Broken trend stationarity of hours worked

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Abstract

The estimated impact of a technology shock on hours worked using structural vector autoregressions depends to a great extent on whether or not the latter variable is considered to be integrated of first order. It is shown in this paper that a widely analyzed time series of hours worked per capita in the U.S. business sector evolves around a broken linear trend. When this fact is taken into account the unit root null is rejected by recently proposed suitable tests. Therefore, it can be stated that empirical specifications with hours in first differences are not recommended. On the contrary, it seems more appropriate to control for the presence of this shift in the deterministic component. We do this using a bivariate model for both productivity growth and hours worked. Our results suggest that technology improvements have a contractionary effect on hours only in the short run.

JEL codes: C12, C22, E00.

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1 Motivation

A burgeoning literature in recent years has tried to disentangle the effects that technology improvements have on hours worked. It was motivated by the different points of view in Galí (1999) and Christiano et al. (2003) on how the latter variable should be specified in the empirical model.

On the one hand, Galí (1999) introduced hours in first differences and found evidence of a negative impact of a positive technology shock on hours. This was interpreted as evidence against standard Real Business Cycle (RBC) models and more consistent with the New-Keynesian framework. On the other hand, Christiano et al. (2003) worked with this variable in levels obtaining

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a positive effect of technology improvements on hours worked. Some papers have tried to circumvent this problem by analyzing this issue using methodologies that are valid regardless of the way hours are specified (Pesavento and Rossi, 2005; Basu et al., 2006; Basistha et al., 2009).

The main cause behind this uncertainty about the most appropriate specification for hours worked is that applied procedures are not able to reject the null hypothesis of a unit root. Although there is no reason to think that hours worked per capita is a non-stationary variable, most studies work with this variable in first differences. This could easily be a consequence of the little effort carried out to empirically corroborate this impression. One reasonable explanation for concluding that hours is an I(1) variable might be the low power of conventional univariate unit root tests. Kappler (2008) has recently tried to overcome this limitation by the use of panel methods. However, his results do not give evidence against the unit root null. As an alternative, Gil-Alana and Moreno (2008) have used a multivariate fractional integration approach obtaining more evidence in favor of the stationarity of hours worked.

This paper focuses on the argument that the reason for the lack of evidence against the non-stationarity of hours worked is that applied unit root tests to date have neglected the presence of a shift in the deterministic component of this variable, leading to a reduction in their power (Perron, 1989). This changing trend has already been tackled by means of a quadratic specification in Galí (1999), Christiano et al. (2003) and Galí and Rabanal (2005), among others. More interestingly, Shields and Shields (2008) and Francis and Ramey (2008) give some insights into the determinants of the evolution of hours worked. Nonetheless, the influence of this changing trend on unit root testing has not been analyzed yet.

The suitability of a broken linear trend for a widely analyzed time series of hours worked per capita in the U.S. business sector is going to be tested for by the statistic proposed in Perron and Yabu (2007). Given the overlaps between unit root and structural break testing (Perron, 2006), this procedure is intended to be used as a pre-test because it is valid for both I(0) and I(1) time series. Once the presence of a broken linear trend is established, the unit root null hypothesis will be assessed under this specification for the deterministic component by the tests developed in Carrión-i-Silvestre et al. (2008). Jointly with those in Kim and Perron (2009), their main advantage with respect to existing ones is that they consider the presence of shifts both under the null and the alternative hypotheses. Our results show that hours per capita evolve around a broken linear trend in a stationary manner. Therefore, it is concluded that introducing this variable in first differences into structural vector autoregressions (SVARs) does not make sense and can lead to biases (Erceg et al., 2005).

In line with the suggestions in Ng and Vogelsang (2002) and Lütkepohl and Krätzig (2004), a further contribution of this paper is to analyze the effects that a technology shock has on hours worked by simultaneously controlling for the presence of shifts in the deterministic components of the two variables involved in the bivariate SVARs commonly used in this context. It tries to complement the attempts made by previous studies that only deal with shifts in the deterministic component of one of the two time series (see Fernald, 2007; and Canova et al., 2008; for some recent examples). The impulse-response function analysis shows that a technology improvement has a negative impact on hours within the first year. After that, hours worked per capita increase.

These results are similar to those found in Basu et al. (2006) both using an augmented-growth-accounting approach and introducing a 'purified' technology measure in the SVAR framework. These authors interpret their findings as consistent with Dynamic General Equilibrium (DGE) models with sticky prices.

The rest of the paper is structured as follows. After describing the data source and the variables analyzed in Section 2, the presence of a broken linear trend in hours worked per capita is established in Section 3. In addition, this specification for the deterministic component is compared with that of a quadratic trend and the unit root null hypothesis is tested. Section 4 estimates the effect that a positive technology shock has on hours worked controlling for the presence of deterministic shifts in both productivity growth and hours worked. Finally, Section 5 concludes.

2 Data source

The variables analyzed in this paper are an updated version of those in Christiano et al. (2003). They have been extracted from the Federal Reserve Bank of St. Louis (FRED®) which, at the same time, are compiled from the Bureau of Labor Statistics (U.S. Department of Labor). The data has a quarterly frequency and cover the period from 1948:Q1 to 2007:Q4.

Hours worked refer to the business sector of the economy (HOABS). They are reported in terms of the year 1992 and have been previously seasonally adjusted. This variable has been converted into per capita terms dividing it by the civilian non-institutional population over 16 (CNP16OV). The latter has been changed to a quarterly frequency by averaging the monthly observations and then transformed into an index.

As has already been noted, we will study to what extent the properties of the deterministic component of hours worked affect the results obtained from the analysis of the effects that a technology shock has on this variable. For this reason, the output per hour worked in the business sector (OPHPBS) will also be used below. As was also the case of hours worked, this variable is reported in terms of the year 1992 and its seasonal component has been removed. Finally, note that the analysis will be carried out with the natural logarithm of hours worked per capita and output per hour (productivity, hereafter).

3 Unit root and shift in trend testing

The presence of a unit root in hours worked per capita will first be analyzed by means of standard univariate unit root tests. The alternatives implemented are those discussed in Ng and Perron (2001) because they have good size and power performance. One way to achieve these is by the application of a Generalized Least Squares (GLS) detrending method (Elliot et al., 1996) before testing the unit root null hypothesis.

Denoting the time series under scrutiny as $\{y_t\}_{t=0}^{T}$, assume that its data generating process can be expressed as the sum of a deterministic (d_t) and an autoregressive (u_t) component:

$$\begin{aligned} y_t &= d_t + u_t \\ d_t &= \psi z_t \\ u_t &= \alpha u_{t-1} + v_t \end{aligned} \tag{1}$$

where

$$\begin{aligned} v_t &= \theta(L) e_t \\ \theta(L) &= \sum_{i=0}^{\infty} d_i L^i \\ -1 &< \alpha \leq 1 \\ e_t &\sim iid(0, \sigma_e^2) \end{aligned} \tag{2}$$

L is the lag operator and the deterministic component will be assumed to be made up by both a constant and a trend $z_t = (1, t)'$.

Let $(y_0^{\bar{\alpha}}, y_t^{\bar{\alpha}}) = (y_0, (1 - \bar{\alpha}L) y_t)$ for some $\bar{\alpha} = 1 + \frac{\bar{c}}{T}$ and $t = 1, ..., T$. Using this notation, the GLS detrended time series for y_t is given by:

$$\tilde{y}_t = y_t - \hat{\psi} z_t \tag{3}$$

where

$$\begin{aligned} \hat{\psi} &= \arg\min S(\bar{\alpha}, \psi) \\ S(\bar{\alpha}, \psi) &= (y^{\bar{\alpha}} - \psi z^{\bar{\alpha}})'(y^{\bar{\alpha}} - \psi z^{\bar{\alpha}}) \end{aligned} \tag{4}$$

Given the specification for the deterministic component, $\bar{c}$ has been set to -13.5.

In addition, Ng and Perron (2001) also proposed a modification to the information criteria used to choose the number of autoregressive lags to control for the possible presence of autocorrelation when testing for unit roots. It consists of imposing the null hypothesis and including a stochastic term in the penalty factor. Following the suggestion in Perron and Qu (2007), OLS detrended data is considered when calculating these modified criteria.

[**Insert Table 1 here**]

Our results from the application of the GLS detrending-based unit root tests to the hours worked per capita time series are presented in the second column of Table 1. As has been found in previous studies, the null hypothesis of a unit root cannot be rejected at conventional significance levels by any of the applied tests. This would lead us to think of introducing this variable in first differences in multivariate systems that do not allow for the presence of cointegration relationships.

The evolution of hours worked per capita during the sample period analyzed has been plotted in Figure 1. It can be observed that this time series evolves around a more complicated trend than a simple linear one. Moreover, this fact is a reasonable explanation for the inability of the standard unit root tests applied before to reject the unit root null because they have low power in the presence of changes in the deterministic components (Perron, 1989).

[Insert Figure 1 here]

As noted by Perron (2006), not only testing for unit roots is complicated by structural breaks. In addition, testing for shifts in the deterministic component depends on the order of integration of the variable analyzed. For this reason, Perron and Yabu (2007) proposed a method for detecting trend shifts that is valid for both I(0) and I(1) time series. Given the virtues of this procedure, it is intended to be used as a pre-test before studying the integration order of a given time series.

The most general case where the possibility of a structural change in both the intercept and the slope are allowed has been considered. According to (1), it is going to be specified that:

$$\begin{aligned} z_t &= (1, DU_t, t, DT_t)' \\ \psi &= (\mu_0, \mu_1, \beta_0, \beta_1) \end{aligned} \tag{5}$$

where

$$\begin{aligned} DU_t &= 1(t > T_b) \\ DT_t &= 1(t > T_b)(t - T_b) \end{aligned} \tag{6}$$

T_b is the break date, which can be expressed as a proportion $\lambda \in (0, 1)$ of the sample size: $T_b = [\lambda T]$. $[\cdot]$ denotes the function that returns the largest integer that is less than or equal to the argument and $1(\cdot)$ is the indicator function.

The null hypothesis of interest is that of no trend shift. That is, $H_0 : \mu_1 = \beta_1 = 0$. Assuming that the autoregressive order is higher than one and also assuming an unknown breakpoint, Perron and Yabu (2007) proposed an approach based on a Feasible Quasi-GLS super-efficient estimator of α when the sum of the autoregressive parameters is equal to one. As a result, an Exp-Wald type (Andrews and Ploberger, 1994) test statistic is obtained.

The resulting value of the Exp-W test from the application of this procedure to the hours worked per capita in the U.S. business sector time series is reported in the lower panel of Table 1. A trimming of 25% both at the beginning and the end of the sample has been applied. It implies that only shifts in the central part of the period analyzed are allowed. This is the equivalent of saying that the possibility of detecting shifts in the extremes has been avoided because we are working with the case of a single break. According to the asymptotic critical values reported in the original paper for this trimming parameter and trend shift model, the null hypothesis of no break can be rejected at a 10% significance level. Given this evidence, it seems appropriate to take this trend shift into account when testing for the unit root null.

Most of the existing literature dealing with the issue of unit root testing in the presence of unknown breaks in the deterministic component only considered them under the alternative hypothesis of stationarity. Kim and Perron (2009) have recently proposed a more suitable testing procedure that allows a break both under the null and the alternative hypotheses. If there is a true break, its limiting distribution is equal to that of the known break case and the test has a correct size and improved power. The tests that have been applied in this paper are those developed by Carrión-i-Silvestre et al. (2008). These authors have extended the GLS detrending-based tests described at the beginning of this section to allow for multiple breaks both under the null and the alternative hypotheses. Nevertheless, only the case of a single trend shift is going to be analyzed.

Our test statistics for the hours worked per capita time series are reported in the third column of Table 1. The break date has been estimated by the minimization of the sum of squared residuals and is located in the last quarter of the year 1974. It can be observed that the unit root null that hours per capita evolve around a broken linear trend in a non-stationary manner is rejected at the 5% significance level by all the applied tests.

[**Insert Table 2 here**]

Previous studies that have specified a deterministic component for hours worked that is not a linear trend, have opted for a quadratic one. For this reason, a comparison of simple adjustment and diagnostic statistics for three alternative empirical specifications of the deterministic component of hours are reported in Table 2. Reinforcing the previous analysis, the best fit is obtained for the broken linear trend specification, for which the highest coefficient of determination and log-likelihood are obtained. In addition, this specification also leads to the lowest standard errors, sum of squared residuals and information criteria. On the contrary, the worst fit corresponds to the simple linear trend.

Finally, it should be noted that Ayat and Burridge (2000) and Harvey et al. (2008) have analyzed the issue of testing for a unit root in the presence of a quadratic trend. The value of the Dickey-Fuller unit root test with GLS detrending using this specification of the deterministic component[1] is -3.38. Ayat and Burridge calculated finite sample critical values[2], those for a sample size of 250 observations being equal to -3.48 and -3.20 for the 5 and 10% significance levels, respectively. Therefore, it can be concluded that the evidence against the unit root null is slightly weaker when a quadratic trend is specified for the deterministic component instead of a broken linear one.

[1] In this case $\bar{c}$ has been set to be -18.5.

[2] Appendix B, page 95.

4 Effects of productivity shocks on hours

In light of the evidence presented in the previous section, it can be stated that working with hours per capita as if they were difference stationary might not be correct and can lead to biased estimations (Erceg et al., 2005). The alternative chosen below consists of controlling for the trend shift in hours rather than trying to achieve stationarity by first-differencing this variable when analyzing the effects of a technology shock on it.

As noted in Ng and Vogelsang (2002), although there is evidence of breaks in the form of mean and/or trend shifts in many macroeconomic time series, vector autoregression (VAR) estimations usually ignore them. However, it seems more appropriate to introduce breaks into multivariate analyses when they are observed at a univariate level. In addition, these authors also demonstrated that inference based on the estimated VAR is invalid when mean shifts are omitted.

The importance of correctly specifying the deterministic components in VARs has also been emphasized by Lütkepohl and Krätzig (2004). As has already been mentioned, in cases where the effects of technology shocks on hours are analyzed, Galí (1999) was the first to use a quadratic specification to detrend the hours worked time series. Fernald (2007) took into account two mean shifts in productivity growth during the period 1950-2004. More recently, Canova et al. (2008) have checked the robustness of the estimated effects to several detrending alternatives for both productivity growth and hours worked. Nevertheless, to date, no study has considered the presence of changing deterministic components simultaneously in productivity growth and hours worked.

Following Fernald (2007), the methodology developed in Bai and Perron (1998) has been applied to our updated version of the productivity growth in the U.S. business sector time series. Our test statistics are shown in Table 3. The results are slightly different to those in Fernald's paper, which may be a consequence of the different time periods analyzed. Using a longer time span only evidence of a single level shift in productivity growth is found at the 10% significance level. These findings are confirmed by the application of the Exp-W test presented in the previous section for the case of one shift in the intercept. The estimated break date is relatively close to that for hours worked and is located in the fourth quarter of 1972.

[**Insert Table 3 here**]

The productivity growth time series and its estimated sub-sample means have been plotted in Figure 3. It can be observed that the level shift has implied a reduction in mean productivity growth. In addition, the estimation results for three different specifications for the mean have also been reported in Table 3. Although an increase in the fit is obtained when introducing a level shift, the consideration of an additional one does not lead to a significant improvement and its estimated value for the mean is not very different to that in the second subsample. Moreover, it should be noted that working with the shifts detected by Fernald (2007) in 1973:Q2 and 1997:Q2 does not change the conclusions drawn below[3].

[3] These results are available from the author upon request.

[Insert Figure 2 here]

To estimate the effect that a technology shock has on hours worked per capita, begin by considering a bivariate VAR of order p for the first difference of productivity (Δp_t) and hours worked (n_t). This VAR(p) representation captures dynamic interactions between these two variables:

$$Y_t = A_1 Y_{t-1} + ... + A_p Y_{t-p} + \xi_t \tag{7}$$

where $Y_t = (\Delta p_t^*, n_t^*)'$. The asterisks reflect that these two variables have been adjusted for their deterministic components before entering the bivariate empirical model[4]. $\xi_t = (\xi_{1t}, \xi_{2t})'$ are the unobservable error terms assumed to be independently distributed as $\xi_t \sim (0, \Sigma_\xi)$, where $\Sigma_\xi = E(\xi_t^{'}\xi_t)$. Four autoregressive lags will be considered in what follows.

VARs are known to be 'reduced form' models because they basically summarize the dynamic properties of the data. However, our interest is in disentangling the effects that a shock in one of the variables has on the rest of the system. These can be determined by the use of structural VARs where identification focuses on the errors which are interpreted as a linear combination of exogenous shocks. Under the assumption of orthogonality, it is possible to analyze the dynamic impact of isolated impulses. It should be emphasized once more that, although the deterministic terms are not affected by these shocks, it is also necessary to adjust for their presence before analyzing the dynamic interactions between the variables.

Because our two variables of interest are stationary, they can be expressed as a distributed lag of two types of shocks:

$$\begin{bmatrix} \Delta p_t^* \\ n_t^* \end{bmatrix} = \begin{bmatrix} B_{11}(L) & B_{12}(L) \\ B_{21}(L) & B_{22}(L) \end{bmatrix} \begin{bmatrix} \varepsilon_t^p \\ \varepsilon_t^n \end{bmatrix} \tag{8}$$

ε_t^p and ε_t^n are the technology (productivity) and non-technology shocks, respectively.

The main issue when dealing with SVARs is the identification of the shocks. Following Galí (1999), this identification has been achieved by imposing long-run restrictions *à la* Blanchard and Quah (1989). It is assumed that non-technology shocks do not affect productivity growth. This identification restriction implies in (8) that $B_{12}(L) = 0$. That is to say, the matrix of long-horizon multipliers is lower-triangular.

The dynamic effects of structural shocks have been investigated by impulse-response functions. They have been plotted in Figure 3 for all possible combinations of shocks and affected variables. Studentized Hall (1992)'s bootstrap 95% confidence bands[5] have also been reported.

[Insert Figure 3 here]

[4] The reason behind substracting the deterministic component first is that, given that it contains the behavioral relations, the stochastic term is that of primary interest in econometric analyses.

[5] Calculated using 200 replications.

Our effects of non-technology shocks on productivity and hours are similar to those in Galí (1999). This is true for both the first-differenced and the detrended specifications of hours worked. A positive non-technological shock increases productivity at first but this effect turns negative after four periods. The influence of this shock on hours is always positive and the estimated impulse-response function is hump-shaped. The impact of a positive technology shock on productivity is also similar to those already established in the literature because technology improvements always increase productivity in the subsequent periods.

Our findings with respect to the effect of a positive technology shock on hours worked are different from most previous papers. After an initial negative impact, the impulse-response function becomes positive after four periods. These results are similar to those in Basu et al. (2006) who, using an augmented-growth-accounting framework found a very similar pattern for the response of hours to technology improvements. It should be emphasized that their conclusions did not depend on the way hours were specified. Moreover, the same effects were obtained when using a 'purified' technology series in a SVAR with long-horizon restrictions. These authors interpreted their findings as consistent with DGE models with sticky prices.

Finally, note that, consistent with the arguments in Ng and Vogelsang (2002), these estimated effects do not significantly change when shifts in the deterministic components are modeled directly in the VAR instead of removing them in a first step[6].

5 Conclusions

This paper has empirically established that the deterministic component of a widely analyzed U.S. hours worked per capita in the business sector time series can be specified as a broken linear trend. Taking this into account, recently proposed suitable tests give evidence against the unit root null hypothesis for this variable. Therefore, it is concluded that there is no reason to introduce it in first differences into SVARs when estimating the effects of a technology shock. As an alternative and trying to complement previous attempts in the literature, we proposed to control for the presence of shifts in the deterministic components of both productivity growth and hours worked, simultaneously. Our results contrast with the predictions of standard RBC models, but only in the short run.

[6] These results are also available from the author upon request.

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Table 1: Unit root and trend shift testing. Hours worked per capita, 1948:Q1-2007:Q4.

	No structural break	Single structural break
Augmented Dickey-Fuller (ADF)	-1.44	-3.52**
Elliot-Rothenberg-Stock (P_T)	18.51	6.60**
Modified P_T (MP_T)	17.60	6.59**
Phillips-Perron (Z_α)	-5.20	-26.75**
Modified Z_α (MZ_α)	-5.16	-26.63**
Modified Sargan-Bhargava (MSB)	0.31	0.13**
Modified Z_t (MZ_t)	-1.59	-3.58**
Exp-W	—	2.16*
Break date	—	1974:Q4

Note: Hours worked per capita are expressed in natural logarithms. They are equal to the hours worked in the U.S. business sector divided by the civilian non-institutional population over 16 (1992=100). Unit root tests are those based on the quasi-GLS detrending method discussed in Ng and Perron (2001). The deterministic component is made up of a constant and a trend. The number of augmentation lags have been selected using the MAIC criterion calculated following the suggestion in Perron and Qu (2007). The Exp-W test is that of Perron and Yabu (2007) for Model III with a trimming of 25%. The break date has been estimated by minimizing the sum of squared residuals. ***, ** and * denote rejection of the null at the 1, 5 and 10% significance level, respectively.

Table 2: Comparison of raw deterministic component adjustments. Hours worked per capita, 1948:Q1-2007:Q4.

	Linear trend	Quadratic trend	Broken linear trend
Constant	4.71^{***}	4.78^{***}	4.77^{***}
Trend	$-4.5\cdot10^{-4\ ***}$	$-2.23\cdot10^{-3\ ***}$	$-1.52\cdot10^{-3\ ***}$
Trend^2	—	$7.42\cdot10^{-6\ ***}$	—
Constant after shift	—	—	4.75^{**}
Trend after shift	—	—	$5.31\cdot10^{-4\ ***}$
Adjusted R^2	0.30	0.61	0.69
Standard error	0.05	0.04	0.03
SSR	0.55	0.31	0.24
Log-likelihood	388.42	458.34	489.62
AIC	-3.22	-3.79	-4.05
BIC	-3.19	-3.75	-3.99
F-statistic	102.04^{***}	184.67^{***}	182.58^{***}

Note: Hours worked per capita are expressed in natural logarithms. They are equal to the hours worked in the U.S. business sector divided by the civilian non-institutional population over 16 (1992=100). Trend shift takes place in 1974:Q4. The break date has been estimated by minimizing the sum of squared residuals. Inferences drawn from Newey-West standard errors. ***, ** and * denote significant at the 1, 5 and 10% significance level, respectively.

Table 3: Level shift testing. Productivity growth, 1948:Q2-2007:Q4.

	Statistic	Break date(s)
UD max	9.32*	
WD max	9.32*	
SupF(1)	9.32*	1972:Q4
SupF(2)	6.20	1972:Q4, 1982:Q2
SupF(3)	4.86	1972:Q4, 1972:Q4, 1996:Q4
SupF(4)	3.95	1960:Q3, 1972:Q4, 1972:Q4, 1996:Q4
SupF(2|1)	4.61	
SupF(3|2)	2.64	
SupF(4|3)	0.41	
Exp-W	1.43*	

Note: 15% of sample trimmed. Productivity growth is equal to the first difference of the natural logarithm of output per hour worked in the U.S. business sector (1992=100). The Exp-W test is that of Perron and Yabu (2007) for Model I. The rest of the test statistics are those proposed by Bai and Perron (1998). Covariance matrices are robust to heteroskedasticity and autocorrelation and AR pre-whitening has been used. Heterogeneous moment matrices across subsamples have also been allowed . ***, ** and * denotes rejection of the null at the 1, 5 and 10% significance level, respectively. Break dates have been estimated by minimizing the sum of squared residuals.

Table 4: Comparison of raw deterministic component adjustments. Productivity growth, 1948:Q2-2007:Q4.

	No shift	Single level shift	Two level shifts
Constant 1	0.60***	0.78***	0.80***
Constant 2	—	0.47***	0.36***
Constant 3	—	—	0.30*
Adjusted R^2	0.00	0.03	0.04
Standard error	0.88	0.87	0.86
SSR	184.45	178.89	175.13
Log-likelihood	-308.16	-304.51	-301.97
AIC	2.59	2.56	2.55
BIC	2.60	2.59	2.60
F-statistic	—	7.37***	6.28***

Note: Productivity growth is equal to the first difference of the natural logarithm of output per hour worked in the U.S. business sector (1992=100). Level shift takes place in 1972:Q4 when a single break is considered and in 1973:Q2 and 1997:Q2 in the two level shifts case. Break dates have been estimated by minimizing the sum of squared residuals. Inferences drawn from Newey-West standard errors. ***, **and * denote significant at the 1, 5 and 10% significance level, respectively.

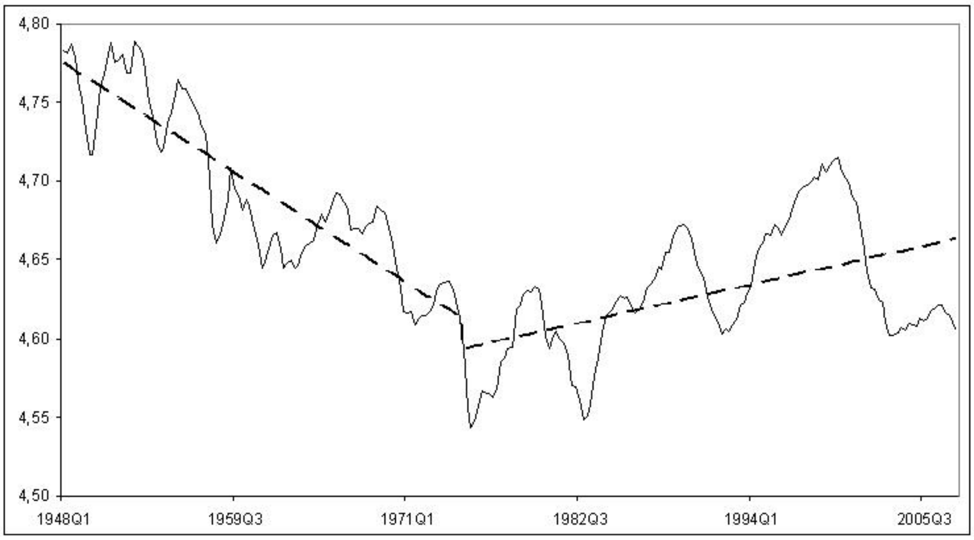

Figure 1: Hours worked per capita (1992=100, in natural logs) and adjusted broken linear trend, 1948:Q1-2007:Q4.

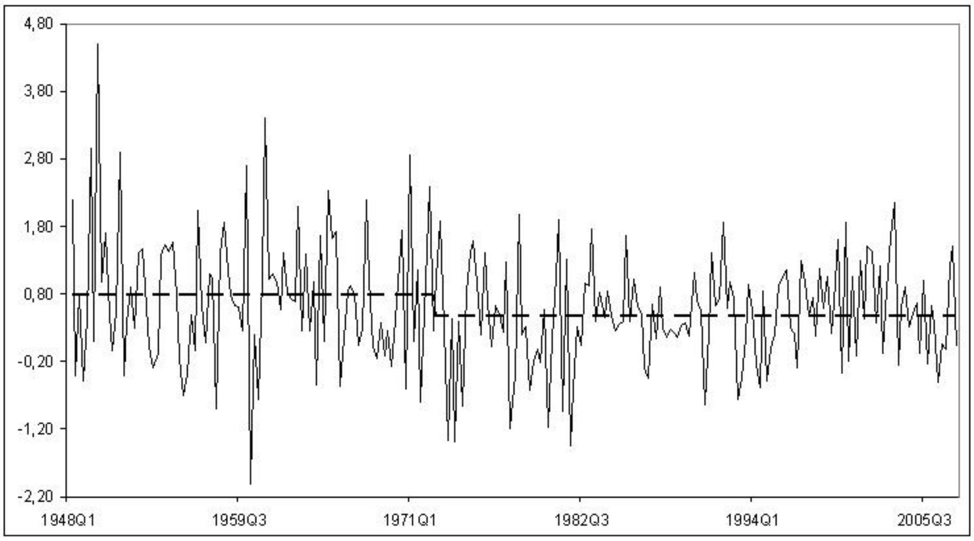

Figure 2: Productivity growth (percentage terms) and adjusted mean shift, 1948:Q2-2007:Q4.

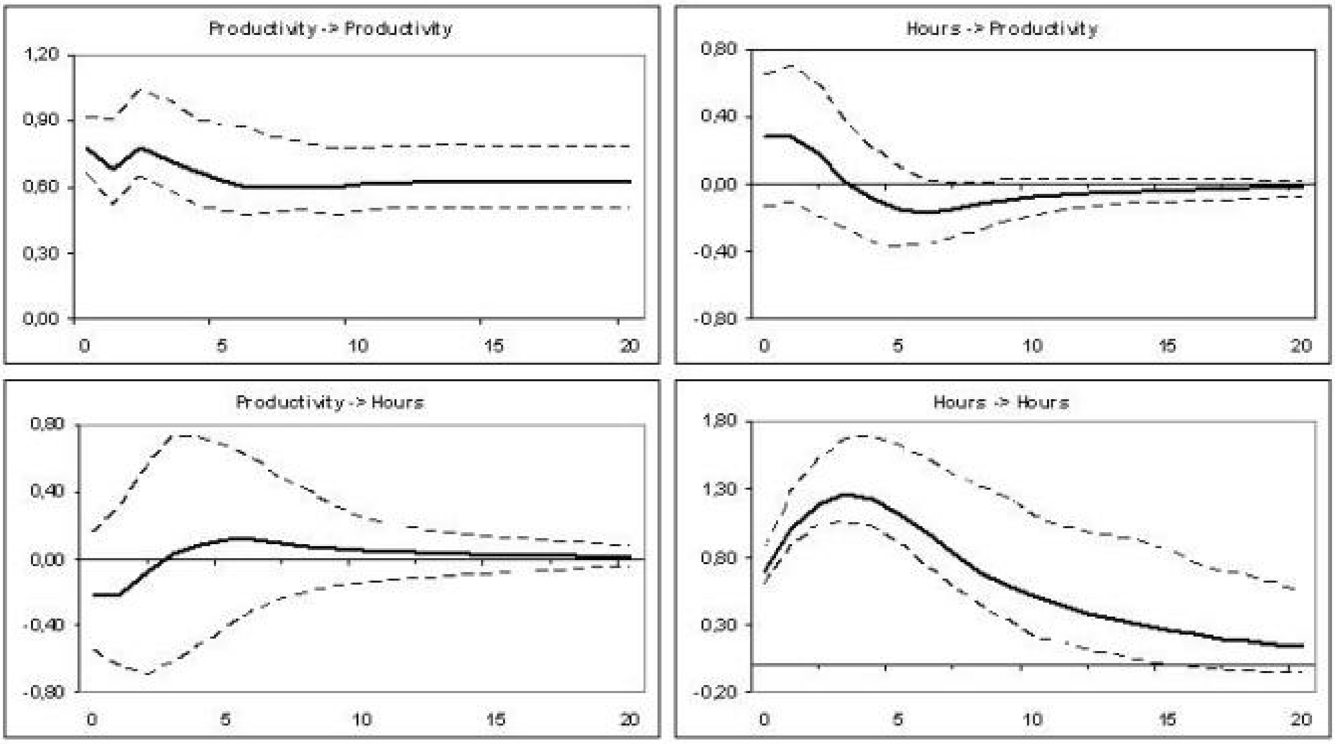

Figure 3: Impulse-response functions (percentages) from a bivariate SVAR with demeaned productivity growth and detrended (log) hours per capita. Studentized Hall (1992)'s 95% bootstrap confidence bands reported (200 replications, dotted lines).