954	2.495	0.061
k=2	3.059	2.361	3.286	0.004	1972.7	2.927	3.879	0.045
k=3	3.389	2.742	3.633	0.016	1973.8	3.256	4.217	0.078
k=6	3.514	2.821	3.724	0.036	1972.1	3.395	4.250	0.135
k=12	3.740	4.117	3.080	0.003	1987.10	3.658	4.248	0.302
Remaining (non-benchmark) revisions ${}_fX_t - {}_{t+k}X_t$								
k=1	4.004	4.480	3.170	0.000	1987.10	3.888	4.722	0.125
k=2	2.885	3.290	2.217	0.000	1987.3	2.849	3.106	0.445
k=3	2.028	2.383	1.444	0.000	1987.3	2.095	1.607	0.108
k=6	1.291	1.795	0.461	0.000	1987.3	1.386	0.697	0.034
k=12	0.952	1.335	0.355	0.000	1986.9	0.996	0.678	0.310
k=24	0.433	0.843	0.189	0.000	1977.7	0.451	0.326	0.634

Notes: The table contains summary statistics for the mean and variance of real-time data on annualized monthly growth rates of seasonally adjusted Industrial Production over the period 1963.1-2001.12, based on data vintages for 1963.1-2004.1. See Table 1 for further details.

Table 3: Structural Change and B.C. Asymmetry in Mean and Volatility: Real-Time Seasonally Unadjusted Producer Price Index for Finished Goods

		Structural Change				B.C. Asymmetry		
	μ	μ_1	μ_2	$\mu_1=\mu_2$	τ_B	μ_1	μ_2	$\mu_1=\mu_2$
Fully revised	2.997^a	10.046^a	1.686^a	0.000	1981.10	2.915^a	3.553^b	0.739
First available	2.992^a	9.986^a	1.691^a	0.000	1981.10	2.906^a	3.559^b	0.724
Complete revision	0.005	−0.010	0.053	1.000	1996.4	0.009	−0.026	0.799
Non-benchmark revision	0.000	−0.036	0.024	1.000	1987.6	0.002	−0.010	0.922
Benchmark revision	0.011	0.121	−0.030	0.928	1980.9	0.019	−0.026	0.683
Fixed width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+k}X_t$								
k=1	−0.001	−0.005	0.000	0.992	1985.6	−0.002	0.000	0.311
k=3	-0.463^a	-1.517^a	-0.229^b	0.000	1982.5	-0.395^a	-0.907^a	0.073
k=4	0.464^a	1.533^a	0.233^b	0.000	1982.4	0.408^a	0.831^a	0.181
Increasing width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+1}X_t$								
k=1	−0.001	−0.005	0.000	0.992	1985.6	−0.002	0.000	0.311
k=3	-0.464^a	-1.517^a	-0.231^b	0.000	1982.5	-0.397^a	-0.907^a	0.074
k=4	0.000	−0.016	0.053	1.000	1996.4	0.012	−0.076	0.478
k=6	−0.011	−0.082	0.019	0.964	1985.1	−0.011	−0.010	0.998
k=12	0.000	−0.036	0.024	1.000	1987.6	0.002	−0.010	0.922
Remaining (non-benchmark) revisions ${}_fX_t - {}_{t+k}X_t$								
k=1	0.000	−0.036	0.024	1.000	1987.6	0.002	−0.010	0.922
k=2	0.002	−0.045	0.021	1.000	1985.1	0.003	−0.010	0.912
k=3	0.464^a	1.522^a	0.236^b	0.000	1982.4	0.398^a	0.897^a	0.131
k=4	−0.000	−0.017	0.007	1.000	1985.1	−0.010	0.066	0.230
k=6	0.011	0.073	0.000	0.684	1981.8	0.013	0.000	0.150

		Structural Change				B.C. Asymmetry		
	σ	σ_1	σ_2	$\sigma_1=\sigma_2$	τ_B	σ_1	σ_2	$\sigma_1=\sigma_2$
Fully revised	6.205	9.831	5.530	0.000	1981.10	5.829	8.668	0.016
First available	6.239	9.714	5.593	0.000	1981.10	5.913	8.376	0.046
Complete revision	1.189	1.301	1.098	0.523	1988.10	1.190	1.183	0.971
Non-benchmark revision	1.116	0.927	1.191	0.320	1984.11	1.133	1.001	0.478
Benchmark revision	0.688	0.933	0.522	0.001	1982.1	0.672	0.754	0.656
Fixed width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+k}X_t$								
k=1	0.004	0.007	0.002	0.724	1985.6	0.004	0.002	0.311
k=3	1.241	1.596	1.178	0.079	1981.8	1.224	1.348	0.375
k=4	1.246	1.587	1.184	0.091	1981.9	1.240	1.284	0.802
Increasing width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+1}X_t$								
k=1	0.004	0.007	0.002	0.724	1985.6	0.004	0.002	0.311
k=3	1.239	1.595	1.176	0.078	1981.8	1.222	1.348	0.369
k=4	1.110	0.929	1.184	0.346	1985.1	1.123	1.019	0.566
k=6	1.130	0.973	1.193	0.554	1984.11	1.150	1.001	0.426
k=12	1.116	0.927	1.191	0.320	1984.11	1.133	1.001	0.478
Remaining (non-benchmark) revisions ${}_fX_t - {}_{t+k}X_t$								
k=1	1.116	0.927	1.191	0.320	1984.11	1.133	1.001	0.478
k=2	1.118	0.927	1.194	0.307	1984.11	1.136	1.002	0.471
k=3	1.118	0.927	1.194	0.307	1984.11	1.136	1.002	0.471
k=4	1.273	1.754	1.186	0.004	1981.9	1.264	1.336	0.685
k=6	0.056	0.303	0.013	0.000	1981.8	0.063	0.013	0.031

Notes: The table contains summary statistics for the mean and variance of real-time data on annualized monthly growth rates of the seasonally unadjusted Producer Price Index for Finished Goods over the period 1978.2-2001.12, based on data vintages for 1978.2-2004.1. For completeness, we planned to include k=2 in all of the panels in the above table, and k=12 to increasing width and remaining revision panels. However, note that these values of k are still not reported for some types of revisions. The reason for this is that all revisions for these values of k are identically zero. See Table 1 for further details.

Table 4: Structural Change and B.C. Asymmetry in Mean and Volatility: Real-Time Seasonally Adjusted Producer Price Index for Finished Goods

		Structural Change				B.C. Asymmetry		
	μ	μ_1	μ_2	$\mu_1 = \mu_2$	τ_B	μ_1	μ_2	$\mu_1 = \mu_2$
Fully revised	3.012^a	10.268^a	1.733^a	0.000	1981.8	2.919^a	3.620^b	0.703
First available	3.103^a	10.806^a	1.746^a	0.000	1981.8	3.013^a	3.694^b	0.726
Complete revision	−0.091	−0.532	−0.007	0.719	1981.11	−0.094	−0.075	0.953
Non-benchmark revision	−0.014	−0.062	0.077	1.000	1993.9	0.010	−0.171	0.442
Benchmark revision	−0.091	−0.532	−0.007	0.719	1981.1	−0.104	0.097	0.318
Fixed width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+k}X_t$								
k=1	0.017	0.026	−0.015	0.725	1996.12	0.018	0.011	0.785
k=2	−0.021	-0.060^b	0.001	0.630	1986.11	−0.017	−0.048	0.405
k=3	-0.121^b	−0.049	-0.292^b	0.696	1994.11	-0.133^b	−0.037	0.409
k=4	0.117^b	0.030	0.320^b	0.499	1994.10	0.132^b	0.016	0.515
Increasing width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+1}X_t$								
k=1	0.017	0.026	−0.015	0.725	1996.12	0.018	0.011	0.785
k=2	−0.004	−0.037	0.041	0.674	1991.11	0.001	−0.037	0.400
k=3	-0.125^b	−0.050	-0.309^b	0.695	1995.1	-0.132^c	−0.074	0.620
k=4	−0.008	−0.061	0.078	0.812	1992.10	−0.000	−0.058	0.660
k=6	−0.043	-0.146^b	0.039	0.555	1988.8	−0.020	−0.191	0.308
k=12	0.001	0.097	−0.050	1.000	1986.5	0.016	−0.102	0.545
Remaining (non-benchmark) revisions ${}_fX_t - {}_{t+k}X_t$								
k=1	−0.014	−0.062	0.077	1.000	1993.9	0.010	−0.171	0.442
k=2	−0.031	−0.073	0.048	1.000	1993.8	−0.008	−0.183	0.451
k=3	−0.010	−0.040	0.069	1.000	1995.4	0.009	−0.135	0.522
k=4	0.111	0.022	0.312^b	0.654	1994.8	0.143	−0.097	0.351
k=6	0.021	0.062	−0.061	0.977	1994.1	0.026	−0.014	0.822
k=12	0.051	−0.036	0.109^c	0.504	1987.8	0.055	0.023	0.835

		Structural Change				B.C. Asymmetry		
	σ	σ_1	σ_2	$\sigma_1 = \sigma_2$	τ_B	σ_1	σ_2	$\sigma_1 = \sigma_2$
Fully revised	5.222	9.305	4.502	0.000	1981.8	4.913	7.244	0.107
First available	5.783	9.952	5.048	0.000	1981.8	5.464	7.873	0.076
Complete revision	2.265	3.493	2.043	0.000	1981.9	2.264	2.271	0.983
Non-benchmark revision	1.646	1.777	1.204	0.051	1996.6	1.638	1.693	0.852
Benchmark revision	1.721	2.987	1.498	0.000	1981.8	1.741	1.595	0.593
Fixed width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+k}X_t$								
k=1	0.110	0.128	0.022	0.566	1997.12	0.122	0.035	0.001
k=2	0.109	0.131	0.027	0.473	1996.11	0.112	0.084	0.507
k=3	0.594	0.339	1.172	0.000	1994.8	0.599	0.562	0.864
k=4	0.618	0.366	1.177	0.000	1994.7	0.609	0.675	0.780
Increasing width (non-benchmark) revisions ${}_{t+k+1}X_t - {}_{t+1}X_t$								
k=1	0.110	0.128	0.022	0.566	1997.12	0.122	0.035	0.001
k=2	0.179	0.219	0.024	0.154	1996.12	0.194	0.079	0.024
k=3	0.729	0.523	1.197	0.000	1994.8	0.745	0.623	0.572
k=4	0.791	0.531	1.368	0.000	1994.7	0.794	0.771	0.933
k=6	0.966	0.694	1.260	0.001	1990.6	0.968	0.957	0.970
k=12	1.482	1.207	1.677	0.058	1987.12	1.513	1.280	0.416
Remaining (non-benchmark) revisions ${}_fX_t - {}_{t+k}X_t$								
k=1	1.646	1.777	1.204	0.051	1996.6	1.638	1.693	0.852
k=2	1.634	1.750	1.244	0.115	1996.6	1.627	1.680	0.855
k=3	1.577	1.326	1.754	0.115	1987.12	1.570	1.620	0.859
k=4	1.552	1.672	1.284	0.206	1994.7	1.532	1.682	0.607
k=6	1.070	1.366	0.061	0.000	1996.7	1.070	1.071	0.999
k=12	0.575	0.733	0.094	0.000	1996.1	0.533	0.855	0.331

Notes: The table contains summary statistics for the mean and variance of real-time data on annualized monthly growth rates of the seasonally adjusted Producer Price Index for Finished Goods over the period 1978.2-2001.12, based on data vintages for 1978.2-2004.1. See Table 1 for further details.

Table 5: Unbiasedness of Real-Time Growth Rates of Industrial Production - Linear Model

Release k	α	β	$\overline{\mathrm{R}}^2$	$\alpha = 0$ $\beta = 0$	SC (τ_B)	BCA
			Seasonally Unadjusted			
1	1.028	0.007	−0.001	0.000	0.000	0.394
	(0.209)	(0.010)			(1976.2)	
2	0.720	0.013	0.007	0.000	0.032	0.591
	(0.152)	(0.007)			(1976.2)	
3	0.277	0.004	0.000	0.005	0.035	0.305
	(0.092)	(0.004)			(1979.3)	
4	0.035	−0.002	−0.001	0.675	0.115	0.377
	(0.049)	(0.003)			(1977.10)	
5	0.054	−0.003	0.001	0.362	0.171	0.548
	(0.043)	(0.003)			(1977.10)	
6	0.068	−0.003	0.001	0.210	0.260	0.226
	(0.041)	(0.003)			(1977.10)	
12	0.062	−0.001	−0.002	0.178	0.281	0.138
	(0.034)	(0.002)			(1984.1)	
24	0.010	0.000	−0.002	0.857	0.735	0.857
	(0.019)	(0.001)			(1984.1)	
			Seasonally Adjusted			
1	0.958	−0.053	0.011	0.000	0.208	0.005
	(0.182)	(0.024)			(1976.2)	
2	0.718	−0.014	−0.000	0.000	0.129	0.352
	(0.136)	(0.017)			(1976.2)	
3	0.363	−0.008	−0.001	0.002	0.468	0.012
	(0.102)	(0.014)			(1970.1)	
4	0.123	−0.019	0.006	0.167	0.358	0.025
	(0.076)	(0.012)			(1970.1)	
5	0.136	−0.020	0.007	0.099	0.346	0.033
	(0.073)	(0.011)			(1970.1)	
6	0.145	−0.022	0.010	0.072	0.340	0.016
	(0.073)	(0.012)			(1970.1)	
12	0.108	−0.016	0.007	0.110	0.361	0.025
	(0.062)	(0.009)			(1970.1)	
24	0.032	−0.011	0.004	0.323	0.260	0.161
	(0.045)	(0.007)			(1970.1)	

Notes: The table contains unbiasedness test results for different releases of annualized monthly growth rates of Industrial Production over the period 1963.1-2001.12, based on data vintages for 1963.1-2004.1, and based on estimating equation (4) with $\gamma = 0$ imposed. The column headed $\alpha = 0$ $\beta = 0$, contains the p-value of the Wald statistic for testing the indicated restriction. The column headed SC contains the p-value from the sup-Wald test for testing the hypothesis $\alpha_1 = \alpha_2$ and $\beta_1 = \beta_2$ in equation (7) with $\gamma_1 = \gamma_2 = 0$ imposed, where the change-point, τ_B, is given in parentheses. The column headed BCA contains the p-value from the Wald test for testing the hypothesis $\alpha_1 = \alpha_2$ and $\beta_1 = \beta_2$ in equation (5) with $\gamma_1 = \gamma_2 = 0$ imposed, using NBER-defined recessions and expansions. For all test statistics, heteroskedasticity and autocorrelation-consistent versions are used. Additionally, heteroskedasticity and autocorrelation-consistent standard errors are given in parentheses under coefficient estimates.

Table 6: Unbiasedness of Real-Time Growth Rates of the Producer Price Index for Finished Goods - Linear Model

Release k	α	β	$\overline{\mathrm{R}}^2$	$\alpha = 0$ $\beta = 0$	SC (τ_B)	BCA
			Seasonally Unadjusted			
1	0.040	-0.013	0.002	0.467	0.351	0.439
	(0.050)	(0.011)			(1986.2)	
2	0.042	-0.013	0.002	0.457	0.265	0.435
	(0.050)	(0.011)			(1986.2)	
3	0.042	-0.013	0.002	0.457	0.265	0.435
	(0.050)	(0.011)			(1986.2)	
4	0.430	0.013	0.001	0.001	0.000	0.313
	(0.122)	(0.016)			(1982.4)	
5	0.006	-0.002	-0.002	0.379	0.807	0.154
	(0.009)	(0.003)			(1981.8)	
6	0.010	0.000	-0.003	0.262	0.811	0.261
	(0.006)	(0.003)			(1981.8)	
			Seasonally Adjusted			
1	0.232	-0.079	0.077	0.000	0.023	0.662
	(0.097)	(0.020)			(1981.9)	
2	0.220	-0.080	0.081	0.000	0.000	0.642
	(0.093)	(0.020)			(1981.9)	
3	0.224	-0.076	0.074	0.000	0.001	0.713
	(0.091)	(0.019)			(1981.9)	
4	0.316	-0.069	0.065	0.000	0.001	0.696
	(0.098)	(0.020)			(1981.9)	
5	0.160	-0.054	0.052	0.004	0.000	0.883
	(0.083)	(0.016)			(1981.9)	
6	0.179	-0.052	0.049	0.003	0.001	0.761
	(0.078)	(0.015)			(1981.9)	
12	0.139	-0.029	0.032	0.014	0.041	0.895
	(0.059)	(0.010)			(1981.8)	

Notes: The table contains unbiasedness test results for different releases of annualized monthly growth rates of the Producer Price Index for Finished Goods over the period 1978.2-2001.12, based on data vintages for 1978.2-2004.1. Note that k=12 is not added to the top panel because revisions are in this case equal to zero for all observations. For the same reason, k=12 is not added to the results in Table 10 below. See Table 5 for further details.

Table 7: Efficiency of Real-Time Growth Rates of Industrial Production - Linear Model

						Tests of efficiency				Tests for structural change				Tests for B.C. asymmetry			
Release k	α	β	γ	δ^*	$\overline{\mathrm{R}}^2$	$\alpha=0$, $\beta=0$	$\gamma=0$	$\delta=0$	$\alpha=\beta=0$, $\gamma=\delta=0$	α,β	γ	δ	α,β, γ,δ	α,β	γ	δ	α,β, γ,δ
						Seasonally Unadjusted											
1	2.323	−0.032		5.952	0.035	0.006	0.106	0.004	0.000	0.052	0.021	0.369	0.000	0.221	0.000	0.000	0.000
	(0.788)	(0.026)											(1976.2)				
2	0.483	−0.009	0.251	3.374	0.061	0.488	0.001	0.003	0.000	0.810	0.000	0.108	0.000	0.432	0.025	0.004	0.000
	(0.494)	(0.014)	(0.053)										(1976.2)				
3	0.411	0.001	0.046	0.767	0.022	0.250	0.063	0.750	0.003	0.173	0.176	0.111	0.000	0.066	0.000	0.017	0.000
	(0.261)	(0.009)	(0.031)										(1982.3)				
4	0.337	−0.007	0.050	0.726	0.016	0.098	0.048	0.703	0.360	0.250	0.004	0.268	0.020	0.011	0.039	0.884	0.403
	(0.168)	(0.007)	(0.019)										(1973.4)				
5	0.135	−0.009	0.035	0.774	0.016	0.313	0.069	0.478	0.404	0.055	0.009	0.168	0.001	0.140	0.337	0.849	0.847
	(0.140)	(0.007)	(0.017)										(1976.2)				
6	0.172	−0.009	0.037	0.795	0.007	0.236	0.308	0.608	0.763	0.009	0.184	0.273	0.280	0.349	0.685	0.840	0.956
	(0.139)	(0.006)	(0.017)										(1976.2)				
12	0.026	−0.004	0.028	0.668	0.025	0.554	0.212	0.180	0.246	0.181	0.313	0.188	0.277	0.938	0.268	0.527	0.418
	(0.128)	(0.004)	(0.011)										(1980.4)				
24	0.013	−0.003	0.007	0.430	0.004	0.551	0.763	0.967	0.997	0.425	0.796	0.855	1.000	0.394	0.783	0.965	0.997
	(0.031)	(0.002)	(0.007)										(1984.1)				
						Seasonally Adjusted											
1	1.406	−0.030		3.086	0.064	0.089	0.002	0.007	0.000	0.000	0.000	0.000	0.000	0.499	0.000	0.001	0.000
	(0.666)	(0.029)											(1986.1)				
2	0.848	−0.029	0.337	2.470	0.100	0.020	0.000	0.000	0.000	0.041	0.000	0.000	0.000	0.177	0.115	0.001	0.000
	(0.360)	(0.016)	(0.059)										(1986.1)				
3	0.669	−0.015	0.070	1.217	0.015	0.041	0.021	0.085	0.000	0.994	0.030	0.327	0.000	0.000	0.331	0.001	0.000
	(0.288)	(0.013)	(0.054)										(1973.4)				
4	0.409	−0.021	0.027	0.978	−0.004	0.069	0.404	0.183	0.061	0.310	0.000	0.082	0.000	0.001	0.907	0.446	0.051
	(0.240)	(0.012)	(0.031)										(1973.3)				
5	0.478	−0.024	0.024	0.830	−0.007	0.026	0.412	0.449	0.143	0.055	0.012	0.115	0.000	0.023	0.968	0.480	0.417
	(0.253)	(0.011)	(0.028)										(1968.10)				
6	0.338	−0.024	0.015	0.721	−0.011	0.069	0.591	0.730	0.458	0.092	0.010	0.415	0.011	0.053	0.401	0.629	0.131
	(0.234)	(0.012)	(0.028)										(1968.10)				
12	0.269	−0.018	0.000	0.792	0.001	0.084	0.377	0.412	0.396	0.971	0.762	0.024	0.204	0.042	0.416	0.286	0.045
	(0.209)	(0.009)	(0.026)										(1984.12)				
24	0.085	−0.008	−0.019	0.533	−0.019	0.354	0.823	0.832	0.969	0.603	0.648	0.219	0.463	0.214	0.779	0.825	0.985
	(0.196)	(0.005)	(0.029)										(1968.10)				

Notes: The table contains efficiency test results for different releases of annualized monthly growth rates of Industrial Production over the period 1963.1-2001.12, based on data vintages for 1963.1-2004.1, and based on estimating equation (4) with W_{t+k} defined to include the increasing width revision up to the k^{th} release of data (i.e. ${}_{t+k}X_t - {}_{t+1}X_t$), the fixed width revisions for the three most recent observations pertaining to months $t+k-2$, $t+k-3$, and $t+k-4$, the 3-month T-bill rate, the spread between yields on 10-year Treasury bonds and 3-month T-bills, the spread between Baa and Aaa rated bond yields, the first difference of the logged oil price, and the return on the S&P 500, all observed at the end of month $t+k-1$. The column with header γ contains estimates of the coefficient associated with the regressor ${}_{t+k}X_t - {}_{t+1}X_t$. The column with header δ^* contains values of $\sqrt{\sum_{s=1}^{12}\hat{\delta}_s^2}$, where $\hat{\delta}_s$ is the estimated coefficient for $D_{s,t}$, $s=1,\ldots,11$, and $\hat{\delta}_{12} = -\sum_{s=1}^{11}\hat{\delta}_s$, which measures the magnitude of seasonal patterns in the revision process. The remainder of the columns contain statistics that correspond to those reported in the previous tables. See Table 5 for further details.

Table 8: Efficiency of Real-Time Growth Rates of the Producer Price Index for Finished Goods - Linear Model

						Tests of efficiency				Tests for structural change				Tests for B.C. asymmetry			
Release k	α	β	γ	δ^*	$\overline{\mathrm{R}}^2$	$\alpha=0$ $\beta=0$	$\gamma=0$	$\delta=0$	$\alpha=\beta=0$ $\gamma=\delta=0$	α, β	γ	δ	α, β γ, δ	α, β	γ	δ	α, β γ, δ
								Seasonally Unadjusted									
1	0.088	−0.027		0.850	−0.012	0.105	0.791	0.126	0.017	0.013	0.050	0.005	0.000	0.000	0.020	0.000	0.000
	(0.161)	(0.013)											(1994.1)				
2	0.049	−0.025	−1.307	0.833	−0.016	0.085	0.762	0.190	0.030	0.380	0.484	0.055	0.000	0.005	0.000	0.000	0.000
	(0.160)	(0.011)	(1.327)										(1994.1)				
3	0.121	−0.023	−2.225	0.842	−0.011	0.098	0.413	0.213	0.005	0.475	0.761	0.004	0.000	0.037	0.003	0.000	0.000
	(0.167)	(0.011)	(1.275)										(1995.1)				
4	0.201	−0.011	−0.496	0.717	0.318	0.403	0.000	0.190	0.000	0.142	0.064	0.936	0.000	0.002	0.000	0.262	0.000
	(0.229)	(0.011)	(0.057)										(1987.7)				
5	−0.019	−0.002	−0.034	0.206	−0.015	0.454	0.516	0.750	0.981	0.171	0.049	0.002	0.000	0.238	0.849	0.766	0.825
	(0.031)	(0.002)	(0.016)										(1981.8)				
6	−0.000	−0.001	−0.011	0.174	0.010	0.831	0.504	0.892	0.962	0.425	0.002	0.000	0.000	0.716	0.412	0.850	0.908
	(0.021)	(0.002)	(0.011)										(1981.8)				
								Seasonally Adjusted									
1	0.127	−0.111		1.543	0.114	0.000	0.092	0.005	0.000	0.116	0.397	0.011	0.000	0.035	0.179	0.000	0.000
	(0.247)	(0.025)											(1994.11)				
2	0.103	−0.106	−0.219	1.440	0.124	0.000	0.003	0.026	0.000	0.112	0.180	0.016	0.000	0.002	0.000	0.000	0.000
	(0.245)	(0.021)	(0.191)										(1994.11)				
3	0.168	−0.102	−0.089	1.359	0.119	0.000	0.001	0.070	0.000	0.040	0.060	0.004	0.000	0.078	0.001	0.004	0.000
	(0.222)	(0.019)	(0.169)										(1994.11)				
4	0.451	−0.087	−0.252	1.326	0.160	0.000	0.000	0.036	0.000	0.107	0.620	0.002	0.000	0.128	0.827	0.000	0.000
	(0.235)	(0.017)	(0.084)										(1994.8)				
5	0.027	−0.076	−0.047	1.376	0.137	0.000	0.000	0.033	0.000	0.000	0.001	0.000	0.000	0.000	0.000	0.000	0.000
	(0.170)	(0.015)	(0.061)										(1994.7)				
6	−0.009	−0.072	−0.070	1.400	0.109	0.000	0.001	0.033	0.003	0.000	0.011	0.000	0.000	0.003	0.000	0.000	0.000
	(0.166)	(0.015)	(0.060)										(1994.9)				
12	−0.051	−0.039	−0.047	0.914	0.070	0.005	0.154	0.463	0.403	0.002	0.156	0.031	0.034	0.590	0.398	0.000	0.000
	(0.121)	(0.012)	(0.029)										(1994.9)				

Notes: The table contains efficiency test results for different releases of annualized monthly growth rates of the Producer Price Index for Finished Goods over the period 1978.2-2001.12, based on data vintages for 1978.2-2004.1. See Table 7 for further details.

Table 9: Unbiasedness of Real-Time Growth Rates of Industrial Production - Structural Change Model

Release k	α_1	β_1	α_2	β_2	$\alpha_1 = \alpha_2 = 0$ $\beta_1 = \beta_2 = 0$	$\alpha_1 = 0$ $\beta_1 = 0$	$\alpha_2 = 0$ $\beta_2 = 0$
			Seasonally Unadjusted				
1	2.039	0.055	0.526	−0.020	0.000	0.000	0.021
	(0.359)	(0.017)	(0.230)	(0.010)			
2	1.268	0.030	0.442	0.001	0.000	0.000	0.012
	(0.316)	(0.011)	(0.156)	(0.008)			
3	0.663	0.002	0.003	0.004	0.001	0.000	0.584
	(0.158)	(0.006)	(0.096)	(0.004)			
6	0.221	−0.007	−0.023	0.000	0.136	0.040	0.754
	(0.089)	(0.006)	(0.033)	(0.002)			
12	0.149	0.000	−0.039	−0.002	0.102	0.027	0.494
	(0.056)	(0.003)	(0.036)	(0.003)			
			Seasonally Adjusted				
No evidence for structural change in the unbiasedness test regressions was found							

Notes: The table contains unbiasedness test results for different releases of annualized monthly growth rates of Industrial Production over the period 1963.1-2001.12, based on data vintages for 1963.1-2004.1, and based on estimating equation (7) with $\gamma_1 = \gamma_2 = 0$ imposed. The difference between these results and those reported in Tables 5 and 6 is that equation (7) imposes nonlinearity in the form of structural change on the unbiasedness test regression, while linearity is imposed when equation (4) is estimated (i.e. in Tables 5 and 6). See Table 5 for further details.

Table 10: Unbiasedness of Real-Time Growth Rates of the Producer Price Index for Finished Goods - Structural Change Model

Release k	α_1	β_1	α_2	β_2	$\alpha_1 = \alpha_2 = 0$ $\beta_1 = \beta_2 = 0$	$\alpha_1 = 0$ $\beta_1 = 0$	$\alpha_2 = 0$ $\beta_2 = 0$
			Seasonally Unadjusted				
1	−0.133	0.018	0.085	−0.038	0.085	0.274	0.053
	(0.084)	(0.013)	(0.059)	(0.016)			
3	−0.125	0.017	0.085	−0.038	0.095	0.317	0.053
	(0.085)	(0.013)	(0.059)	(0.016)			
4	1.570	−0.006	0.266	−0.021	0.000	0.000	0.061
	(0.290)	(0.028)	(0.116)	(0.014)			
			Seasonally Adjusted				
1	0.756	−0.065	0.181	−0.120	0.000	0.198	0.000
	(0.420)	(0.039)	(0.099)	(0.024)			
2	0.688	−0.061	0.170	−0.122	0.000	0.212	0.000
	(0.391)	(0.037)	(0.096)	(0.024)			
3	0.663	−0.057	0.178	−0.115	0.000	0.208	0.000
	(0.376)	(0.036)	(0.093)	(0.024)			
4	0.915	−0.056	0.255	−0.112	0.000	0.079	0.000
	(0.428)	(0.035)	(0.099)	(0.022)			
5	0.554	−0.038	0.121	−0.089	0.000	0.160	0.000
	(0.326)	(0.029)	(0.083)	(0.020)			
6	0.510	−0.035	0.144	−0.084	0.000	0.206	0.000
	(0.321)	(0.029)	(0.079)	(0.019)			
12	−0.284	0.022	0.144	−0.046	0.009	0.216	0.004
	(0.163)	(0.014)	(0.061)	(0.014)			

Notes: The table contains unbiasedness test results for different releases of annualized monthly growth rates of the Producer Price Index for Finished Goods over the period 1978.2-2001.12, based on data vintages for 1978.2-2004.1. Note that k=2 has not been added to the top panel of the table because this case yields identical results to those for the k=3 case. See Table 9 for further details.

Table 11: Unbiasedness of Real-Time Growth Rates of Industrial Production - Business Cycle Asymmetry Model

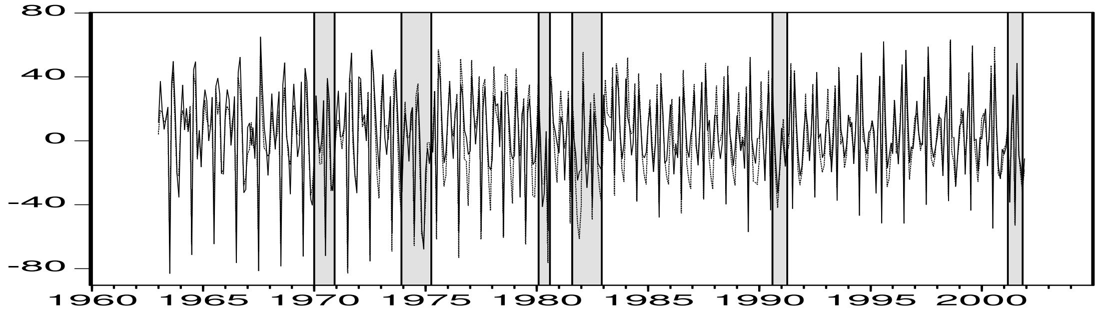
Release k	α_1	β_1	α_2	β_2	$\alpha_1 = \alpha_2 = 0$ $\beta_1 = \beta_2 = 0$	$\alpha_1 = 0$ $\beta_1 = 0$	$\alpha_2 = 0$ $\beta_2 = 0$
			Seasonally Unadjusted				
1	1.156	0.003	0.427	0.018	0.000	0.000	0.660
	(0.230)	(0.011)	(0.563)	(0.028)			
2	0.716	0.010	0.908	0.024	0.000	0.000	0.004
	(0.165)	(0.007)	(0.348)	(0.014)			
3	0.303	0.001	0.271	0.014	0.008	0.010	0.093
	(0.104)	(0.004)	(0.218)	(0.007)			
			Seasonally Adjusted				
1	1.290	-0.092	-0.499	-0.089	0.000	0.000	0.191
	(0.218)	(0.033)	(0.526)	(0.049)			
2	0.861	-0.037	0.574	0.002	0.000	0.000	0.073
	(0.174)	(0.025)	(0.297)	(0.030)			
3	0.551	-0.039	0.515	0.038	0.000	0.000	0.081
	(0.139)	(0.020)	(0.249)	(0.019)			
4	0.310	-0.046	0.036	0.009	0.091	0.029	0.635
	(0.121)	(0.019)	(0.130)	(0.010)			
5	0.310	-0.045	0.023	0.004	0.109	0.029	0.792
	(0.120)	(0.019)	(0.128)	(0.006)			
6	0.336	-0.049	0.005	0.003	0.056	0.012	0.801
	(0.116)	(0.019)	(0.127)	(0.006)			
12	0.260	-0.038	-0.034	0.002	0.089	0.021	0.809
	(0.097)	(0.015)	(0.131)	(0.006)			

Notes: The table contains unbiasedness test results for different releases of annualized monthly growth rates of Industrial Production over the period 1963.1-2001.12, based on data vintages for 1963.1-2004.1, and based on estimating equation (7) with $\gamma_1 = \gamma_2 = 0$ imposed. The difference between these results and those reported in Tables 5 and 6 is that equation (7) imposes nonlinearity in the form of asymmetric business cycle effects on the unbiasedness test regression, while linearity is imposed when equation (4) is estimated (i.e. in Tables 5 and 6). See Table 5 for further details.

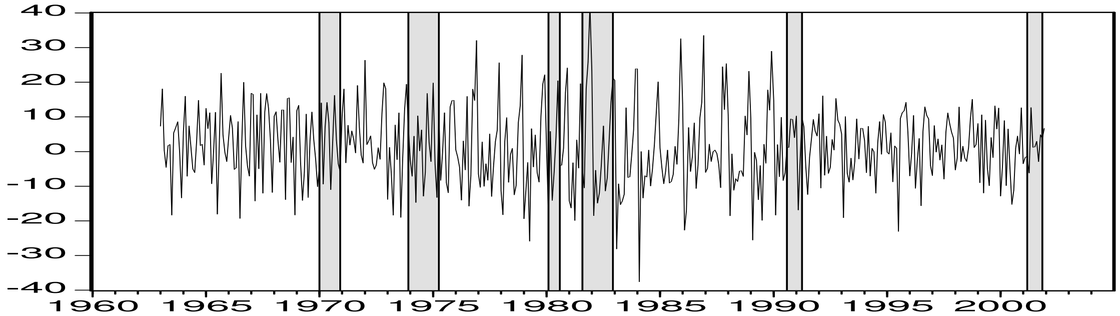
Table 12: Unbiasedness of Real-Time Growth Rates of the Producer Price Index for Finished Goods - Business Cycle Asymmetry Model

Release k	α_1	β_1	α_2	β_2	$\alpha_1 = \alpha_2 = 0$ $\beta_1 = \beta_2 = 0$	$\alpha_1 = 0$ $\beta_1 = 0$	$\alpha_2 = 0$ $\beta_2 = 0$
			Seasonally Unadjusted				
1	0.061	−0.020	−0.049	0.011	0.566	0.275	0.880
	(0.055)	(0.013)	(0.147)	(0.021)			
3	0.063	−0.021	−0.049	0.011	0.552	0.264	0.880
	(0.055)	(0.013)	(0.147)	(0.021)			
4	0.360	0.015	0.880	0.006	0.001	0.008	0.004
	(0.123)	(0.018)	(0.343)	(0.029)			
			Seasonally Adjusted				
No evidence for business cycle asymmetry in the unbiasedness test regressions was found							

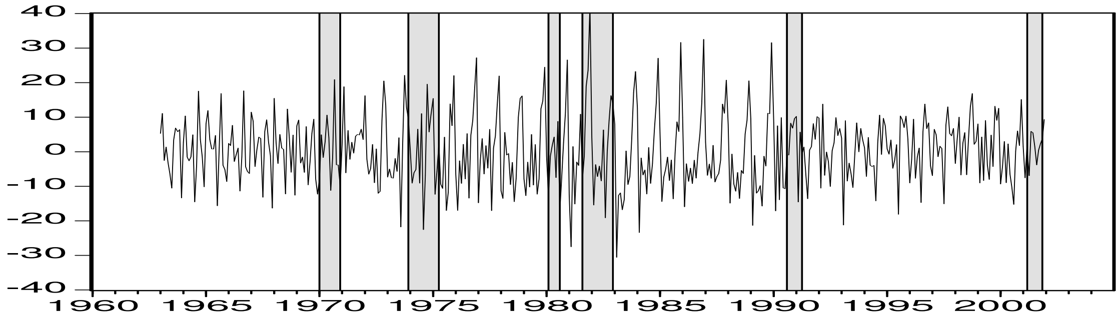
Notes: The table contains unbiasedness test results for different releases of annualized monthly growth rates of the Producer Price Index for Finished Goods over the period 1978.1-2001.12, based on data vintages for 1978.1-2004.1, and based on estimating equation (7) with $\gamma_1 = \gamma_2 = 0$ imposed. Note that k=2 has not been added to the top panel of the table because this case yields identical results to those for the k=3 case. See Tables 5 and 11 for further details.



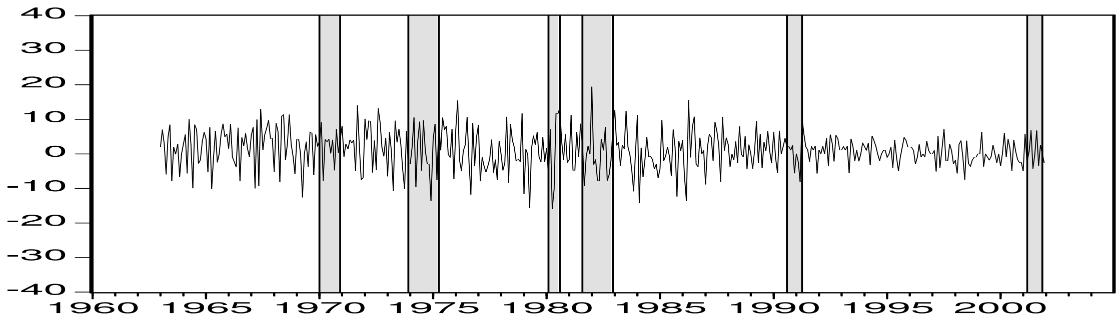
(a) First available (dashed line) and Fully revised (solid line) growth rates



(b) Complete revision

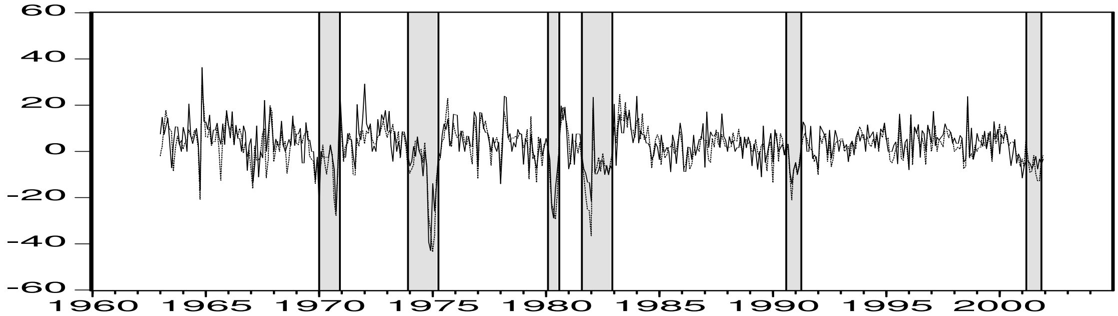


(c) Benchmark revision

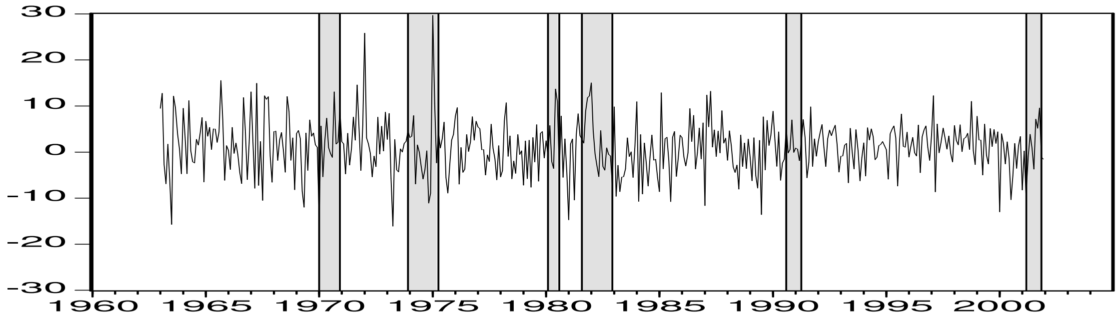


(d) Non-benchmark revision

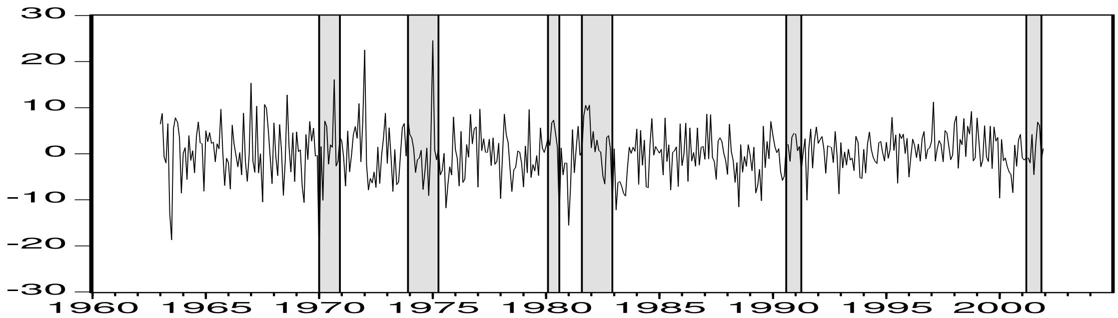
Figure 1: Real-time monthly growth rates of seasonally unadjusted industrial production.



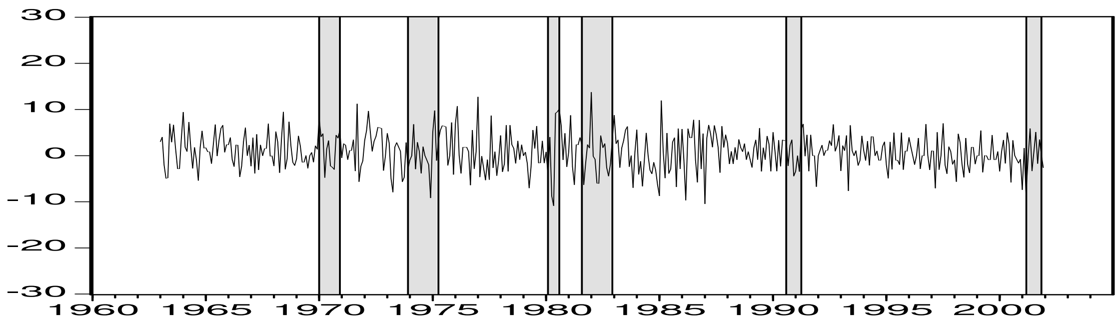
(a) First available (dashed line) and Fully revised (solid line) growth rates



(b) Complete revision

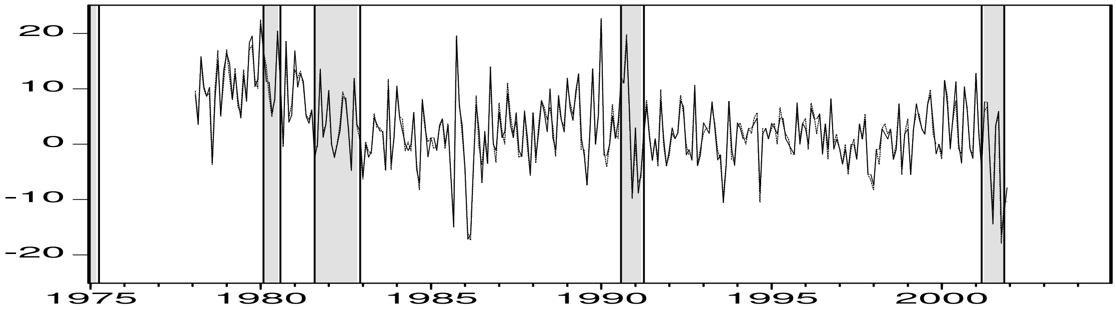


(c) Benchmark revision

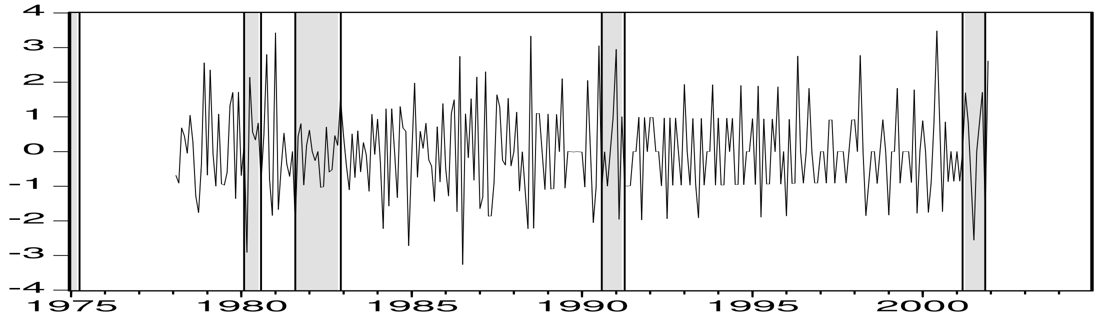


(d) Non-benchmark revision

Figure 2: Real-time monthly growth rates of seasonally adjusted industrial production.



(a) First available (dashed line) and Fully revised (solid line) growth rates



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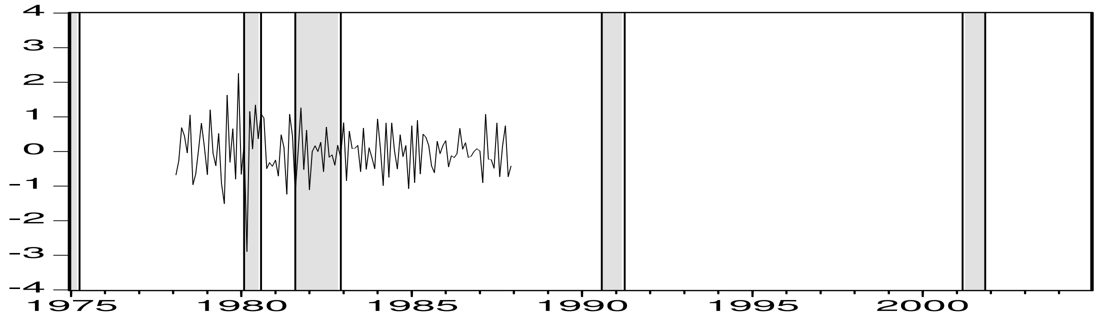
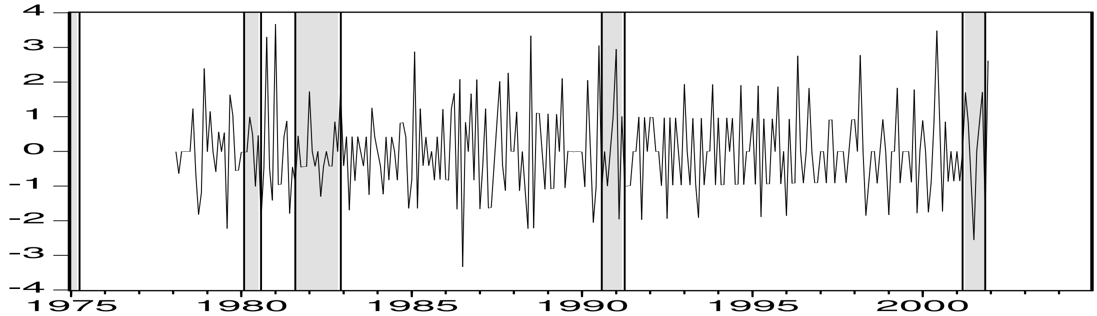
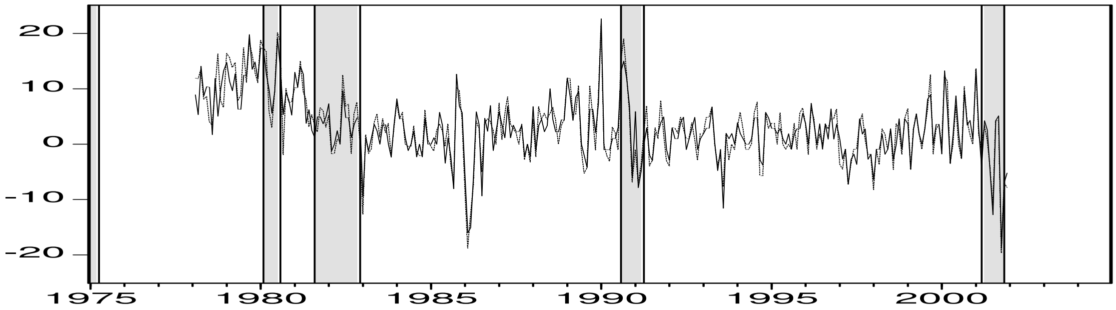


(c) Benchmark revision

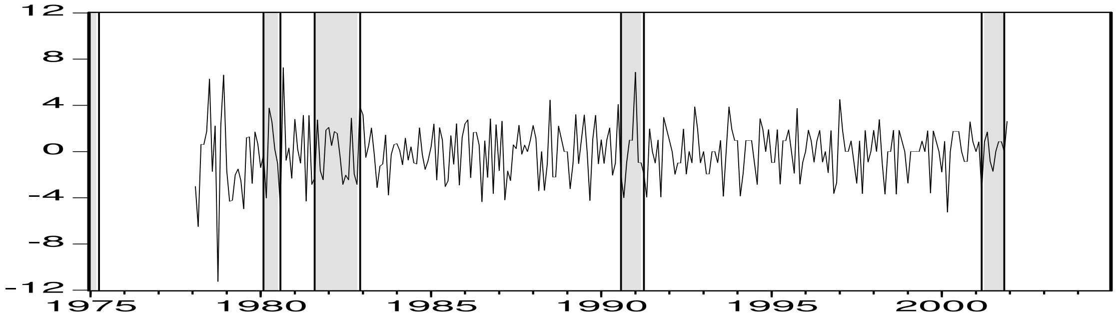


(d) Non-benchmark revision

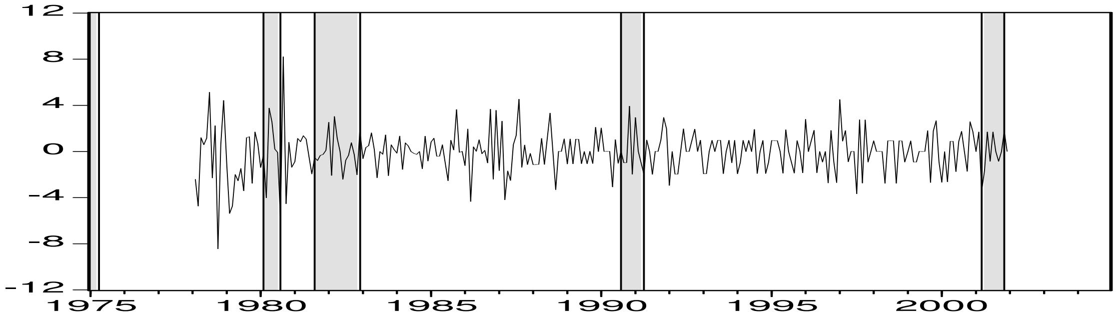
Figure 3: Real-time monthly growth rates of seasonally unadjusted producer price index for finished goods.



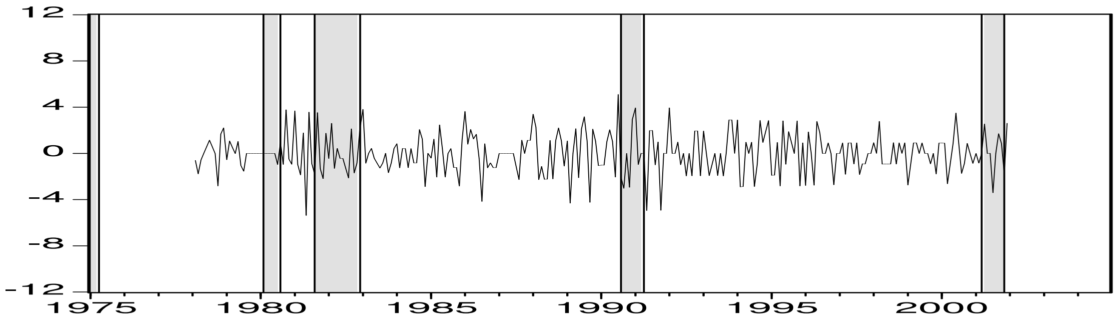
(a) First available (dashed line) and Fully revised (solid line) growth rates



(b) Complete revision



(c) Benchmark revision



(d) Non-benchmark revision

Figure 4: Real-time monthly growth rates of seasonally adjusted producer price index for finished goods.