Commodity Price Booms and Breaks: Detection, Magnitude and Implications for Developing Countries

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Abstract[1]

There has been much interest of late regarding the current commodity "super cycle". However, even sizing the current boom implies knowledge of long-run trends that are notoriously difficult to estimate. This paper uses new techniques to identify breaks in commodity prices and estimate trends and cointegrating relationships and argues that the weight of evidence is against a stable declining commodity terms of trade. The results are used to characterize the current boom and, assuming no new break, how commodity prices would be expected to return to the estimated "equilibrium". The paper also discusses implications for commodity-dependent developing countries.

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1 Introduction and Motivation

Given current relatively high commodity prices, it is easy to forget the Prebisch-Singer hypothesis (developed independently by Raúl Prebisch and Hans Singer around 1951), which suggested that commodity prices are in a long-run steady decline relative to manufactured goods prices. It is also perhaps too easy to forget the sharp busts in commodity prices, particularly those after the huge boom of 1920 and the early 1970s. Early papers that considered previous versions of the dataset we employ below include Grilli and Yang (1989) and Ardeni and Wright (1992), who argued for stable and declining commodity terms of trade, and Cuddington and Urzúa (1989), von Hagen (1989) and Powell (1991), who rejected this hypothesis. This last paper, Powell (1991), applied cointegration techniques and argued that, with just three structural breaks over one hundred years, commodity prices and manufactured unit values (MUVs) were cointegrated with a cointegrating vector of unity. Interestingly the two largest negative breaks occurred after the 1920 and early 1970 booms, which many considered to have a speculative or bubble component. The explanation offered was that a boom might create incentives for innovations to reduce production costs and develop substitutes such that when the reasons for the boom subside, a new equilibrium might emerge with lower relative commodity prices. Subsequent papers including Bleaney and Greenaway (1993), Lutz (1999), Newbold, Pfaffenzeller, and Rayner (2005) and Cashin and McDermott (2002) continue to dispute whether commodity prices are in long-run decline using cointegration techniques and more recent methodologies for detecting breaks.

Understanding commodity price processes is critical for commodity-dependent countries. For example, in the Latin American and Caribbean region some eight countries derive about a third of their fiscal revenues on average from non-renewables, and agricultural commodities are important foreign currency and tax earners in a further set of countries; see Corbacho, Fretes, and Lora (2013). Knowing the persistence of prices, the likelihood that a structural break has occurred and how fast prices tend to revert to an equilibrium value is all critical information for macroeconomic policymaking for such countries. As is well known in the financial economics literature, if a break goes undetected or is ignored, it may induce forecast failures, and non-constancy in model parameters will affect model interpretability. In particular, the identification of breaks and the development of a stable model for commodity prices are important elements for the appropriate design of a stabilization fund for commodity revenues, as implemented by Chile and other countries.[2]

[2] Chile established a stabilization fund for copper revenues in 1985. A committee determines the "long-run"price of copper and then savings placed into the fund or the usage of the fund is dependent on the relationship between the current price and the "long-run"price, see Fuentes and Piedrabuena (2012) on the taxation system for copper in Chile, further details regarding the stabilization fund and discussion.

In this paper we employ new techniques to analyze these issues. Specifically, we apply the Impulse Indicator Saturation (or IIS) and the Step-Indicator Saturation (or SIS) methodologies to endogenously identify breaks in long-term commodity price series. Having estimated relevant breaks, we estimate a set of error correction models for commodity prices. The IIS and the SIS techniques provide a general procedure for analyzing model constancy and for detecting anomalies such as structural breaks. Both IIS and SIS are generic tests for an unknown number of structural breaks, occurring at unknown times, with unknown duration and magnitude, anywhere in the sample; see Hendry (1999), Hendry, Johansen, and Santos (2008), Johansen and Nielsen (2009), Hendry and Santos (2010) and Doornik, Hendry, and Pretis (2013) for further discussion. It may be considered as a generalization of techniques such as recursive estimation, rolling regressions, the Chow (1960) predictive failure statistic, the Andrews (1993) unknown breakpoint test, the Bai and Perron (1998) multiple breakpoint test and intercept correction (in forecasting or estimation). Given the knowledge of structural breaks, we are then in a position to analyze cointegration relationships and to estimate a set of vector error correction models.

With our empirical results we then comment on the current boom in commodity prices. We compare the magnitude of this boom to previous ones. We take as a boom observations that lie outside statistical tolerance bounds. We also estimate the predicted behavior for future commodity prices assuming that a break has taken place, and hence how prices would be expected to move towards the estimated equilibrium and discuss the implications for commodity-dependent countries.

The paper is organized as follows. In the next section we review relevant theory regarding commodity prices and we also review the Prebisch-Singer hypothesis. In Section 3 we introduce the IIS/SIS methodology and apply this to the long-run deflated commodity price series. In Section 4 we then employ these results and estimate a set of vector error correction models for commodity prices; we consider variants that would nest the Prebisch-Singer hypothesis in a more general model and hence construct a test for the existence of a stable declining terms of trade trend. In Section 5 we use our results to analyze the current boom for commodity prices to size the boom, and we consider the implications for commodity producers. Section 6 concludes.

2 Modeling Commodity Prices and the Prebisch-Singer Hypothesis

There are several strands of the commodity price literature. One strand is theoretical and considers the intertemporal process for commodities given the possibility of storage. Early milestones include Hotelling (1931), Samuelson (1957) and Gustafson (1958). Gustafson was the first to develop a model of intermittent commodity production (e.g., annual harvests), but one where agents decide rationally whether to store output to the next crop year or not. A complication is that whether it is profitable to store today to next year, depends on the expectation that storage will be profitable from next year to two years' time, and so on. Williams and Wright (1991) and Deaton and Laroque (1992) independently considered the same question and generalized Gustafson's model to provide a more elegant solution to this dynamic programming problem.

The Deaton and Laroque (1992, 1995) models yield simulated price paths for commodities that exhibit sharp peaks and more persistent valleys. When the availability (the amount stored from the previous year plus this year's production) of the commodity is low, prices rise quite dramatically and price peaks correspond to periods where essentially all availability is consumed. Price volatility may be high and tends to be higher at higher price levels. At the other extreme, when there is ample availability, prices are low and persistent, akin to a random walk. Commodities might be thought of in terms of their perishability affecting the costs of storage. At one extreme, gold hardly deteriorates, and hence prices behave much like a financial asset. On the other hand, for perishable agricultural goods this year's harvest may even be considered a different quality than last year's. These goods are then susceptible to more frequent stocks outs and price peaks. Oil tends to be somewhere in the middle; storage is costly in terms of infrastructure requirements and environmental risks and both demand and supply shocks can be significant.

Deaton and Laroque (1992, 1995) assumed deterministic demand and uncertain supply, with the advance in the 1995 model being the introduction of serially correlated supply shocks, which further complicates finding a solution to the model. In fact, the model can be switched to have known supply and uncertainty in demand, which again may be serially correlated. Divir and Roggoff (2010) solve such a model and apply the results to the oil market, employing the approach to characterize different "epochs"of oil. However, while such models appear to capture some essential features of the nature of commodity prices, they do not capture longer cycles, nor do they capture all of the persistence, especially within commodity booms. One of the features lacking in these models is investment. Adding investment may produce longer commodity cycles and greater persistence. However, adding investment to these models increases the state space, and good data on investment and production costs tend to be lacking, complicating an empirical analysis. While still not providing a full explanation for the nature of long-term commodity prices, this strand of the literature explains why commodity prices tend to exhibit high volatility and the existence of significant booms and busts and apparent and actual regime shifts depending on the nature of shocks.

A second strand of the literature has developed that addresses possible long-run trends in commodity prices. Prebisch (1951) and Singer (1950) both argued that commodity prices would be expected to decline relative to other, and especially manufactured goods', prices. One view was that as manufactured goods tend to be differentiated, oligpolistic products, and labor inputs to these goods are normally organized with unions demanding an increasing share of the cost to price mark-up, then the prices of these goods will tend to rise relative to homogeneous commodity inputs sold in competitive markets; see Kaldor (1987). A second view was simply that income elasticities of demand for manufactured goods and services would be higher than for basic commodities and hence, as consumers became more wealthy, commodity prices would tend to fall relative to other prices; see Bértola and Ocampo (2012) for a recent review and a discussion of the implications for one commodity-producing region: Latin America.

In 1988, Grilli and Yang (1988) provided a consistent dataset for a broad commodity index, which has been updated by Pfaffenzeller, Newbold and Rayner (2007)[3], and this prompted a third strand of the literature which has been the econometric analysis of this (and other) longer-term commodity price time series. The series graphed in Figure 3 illustrates that commodity prices have declined over the long term in relation to manufactured goods unit values (MUV's), but with significant volatility and punctuated by large booms and busts as predicted by commodity price theory. The econometric methodologies employed to analyze this price series have followed developments in econometric methodology more generally. As the literature started to consider whether economic time series had unit roots and/or time trends and whether multiple series were cointegrated or not (Nelson and Plosser (1982), Perron (1988), Stock and Watson (1988), Engle and Granger (1987)[4]), these types of techniques were then applied to the Grilli-Yang dataset. Subsequently, as tests became available to test for breaks in univariate[5] or multivariate contexts or specialist models became available, these were also applied to the commodity price time series.[6] In addition, tests of whether to include impulse dummies[7] or step dummies,[8] or a combination thereof[9] or a break[10]

[3] For recent updates, see the site: `www.stephan-pfaffenzeller.com/cpi.html`.

[4] For examples of this extensive literature, see: Banerjee and Hendry (1992), Banerjee, Dolado, Galbraith, and Hendry (1993), Chan and Wei (1988), Dickey and Fuller (1979, 1981), Hall and Heyde (1980), Hendry (1995), Johansen (1988, 1991, 1992a,b, 1995), Johansen and Juselius (1990, 1992), Phillips (1986, 1987, 1988), Park and Phillips (1988, 1989) Phillips and Perron (1988) and Stock (1987).

[5] See Perron (1990), Bunzel and Vogelsang (2005) and Harvey, Leybourne, and Taylor (2007).

[6] See Ardeni and Wright (1992) and Balagtas and Holt (2009).

[7] See originally Grilli and Yang (1988) did it and von Hagen (1989), Powell (1991), Cuddington (1992), León and Soto (1997) followed–sometimes with different dummies.

[8] As Cuddington and Urzúa (1989), Cuddington and Wei (1992), Ardeni and Wright (1992), Bleaney and Greenaway (1993) and Lutz (1999).

[9] See Newbold and Vougas (1996).

[10] See Perron (1990), Ocampo and Parra (2003) and Harvey, Leybourne, and Taylor (2012).

were also considered. Some papers like Perron (1989, 1990) proposed a test between a null (unit root) and an alternative (trend stationary) with the presence of a break under both hypotheses. But perhaps more importantly, this line of research has established that, under the alternative, an undetected break carries low power for unit root tests. Over the last two decades the research has evolved to develop new tests: to allow for several breaks; to find the breaks endogenously; and to remain stable with the magnitude of the shifts.[11]

A further strand of the literature considers the data itself. Some authors have attempted to consider the behavior of individual commodity prices, while others have considered alternative indices such as the industrial commodity-price index of *The Economist.*[12] In this paper we chose to work with the overall Grilli and Yang (1988) commodity price index (GYCPI), which has been employed widely in the literature. Moreover, our aim is to consider the behavior of commodity prices in general, as we wish to consider the potential implications for developing countries that may export several commodity products.

[11] We employ the following tests: Lanne, Lütkepohl, and Saikkonen (2002) [LLS], Perron and Yabu (2009) [PY], Carrion-i-Silvestre, Kim, and Perron (2009) [CKP] and Harvey, Leybourne, and Taylor (2013) [HLT]. Our contribution is not to develop further types of tests but rather to apply these relatively new techniques to the commodity data. For the LLS test we use JMulTi; for PY we use Gauss with the code provided by Pierre Perron and Tomoyoshi Yabu; for CKP test we use Gauss with the code provided by Josep Lluís Carrion-i-Silvestre and Pierre Perron: `http://people.bu.edu/perron/code.html`; and for HLT we use Gauss with the code provided by the authors at The Granger Centre for Time Series Econometrics at the University of Nottingham: `http://www.nottingham.ac.uk/economics/grangercentre/code.htm`.

[12] See, for example, Cashin and McDermott (2002) and Ocampo and Parra (2003).

3 Identifying Structural Shifts in the Deflated Commodity Price Index

We use the updated version of Grilli and Yang (1988) commodity price index (GYCPI) as developed by Pfaffenzeller, Newbold, and Rayner (2007) to analyze commodity prices. The GYCPI is a weighted average of 24 commodities. It does not include oil. The weights were calculated using the average export share from 1977 to 1979.[13] We construct an Index of Relative Commodity Prices by taking logarithms of the GYCPI and the manufacturing unit value index (MUV), and then compute their difference: $\text{LGYCPI}_t - \text{LMUV}_t$. The first step in our analysis was to search for outliers and ultimately for breaks. We use the Impulse Indicator Saturation (IIS) and the Step-Indicator Saturation (SIS) techniques to conveniently help us to detect data contamination and breaks in the data.[14] The user only has to specify the model and select a target size.[15] In the first stage, the IIS/SIS adds one impulse dummy for every observation, then it splits the dummies into blocks, regressing one block of dummies at a time against the data and keeping the significant dummies. It proceeds with the next block and so on until all periods are completed. Finally, the kept dummies in every block are combined and regressed again in recursive fashion; at the end only the significant dummies remain. This procedure is valid for both integrated and cointegrated data: see Johansen and Nielsen (2009).

Our methodology in employing the IIS and SIS techniques is as follows. We consider the Index of Relative Commodity Prices as the endogenous variable and a constant as the exogenous variable.[16] Using IIS we find 60 significant impulse dummies, and using SIS we find 18 significant step dummies for the whole period. Then we combine the two methods (SIS and IIS), as in Ericsson and Reisman (2012), and find 12 step dummies and 4 impulse dummies, all statistically significant. In each case and for the target sizes that we experimented–0.05, 0.025 or 0.01–the results turn out to be practically the same.[17] Doornik, Hendry, and Pretis (2013) find that SIS performs better than IIS in simulations, but there are some cases in which IIS does better–e.g., when there is a single impulse in the middle of an artificial sample. They also show that combining SIS and IIS does not have detrimental effects when there are step dummies, but it could reduce the power for the identification of the impulse indicators. Therefore, it is highly recommended to use the two techniques and the combination to improve the analysis.

[13] For more detail see Grilli and Yang (1988), Cuddington and Wei (1992), and Pfaffenzeller, Newbold, and Rayner (2007).

[14] See Hendry, Johansen, and Santos (2008); Santos (2008); Castle, Doornik, and Hendry (2012); Doornik, Hendry, and Pretis (2013).

[15] This accounts for the number of irrelevant variables that on average would survive. As one increases the size, the process is less likely to keep irrelevant variables but at the same time is more likely to exclude variables of relevance.

[16] The IIS technique identifies an unknown number of impulse dummies while the SIS technique identifies an unknown number of step dummies.

[17] The target size determines the statistical tolerance for the procedure. A target size of 0.05 for the IIS implies the acceptance on average of 5 impulse dummies that may not be in the data generating process for each 100 observations.

Figure 1 shows the fitted values of the grouped dummies for the three cases: IIS, SIS and SIS+IIS. One can see two major negative shifts: one in 1921 and the other in 1985. There are three upward steps in 1923-29, 1950-51 and 1973-74, and there are other more minor steps. We also considered a second methodology to employ the IIS+SIS technique, not reported here. In that case, we assumed no breaks and estimate an error correction model directly. We then apply IIS+SIS to the error correction term, which would be roughly equivalent to applying IIS+SIS to the error term of the long-run cointegrating equation. Clearly this methodology assumes cointegration and the assumption must then be that the breaks are not required to gain stationarity of the residuals. We prefer the first methodology, which makes no assumptions about the stationarity of the Index of Relative Commodity Prices. Still, we find similar results using both approaches.

A difficult question is how to classify the last four observations, from 2007 to 2010. The IIS/SIS methodology allows the researcher to consider if shifts are considered permanent or not. One possibility is that the recent 2007-2010 boom is transitory and that commodity prices will return to the pre 2007 “long-run”relationship which prevailed from 1921-1985. An alternative would be that there is a positive and permanent shift such that prices would not be expected to return to those levels or, to put that another way, that the lower prices from 1985-2007 actually represented a transitory negative break and not a permanent one. In favor of the first view is that, consistent with this hypothesis, in the 100 years plus of data that we have, there have been several positive booms that have all turned out to be transitory, perhaps explained by a Deaton and Laroque-type model, although with some booms exhibiting substantially greater persistence than typically found in their simulations. Interestingly some of these booms have been identified with positive dummies by the IIS/SIS methodology, and one has been almost as persistent as the current boom (namely 1923-29). But we are fully aware that in the end this is somewhat subjective and not determined by the statistical techniques. For now we assume that the 1921 and 1985 shifts are permanent in our model specifications, but we will return to this question when we consider the size of the current boom and the implications for future commodity prices in Section 5.[18]

[18] While we do not claim to be clairvoyants, other tests do lead us to conclude that the 1985 shift might be permanent. For example, we employed similar saturation tests but with step dummies using *Autometrics*; a Bai-Perron test (Bai and Perron, 1998, 2003), and an estimate of a Markov switching regression–switching the intercept, using 3 regimes, uniform probabilities and switching residual variance. These tests appear to be consistent in suggesting that the 1985 shift might be permanent.

4 Developing Vector Error Correction Models for the Long-Run Commodity Price

In this section we consider a set of models to describe long-run commodity prices. Since Raul Prebisch and Hans Singer postulated in the early 1950s that real commodity prices are in a long-run secular decline, there has been particular interest in considering whether a trend model of commodity prices is a reasonable characterization. At the same time, it is clear that in more than 100 years of data there are likely to have been several structural breaks provoked by shifts in the cost function of production, innovation in substitutes or other developments. In the previous section the extremely general IIS/SIS methodology was used to endogenously test the number and duration of structural breaks. In this section we then combine the two issues and specifically test whether, given these identified breaks, a trend model or other models are validated. We then develop the error correction interpretation of the model. More specifically we investigate three hypotheses:

- Taking into account the endogenously identified breaks, are commodity prices and manufactured units values cointegrated?
- If so, does the linear combination (of LGYCPI_t and LMUV_t) remove the stochastic trend but not the deterministic trend, such that the commodity price ratio is then trend-stationary (down-trending accordingly to Prebisch and Singer).
- Is the appropriate linear combination cointegrated with a cointegrating vector of unity? If so then, does this imply that the simple real commodity index is stationary (or trend-stationary)?

In addition, and assuming we do establish cointegration, we are interested in the average speed that commodity prices return to the identified long-run equilibrium, or the half-life of the error correction process. In order to investigate these issues, we estimated three vector error correction (VEC) models that include unrestricted deterministic impulse and break dummies (i.e., that may lie outside the cointegration space). We refer to the first model as the Restricted Model, as this has no linear deterministic trend in the cointegration space nor in the the levels: we only include an intercept in the cointegration space. The second model, which we call the Unrestricted Model, also has no linear trend in the cointegration space but, unlike the Restricted Model, it assumes a non-zero growth rate in the data: the intercept lies outside the cointegration space (unrestricted constant). Finally, we estimated a Trend Model which includes a linear trend inside the cointegration space and is then consistent with the Prebisch-Singer hypothesis. We subjected the three models to a battery of specification tests to determine which model might be considered a more valid description of the data. The three VEC models may be written as follows:

$$\Delta \mathbf{Y}_t = \boldsymbol{\alpha}\mu_0 + \boldsymbol{\alpha}\boldsymbol{\beta}' \mathbf{Y}_{t-1} + \boldsymbol{\Gamma}(L)\Delta \mathbf{Y}_t + \boldsymbol{\delta}_3 \mathbf{D}_t + \mathbf{u}_t \qquad \text{(Restricted Model)}$$

$$\Delta \mathbf{Y}_t = \boldsymbol{\delta}_0 + \boldsymbol{\alpha}\boldsymbol{\beta}' \mathbf{Y}_{t-1} + \boldsymbol{\Gamma}(L)\Delta \mathbf{Y}_t + \boldsymbol{\delta}_3 \mathbf{D}_t + \mathbf{u}_t \qquad \text{(Unrestricted Model)}$$

$$\Delta \mathbf{Y}_t = \boldsymbol{\delta}_0 + \boldsymbol{\alpha}\mu_1 t + \boldsymbol{\alpha}\boldsymbol{\beta}' \mathbf{Y}_{t-1} + \boldsymbol{\Gamma}(L)\Delta \mathbf{Y}_t + \boldsymbol{\delta}_3 \mathbf{D}_t + \mathbf{u}_t \qquad \text{(Trend Model)}$$

where $\mathbf{Y}$ is a (2×1) vector of the log of the commodity price and the manufactured unit value index, $\mathbf{D}$ is a (6×1) vector of impulse and step dummies as identified in the previous section, t is a trend and $\mathbf{u}_t \sim \mathsf{IN}_2[\mathbf{0}, \boldsymbol{\Omega}]$. Then we have a set of parameters $(\boldsymbol{\delta}_0, \boldsymbol{\delta}_3, \boldsymbol{\Gamma})$ separate from the cointegration space; and a set of parameters that are within the cointegration space: $(\mu_0, \mu_1, \boldsymbol{\alpha}, \boldsymbol{\beta})$.[19]

Given a cointegration relation, the three models assume different processes for the data. The Restricted Model assumes that there is no linear deterministic trends in the series as $\mathsf{E}[\Delta \mathbf{Y}_t | \mathbf{D}_t = \mathbf{0}] = \mathbf{0}$, but that there is a restricted constant; the long-run equilibrium may then have a non-zero mean. This would imply, first, that neither the commodity prices nor manufactures unit values have a deterministic trend; and second, that in the long run, the ratio of commodities to manufactures, as estimated in our models, has an equilibrium mean different from zero (μ_0). On the other hand, the Unrestricted Model assumes that $\mathsf{E}[\Delta \mathbf{Y}_t | \mathbf{D}_t = \mathbf{0}] = \boldsymbol{\delta}_0 \neq \mathbf{0}$ and $\mu_0 = 0$; that is, each variable has a deterministic trend in its level but these trends cancel each other out in long-run relationship.[20] Finally, the Trend Model includes a deterministic trend in the cointegration relationship and, like the Unrestricted Model, each variable has a deterministic trend in its level but these trends do not cancel each other in the long-run relationships, implying a trend-stationary relationship between the variables.[21] In this setup, the long run ratio of commodities to manufactures, will decline (or rise) permanently by a factor of μ_1. In this case a decline will be consistent with the Prebish-Singer hypothesis.

On Model Specification and Results

In order to estimate a model, we first consider the appropriate lag structure in an unrestricted VAR representation of each model following Johansen (1995) and Juselius (2006). The information criteria (SCB, HQ, AIC), Lagrange multiplier (LM) and Wald tests[22] suggest that we need two lags in the levels, i.e., one lag in the VEC in differences for each model. Table 3 shows that adding one

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[19] Note that $\boldsymbol{\Gamma}_0 = \mathbf{0}$, and $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$ under the hypothesis of cointegration. For a detailed exposition of the all the different cases see Johansen (1994), Hendry and Juselius (2001) or Juselius (2006).

[20] One can perfectly test whether $\boldsymbol{\delta}_0 = \boldsymbol{\alpha}\mu_0$.

[21] But as Hendry and Juselius (2001) note, this could also happen if one has a single trend-stationary variable instead of an equilibrium relationship. We will come back to this point later.

[22] One normally uses the Wald test when he or she “knows”the standard errors.

or two additional lags will harm the fit of the models given the results of the LM test. On the other hand, for the Wald test, we reject other lag structures in favor of two lags for all the models we consider.

Simultaneously we considered two types of likelihood ratio tests to establish the rank of the Π matrix.[23] Under the hypothesis of cointegration, $\Pi = \alpha\beta'$. The most important thing to note about these tests is that the distribution under the hypothesis of cointegration depends on the lagged differences and other terms included, i.e., constant, trend, dummies. Johansen's test is the standard and common practice given this setup. The main difference between Johansen's test and Saikkonen and Lütkepohl's (S&L) is that S&L's test uses a two-step procedure. It first estimates the deterministic terms by GLS and then it runs the standard tests. Although the power of S&L's test is superior in specific cases (Saikkonen and Lütkepohl, 1999), it is not superior in all cases (Lütkepohl and Krätzig, 2004).

Table 2 shows the results of the cointegration tests. Interestingly, we find only one significant long-run relationship for the Restricted and the Unrestricted models, and no significant long-run relationship in the Trend Model, with or without step dummies.[24] The step dummies increase the test value and thus the significance of the tests for all models, but only for the Restricted and Unrestricted specifications do the values exceed the 90 percent significance level. The tests on cointegration with a trend specification remain insignificant, even when we include dummy variables that we found using the IIS/SIS method on the real price index with a predetermined fixed trend. Although we do not report the results, we also found that the recursive eigenvalues of the Restricted and Unrestricted models are fairly stable over time, while those of the Trend model decrease as the sample increases. These results and other specification tests shown in Table 4 indicate that the Trend model should be rejected in favor of either the Restricted or the Unrestricted model.[25]

Finally, before estimating the VEC, we need to decide which dummies would be in the cointegration space and which should not. The deterministic shifts that we include in all models come from our first estimations of the Impulse Indicator Saturation (IIS) and Step-Indicator Saturation (SIS) regressions on the index of relative prices. We included other impulse dummies in the following years: 1921, 1938, 1975, 1986.[26] We tested each dummy individually and together, inside

[23] See the Johansen test in Johansen (1995) and Saikkonen and Lütkepohl in Lütkepohl and Saikkonen (2000), Saikkonen and Lütkepohl (2000a,b), Saikkonen and Luukkonen (1997). We did most of the estimations in Anders Warne's Structural VAR (update November 11, 2011) and JMulTi, and our reference to this software is in Lütkepohl and Krätzig (2004).

[24] We define the step dummy, s1921 = 1 if year $\geq$ 1921 and zero otherwise.

[25] The final choice can be done by a consistent procedure as Johansen (1995) and Doornik, Hendry, and Nielsen (1998) suggest. This procedure would select the Restricted over the Unrestricted but the fact that $\delta_0 = \alpha\mu_0$ statistically implies that one can use either (see Table 4).

[26] These dummies represent outliers in the series. Most of them were previously identified in the literature; see von Hagen (1989), Powell (1991), Newbold and Vougas (1996), Lutz (1999). The 1986 dummy was not identified

and outside the cointegration space using LR tests. The tests suggest that most of the dummies should be included outside the cointegration space (see Table 5 and Table 6). Although the step dummy `s1985` is individually insignificant at standard levels in the Restricted and the Unrestricted models, it is significant when tested with the other dummies; see Table 5. We suspect this happens because, in the recent boom, observations are considerably higher than in previous ones. When we estimate the model to 2007, we find this dummy individually significant at standard levels. Then, in Table 6, the `i1986` dummy is statistically significant and clearly is affecting the overall results, but when we estimated the models and included it inside the cointegration space we find that it is actually statistically insignificant.

Model Results

All the coefficients estimated inside the cointegration relation, either in the Restricted and Unrestricted models, have the value and sign that one would expect and are statistically significant (see Table 7). Moreover, the coefficients on $LMUV_t$ are recursively stable and normally distributed around minus one (see Figure 4). When we restrict the long-run coefficients on $LMUV_t$ to -1, the test shows that this restriction is statistically plausible and stable over time (see panels (a) and (b) in Figure 6).[27] On the other hand, the Trend model has the expected sign on the $LMUV_t$ coefficient but not on the trend; instead, the trend appears to be negative within the cointegrating space (this implies real commodity prices would increase over time–see Table 7). Besides, the recursive estimates of the trend coefficient appear to be unstable (with upward shifts) and right-biased (see Figure 4)–although not different than minus one, as the restriction test indicate (see panel (b) in Figure 6). Figure 5 finally shows the recursive and bootstrapped coefficient values for the constant and the trend in the Restricted and Trend models, respectively. One can see that the Restricted model has a stable and non-zero coefficient for the constant term, while the Trend model has a relatively stable coefficient (negative) and not statistically different from zero. The restriction tests in Figure 6 show that the Restricted model rejects the constant as being zero–particularly taking the more recent observations–and the $LMUV_t$ coefficient as being minus one simultaneously, but the Trend model fails to reject the trend as being zero–except in very few years–and both the trend as being zero and the $LMUV_t$ coefficient as being minus one simultaneously.

All the tests indicate that a preferred model would be one where, taking into account the endogenously identified breaks and shifts in the series, commodity prices and manufactured unit values are cointegrated and where the cointegrating vector may be equal to unity and where there

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previously; however, we think this dummy reflects significant changes in international trade policies around 1986 as agreed on in the General Agreement on Tariffs and Trade (GATT), which may account for this break, as may the large currency depreciations of developing countries.

[27] To save space we omit the figure on the LR test for the Unrestricted model, but it looks almost like the Restricted model.

is no deterministic trend component. In this sense the results do not favor a strict interpretation of the Prebisch-Singer hypothesis of a stable declining commodity terms of trade trend. That said, it is still clear that commodity prices have fallen against manufactured unit values and that all of the permanent shifts in the relationship have been negative. Naturally, this raises interesting questions regarding the nature of the current boom, and we come back to this point in the next section below.[28]

The estimates on the short-run dynamics are very similar across models. First, in the three models, the coefficient on $\Delta LMUV_{t-1}$ in the $\Delta LGYCPI_t$ equation that links the growth rate of unit value of manufactures with the change in commodity prices is negative but insignificant, suggesting no influence in the short run between these variables. Nevertheless, the coefficient on $\Delta LGYCPI_{t-1}$ in both the $\Delta LGYCPI_t$ and the $\Delta LMUV_t$ equations is positive and significant with a value of around 0.5 on $\Delta LGYCPI_t$ and 0.3 on $\Delta LMUV_t$. That is, the growth rate of commodity prices has a larger influence in the short run in both equations and in all three models.

The dummy variables also exhibit a similar pattern across models. For example, the coefficientd on the step dummies of 1921 and 1985 are very close in value within the models and among the three different specifications, and only significant in the equation for the change in commodity prices. Accordingly, the other impulse dummies have negative and significant effects on the commodity price change, with the only exception being the dummy of 1986, which has no influence on the commodity price changes; instead, it is positive and significant for the change in manufactures unit value.

Finally the feedback coefficients have similar values and significance across models. The feedback coefficient on the $\Delta LGYCPI_t$ equations is negative and significant in all three models, in accordance with a standard error-correction mechanism. But then the coefficients on the $\Delta LMUV_t$ equations are insignificant, so that when we tested for weak exogenity, we find that the $\Delta LMUV_t$ equations are weakly exogenous in all three models (see Table 8). This implies that, while changes in manufactures unit values drive changes in commodity prices, there is little evidence in favor of the reverse causality.

[28] As mentioned above, we also employed the IIS/SIS methodology in a very different manner, estimating first the VEC and then considering whether there were breaks and shifts in the error correction term. This second methodology presupposes cointegration even in the absence of the relevant breaks. Even so, we found results similar to those reported here.

5 The Current Boom in Commodity Prices and Implications

To consider the statistics of the current boom in commodity prices in relation to historical commodity price booms, we use the estimated model and in particular the estimated real commodity price in each regime (the period between the identified permanent shifts in real commodity prices) to compare various statistics regarding the current and previous booms relative to the mean real price of each regime. As an illustration of the procedure, Figure 3 plots the deflated Grilli-Yang commodity price index and indicates the mean real level of each regime and the boom periods that we consider. As can be seen, we have just three regimes over the 100-plus years of data, and we work with six boom periods including the current one. Table 11 compares various statistics across these six booms. The longest boom prior to the current one is that of 1924-1928, and the largest boom in terms of the distance of the mean actual price during the boom from the mean of the regime is that of 1974. However, even with the dataset ending in 2010 the current boom is by far the largest in the sense of the integral of the distance between the actual commodity price and the mean of the regime over the boom period–the total boom size as detailed in the Table 11.

In this sense, the current boom does indeed merit the name of a "super boom" or "super cycle". The next question we investigate is, then, what would be the expected movement of commodity prices if this boom is expected to be temporary in nature and prices were expected to return to the regime mean? Assuming prices followed our preferred model specification above, this implies that prices would return roughly to the regime mean at the same speed as in the sample, as specified by the coefficient on the error correction term corrected for any other short-run dynamics. As prices at the end of the dataset are considerably higher than the regime mean, this implies that real prices would fall quite substantially. As we find in Figure 7, commodity prices respond to MUVs, but essentially MUVs do not respond to commodity prices, thus almost all of the movement would be expected to be in commodity prices. This relationship notwithstanding, forecasting the VEC forward does imply a steadily increasing MUV index in line with historical levels of inflation, which limits the downward fall in nominal commodity prices. Figure 2 below illustrates the expected movement in real commodity prices. Given our estimates of the speed of the error correction, it turns out that prices would be expected to have fallen by more than 50 percent of the distance to the regime mean by 2014 and 75 percent by 2018.

Of course, these forecasts are critically dependent on the assumption that no new permanent shift in commodity prices has occurred. Interestingly, in 100-plus years of data all positive booms have been temporary in nature, while we have identified only two negative permanent shifts. Indeed, these shifts have followed particularly striking booms, particularly the boom leading up to the shift in the early 1920s and the boom of the 1970s that preceded a negative shift in the 1980s. As suggested in Powell (1991), there may be incentives for greater investment and innovation that may then serve to reduce productions costs in these boom periods, which then alter the new long-run equilibrium going forward. On the other hand, over the last 100 years, the world has not seen an event quite the same as the transformation of the Chinese economy with its globally systemic impacts. Whether this may result in a permanent positive shift or simply a long but eventually temporary boom remains to be seen.

The answer to that question is critical for a number of commodity-dependent countries. In particular a number of countries are highly dependent on commodities for exports and fiscal revenues in Latin America and the Caribbean. In the Table 12 we present estimates of the potential impact of a fall in commodity prices on the regime mean for this group of countries. The methodology is as follows. We employ a country-specific commodity price index developed by Fernández-Arias and Pérez and employed in the IDB 2013 Latin American and Caribbean Macroeconomic Report[29] and regress each country index on the general Grilli-Yang index for the period that data are available for both. We then take the regression coefficient, which is akin to a beta coefficient (in the Capital Asset Pricing Model), as the sensitivity of the country index to a general movement in commodity prices. We then apply this estimated beta to our forecasts of the general index in order to obtain a forecast for the country-specific index. We then also use the elasticities developed by Fernández-Arias and Pérez regarding the effect of a change in the country index on fiscal revenues to calculate the fiscal impact of a fall in commodity prices on commodity-dependent countries in the region. The results are detailed in Table 12 below.

Table 12 also illustrates that a fall of commodity prices back to the regime mean would have a very significant impact on fiscal revenues for this set of countries, particularly for some of the oil exporters such as Venezuela, Trinidad and Tobago and Ecuador, but also for Bolivia, Peru, Chile, Colombia and Mexico. Argentina is less affected, as its fiscal revenues are less commodity dependent relative to some of the other countries.

These results illustrate the need to consider explicitly how to manage the risk of commodity prices to fiscal sustainability. The data and analysis in this paper illustrate that commodity prices have been extremely volatile for over 100 years with sharp boom and bust cycles. If commodity prices do not revert to the previous regime mean, this will be the first time that that has happened in the available sample. In all other cases, positive booms have been temporary in nature, albeit prolonged in some cases. It is also likely that the costs of mispredicting a boom as permanent when in fact it turns out to be temporary are higher than the costs of mispredicting a boom as temporary when in fact it turns out to be permanent, or at least it would be advised for policymakers to take a conservative view for several years. Given the current context of potentially rising world interest rates and a gradual deceleration in the Chinese economy, as well as recent falls in commodity prices, there is indeed increasing concern regarding potential future declines.

[29] See Powell (2013), Chapter 4.

Our results also carry implications for the design of programs to hedge risks such as the use of futures, options and swaps and the design of instruments to manage risks such as stabilization funds. While we leave to future research the precise design parameters that would be suggested by our results, suffice it to say that countries would be well-advised to consider both sets of instruments and devise a program to manage risks that is appropriate to their particular commodity portfolio taking into account the historical patterns of booms and busts and the speed at which prices may return to a long-run equilibrium, while also considering the possibility of a permanent shift in prices.

6 Conclusions

In this paper we reconsidered the nature of long-term commodity prices, and in particular the validity of the Prebisch-Singer hypothesis, employing new statistical techniques (namely IIS/SIS) to identify potential breaks and permanent shifts in the series. This technique is a very general methodology to endogenously identify breaks anywhere in the time series and of any duration. Our results indicate that, while commodity prices have declined with respect to manufactured unit values, there is little evidence for a stable declining terms of trade trend. Rather, when breaks are carefully accounted for, commodity prices and manufactured goods prices appear to be cointegrated with a cointegrating vector of unity.

These results then allowed us to estimate a preferred vector error correction model, and we find that while commodity prices respond to changes in manufactured unit values there is little evidence that manufactured unit values respond to commodity prices. Hence when there are shocks such that commodity prices are away from long-run equilibrium values, normally it is commodity prices that move to re-establish the long run relationship. Given the results from the estimated VEC model, we were also able to estimate the speed at which a commodity price shock tends to revert to the long-run level.

We employed the estimated results to consider the characteristics of the current commodity boom relative to previous booms. The current boom does indeed merit the name of a super boom or super cycle given its duration and size, although previous booms have experienced higher prices in terms of the distance from the mean actual price level from the regime mean. Assuming no permanent shift in commodity prices has occurred, real commodity prices would be expected to fall to the previous regime mean, implying substantial real commodity prices declines, although nominal prices declines would be limited by an expected rise in MUVs consistent with historical inflation rates. However, it is clearly hard to interpret the last data points as to whether they point to a temporary but prolonged boom or potentially a permanent shift. The fact that all previous booms over the last 100-plus years of available data all turned out to be temporary (albeit some prolonged) may lead analysts and policymakers to be somewhat conservative regarding the nature of the current boom.

Moreover, an analysis of commodity-dependent countries in Latin America and the Caribbean indicates that some countries would suffer substantial falls in fiscal revenues if commodity prices did fall back to the previous regime mean. This suggests that countries in this position should seek means to hedge or manage risks and should follow prudent fiscal strategies that explicitly recognize these risks.

Table 1. Unit Root Tests

Test	Variable Setup	Deterministic Terms	LGYCPI	Break(s)	LMUV	Break(s)	INDEX	Break(s)
ADF	levels	constant	-0.62		-0.31		-2.18	
	levels	constant and trend	-2.68		-2.34		-3.22 *	
	1st difference	constant	-8.28 ***		-6.74 ***		-8.74 ***	
ZA	levels	trend with 1 break	-3.93	1933	-3.21	1932	-3.33	1932
	levels	intercept and trend with 1 break	-4.55	1921	-3.82	1973	-3.99	1990
	1st difference	intercept with 1 break	-7.87 ***	1921	-4.47	1940	-8.98 ***	1994
LLS	levels	constant and step dummy	-0.47	1921	-0.44	1951	-2.60 *	1921
	levels	trend and step dummy	-3.55 ***	1921	-2.02	1951	-2.08	1921
	1st difference	constant impulse dummy	-7.72 ***	1921	-6.36 ***	1951	-8.86 ***	1921
PY	levels	trend with 1 break	-0.10	1937	0.00	1940	-0.18	1975
	levels	intercept and trend with 1 break	12.74 ***	1930	1.93	1974	8.42 ***	1986
	1st difference	intercept with 1 break	-0.03	1932	0.47	1932	-0.10	1993
CKP	levels	trend with 0 breaks	9.12		16.46		5.69 *	
	levels	trend with 1 break	8.15	1921	9.84	1940	8.48	1918
	levels	trend with 2 breaks	15.01	1921, 1932	26.43	1940, 1973	11.49	1918, 1946
	1st difference	constant with 0 breaks	0.71 **		2.36		0.67 **	
HLT	levels	trend with 1 break	-2.86		-2.65		-3.34	
	levels	trend with 2 breaks	-3.21		-3.15		-4.03	
	1st difference	trend with 1 break	-8.20 ***		-6.78 ***		-9.93 ***	
	1st difference	trend with 2 breaks	-8.41 ***		-6.86 ***		-10.07 ***	

Variables: Logs of index of relative prices (GYCPI) and the manufacturing unit value index (MUV), then compute the real INDEX = LGYCPI - LMUV. Tests: Dickey and Fuller (1979) [ADF], Zivot and Andrews (1992) [ZA], Lanne, Lütkepohl, and Saikkonen (2002) [LLS], Perron and Yabu (2009) [PY], Carrion-i-Silvestre, Kim, and Perron (2009) [CKP] and Harvey, Leybourne, and Taylor (2013) [HLT]. The criteria to select the lags was based on Schwarz Criterion with maximum lags set to the following rule: $\text{maxk} = \lfloor 12\left((n/100)^{1/4}\right)\rfloor$. When applicable we use a trim 0.15. Significance: * at 10%, ** at 5% and *** at 1%.

Table 2. Cointegration Tests

Test	No. of lagged differences	Null hypothesis	Test value	Critical values 90%	95%
		Restricted Model (intercept included)			
Johansen Test	2	$r = 0$	16.29	17.98	20.16
		$r = 1$	6.05	7.60	9.14
S&L Test	2	$r = 0$	12.53	10.47	12.26
		$r = 1$	2.55	2.98	4.13
S&L Test	2	$r = 0$	12.91	10.47	12.26
(step dummies[1])		$r = 1$	2.16	2.98	4.13
		Unrestricted Model (intercept excluded)			
Johansen Test	2	$r = 0$	10.18	13.42	15.41
S&L Test	2	$r = 0$	9.53	8.18	9.84
S&L Test	2	$r = 0$	10.80	8.18	9.84
(step dummies[1])					
		Trend Model (trend and intercept included)			
Johansen Test	2	$r = 0$	15.89	23.32	25.73
		$r = 1$	5.15	10.68	12.45
S&L Test	2	$r = 0$	6.34	13.88	15.76
		$r = 1$	2.04	5.47	6.79
S&L Test	2	$r = 0$	9.09	13.88	15.76
(step dummies[1])		$r = 1$	2.44	5.47	6.79
S&L Test	2	$r = 0$	10.48	13.88	15.76
(step dummies[2])		$r = 1$	2.24	5.47	6.79

Note: critical values are taken form Johansen (1995) and Lütkepohl and Saikkonen (2000), Saikkonen and Lütkepohl (2000a,b).

[1] included dummies using the IIS method. The dummies are: s1921, s1985, i1921, i1938, i1975, i1986. We define, for example, the step dummy s1921 = 1 if year $\geq$ 1921 and zero otherwise.

[2] included dummies using the IIS/SIS methods but with a predetermined fixed trend. The dummies are: i1917, i1921, i1932, i1974, i1999, i2010, s9293, s0708.

Table 3. Lag Order Tests

Restricted Model			
Null hypothesis:	2 lags	1 lags	2 lags
Alternative hypothesis:	3 lags	2 lags	4 lags
Statistic	LM(4)	Wald(4)	LM(8)
Statistic value	2.815	60.626	4.085
p-value	0.589	0.000	0.849
Unrestricted Model			
Null hypothesis:	2 lags	1 lags	2 lags
Alternative hypothesis:	3 lags	2 lags	4 lags
Statistic	LM(4)	Wald(4)	LM(8)
Statistic value	3.387	54.196	4.651
p-value	0.495	0.000	0.794
Trend Model			
Null hypothesis:	2 lags	1 lags	2 lags
Alternative hypothesis:	3 lags	2 lags	4 lags
Statistic	LM(4)	Wald(4)	LM(8)
Statistic value	3.159	56.309	4.419
p-value	0.532	0.000	0.817

Note: the estimation was done for with the series in levels.

Table 4. LR-Type Specification Tests

Restricted Model	
H_0 :	$\mu_0 = 0$
H_a :	$\mu_0 \neq 0$
Rank	1
LR-Statistic	5.699
Degrees of Freedom	1
p-value	0.017
Unrestricted Model	
H_0 :	$\delta_0 = \alpha\mu_0$
H_a :	$\delta_0 \neq \alpha\mu_0$
Rank	1
LR-Statistic	2.681
Degrees of Freedom	1
p-value	0.102
Trend Model	
H_0 :	$\mu_1 = 0$
H_a :	$\mu_1 \neq 0$
Rank	1
LR-Statistic	0.093
Degrees of Freedom	1
p-value	0.761

Table 5. Restrictions on Deterministic Variables

Dummy	LR-Test	df	p-value
	Restricted Model		
s1921	8.475	2	0.014
s1985	5.692	2	0.058
i1921	27.214	2	0.000
i1938	11.521	2	0.003
i1975	6.892	2	0.032
i1986	14.853	2	0.001
Overall	74.919	12	0.000
	Unrestricted Model		
s1921	6.232	2	0.044
s1985	5.387	2	0.068
i1921	27.470	2	0.000
i1938	11.504	2	0.003
i1975	6.872	2	0.032
i1986	14.723	2	0.001
Overall	71.600	12	0.000
	Trend Model		
s1921	3.934	2	0.140
s1985	5.332	2	0.070
i1921	22.799	2	0.000
i1938	11.597	2	0.003
i1975	6.933	2	0.031
i1986	14.752	2	0.001
Overall	71.081	12	0.000

Note: the null hypothesis is that the dummies have zero coefficients.

Table 6. Long Run Restrictions on Deterministic Variables

Dummy	LR-Test	df	p-value
	Restricted Model		
s1921	2.321	1	0.128
s1985	0.544	1	0.461
i1921	0.569	1	0.451
i1938	0.442	1	0.506
i1975	0.091	1	0.763
i1986	8.267	1	0.004
Overall	21.433	6	0.044
	Unrestricted Model		
s1921	0.238	1	0.625
s1985	0.563	1	0.453
i1921	1.251	1	0.263
i1938	0.191	1	0.662
i1975	0.014	1	0.907
i1986	8.044	1	0.005
Overall	18.593	6	0.099
	Trend Model		
s1921	0.251	1	0.617
s1985	0.552	1	0.457
i1921	0.476	1	0.490
i1938	0.283	1	0.595
i1975	0.067	1	0.796
i1986	5.611	1	0.018
Overall	18.553	6	0.100

Note: the null hypothesis is that dummies have zero long run effect on all endogenous variables.

Table 7. Main Estimations for the Cointegrated Models

Variables	Restricted Model	Unrestricted Model	Trend Model
LGYCPI	1	1	1
LMUV	-0.971	-0.968	-0.914
	(0.067)	(0.066)	(0.144)
Constant	-0.550	–	–
	(0.193)		
Trend	–	–	-0.002
			(0.004)
	Feedback Coefficients for Rank 1		
α_{11}	-0.332	-0.318	-0.319
	(0.092)	(0.094)	(0.090)
α_{21}	-0.062	-0.044	-0.056
	(0.051)	(0.052)	(0.050)

Note: We estimated the models by maximum likelihood, see Johansen (1995). The standard errors are shown in parentheses.

Restricted Model

$$\begin{bmatrix}\Delta LGYCPI_t\\ \Delta LMUV_t\end{bmatrix} = \begin{bmatrix}\underset{(0.092)}{-0.332}\\ \underset{(0.051)}{-0.062}\end{bmatrix}\left(LGYCPI_{t-1} \;-\; \underset{(0.067)}{0.971}\,LMUV_{t-1} \;-\; \underset{(0.193)}{0.550}\right) + \begin{bmatrix}\underset{(0.104)}{0.492} & \underset{(0.175)}{-0.182}\\ \underset{(0.058)}{0.272} & \underset{(0.097)}{0.158}\end{bmatrix}\begin{bmatrix}\Delta LGYCPI_{t-1}\\ \Delta LMUV_{t-1}\end{bmatrix} +$$

$$\begin{bmatrix}\underset{(0.040)}{-0.103} & \underset{(0.042)}{-0.131} & \underset{(0.112)}{-0.620} & \underset{(0.109)}{-0.298} & \underset{(0.114)}{-0.259} & \underset{(0.112)}{0.067}\\ \underset{(0.022)}{-0.008} & \underset{(0.023)}{-0.033} & \underset{(0.062)}{-0.195} & \underset{(0.061)}{-0.007} & \underset{(0.063)}{-0.027} & \underset{(0.062)}{0.210}\end{bmatrix}\begin{bmatrix}S1921_t\\ S1985_t\\ I1921_t\\ I1938_t\\ I1975_t\\ I1986_t\end{bmatrix} + \begin{bmatrix}u_{1,t}\\ u_{2,t}\end{bmatrix}$$

Unrestricted Model

$$\begin{bmatrix}\Delta LGYCPI_t\\ \Delta LMUV_t\end{bmatrix} = \begin{bmatrix}\underset{(0.094)}{-0.318}\\ \underset{(0.052)}{-0.044}\end{bmatrix}\left(LGYCPI_{t-1} \;-\; \underset{(0.066)}{0.968}\,LMUV_{t-1}\right) + \begin{bmatrix}\underset{(0.104)}{0.486} & \underset{(0.176)}{-0.194}\\ \underset{(0.057)}{0.265} & \underset{(0.097)}{0.139}\end{bmatrix}\begin{bmatrix}\Delta LGYCPI_{t-1}\\ \Delta LMUV_{t-1}\end{bmatrix} +$$

$$\begin{bmatrix}\underset{(0.053)}{0.194} & \underset{(0.043)}{-0.112} & \underset{(0.042)}{-0.125} & \underset{(0.112)}{-0.623} & \underset{(0.109)}{-0.300} & \underset{(0.114)}{-0.259} & \underset{(0.112)}{0.063}\\ \underset{(0.029)}{0.046} & \underset{(0.024)}{-0.022} & \underset{(0.023)}{-0.026} & \underset{(0.062)}{-0.200} & \underset{(0.060)}{-0.009} & \underset{(0.063)}{-0.027} & \underset{(0.062)}{0.205}\end{bmatrix}\begin{bmatrix}Const_t\\ S1921_t\\ S1985_t\\ I1921_t\\ I1938_t\\ I1975_t\\ I1986_t\end{bmatrix} + \begin{bmatrix}u_{1,t}\\ u_{2,t}\end{bmatrix}$$

Trend Model

$$\begin{bmatrix}\Delta LGYCPI_t\\ \Delta LMUV_t\end{bmatrix} = \begin{bmatrix}\underset{(0.090)}{-0.319}\\ \underset{(0.050)}{-0.056}\end{bmatrix}\left(LGYCPI_{t-1} \;-\; \underset{(0.144)}{0.914}\,LMUV_{t-1} \;-\; \underset{(0.004)}{0.002}\,t\right) + \begin{bmatrix}\underset{(0.103)}{0.479} & \underset{(0.175)}{-0.192}\\ \underset{(0.057)}{0.268} & \underset{(0.097)}{0.139}\end{bmatrix}\begin{bmatrix}\Delta LGYCPI_{t-1}\\ \Delta LMUV_{t-1}\end{bmatrix} +$$

$$\begin{bmatrix}\underset{(0.061)}{0.235} & \underset{(0.045)}{-0.126} & \underset{(0.041)}{-0.126} & \underset{(0.112)}{-0.605} & \underset{(0.109)}{-0.302} & \underset{(0.113)}{-0.258} & \underset{(0.111)}{0.063}\\ \underset{(0.034)}{0.060} & \underset{(0.025)}{-0.028} & \underset{(0.023)}{-0.031} & \underset{(0.062)}{-0.193} & \underset{(0.060)}{-0.008} & \underset{(0.062)}{-0.024} & \underset{(0.061)}{0.208}\end{bmatrix}\begin{bmatrix}Const_t\\ S1921_t\\ S1985_t\\ I1921_t\\ I1938_t\\ I1975_t\\ I1986_t\end{bmatrix} + \begin{bmatrix}u_{1,t}\\ u_{2,t}\end{bmatrix}$$

Table 8. Weak Exogenity Test

Equation	df	Wald Test	p-value
		Restricted Model	
LGYCPI	1	13.189	0.000
LMUV	1	1.488	0.223
		Unrestricted Model	
LGYCPI	1	11.551	0.001
LMUV	1	0.715	0.398
		Trend Model	
LGYCPI	1	12.589	0.000
LMUV	1	1.259	0.262

Note: the null hypothesis is that alpha in the given equation is zero.

Table 9. Diagnostic Tests (I)

Test	Statistic	Distribution	p-value
		Restricted Model	
Portmanteau	15.6573	$\chi^2(14)$	0.3347
LM Test	25.7465	$\chi^2(20)$	0.1743
Normality Test[1]	10.1400	$\chi^2(4)$	0.0381
Normality Test[2]	5.8302	$\chi^2(4)$	0.2122
Multi ARCH LM	74.1568	$\chi^2(45)$	0.0040
		Unrestricted Model	
Portmanteau	15.2961	$\chi^2(14)$	0.3582
LM Test	24.9038	$\chi^2(20)$	0.2051
Normality Test[1]	10.6703	$\chi^2(4)$	0.0305
Normality Test[2]	6.5362	$\chi^2(4)$	0.1625
Multi ARCH LM	68.4346	$\chi^2(45)$	0.0137
		Trend Model	
Portmanteau	15.2918	$\chi^2(14)$	0.3585
LM Test	25.5309	$\chi^2(20)$	0.1819
Normality Test[1]	9.8362	$\chi^2(4)$	0.0433
Normality Test[2]	6.2237	$\chi^2(4)$	0.1831
Multi ARCH LM	67.7351	$\chi^2(45)$	0.0158

Note: all test were done taking 5 lags.

/1 Doornik and Hansen (2008).

/2 Lütkepohl (1993), p. 153.

Table 10. Diagnostic Tests (II)

Test	Statistic	Distribution	p-value
Restricted Model			
Portmanteau	35.3869	$\chi^2(34)$	0.4026
LM Test	50.9509	$\chi^2(40)$	0.1149
Normality Test[1]	10.1400	$\chi^2(4)$	0.0381
Normality Test[2]	5.8302	$\chi^2(4)$	0.2122
Multi ARCH LM	105.9129	$\chi^2(90)$	0.1207
Unrestricted Model			
Portmanteau	34.3055	$\chi^2(34)$	0.4531
LM Test	50.1311	$\chi^2(40)$	0.1309
Normality Test[1]	10.6703	$\chi^2(4)$	0.0305
Normality Test[2]	6.5362	$\chi^2(4)$	0.1625
Multi ARCH LM	105.0834	$\chi^2(90)$	0.1322
Trend Model			
Portmanteau	34.6757	$\chi^2(34)$	0.4356
LM Test	50.5539	$\chi^2(40)$	0.1225
Normality Test[1]	9.8362	$\chi^2(4)$	0.0433
Normality Test[2]	6.2237	$\chi^2(4)$	0.1831
Multi ARCH LM	104.7412	$\chi^2(90)$	0.1372

Note: all test were done taking 10 lags.

/1 Doornik and Hansen (2008).

/2 Lütkepohl (1993), p. 153.

Table 11. Statistics of the Booms Size

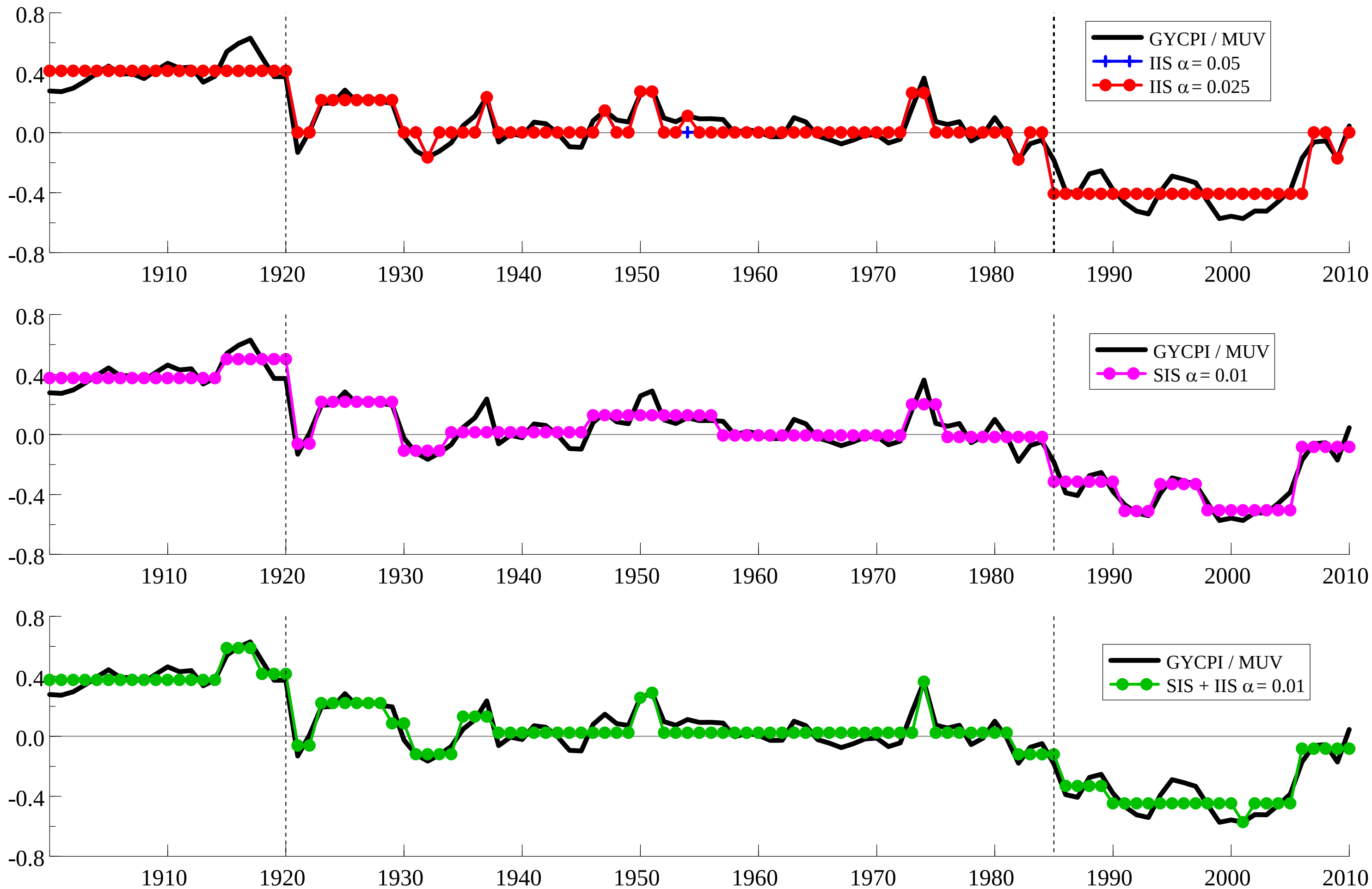
Booming Years	1915-1917	1924-1928	1937	1950-1951	1974	2007-2010
Number of Years	3	5	1	2	1	4
Regime Mean	0.41	0.05	0.05	0.05	0.05	-0.35
Mean of the Boom	0.18	0.18	0.19	0.23	0.32	0.29
Size of the Boom	0.53	0.91	0.19	0.46	0.32	1.17

Table 12. Fiscal Impact of the Change in Commodity Prices (in percent)

Country	Price Change with 2012			Impact on Revenues		
	2013	2014p/	2015p/	2013	2014p/	2015p/
Argentina	0.19	-10.40	-18.91	0.02	-0.94	-1.70
Bolivia	3.72	-9.14	-19.02	1.08	-2.65	-5.52
Chile	4.17	-11.00	-22.27	0.58	-1.54	-3.12
Colombia	-0.22	-11.11	-19.06	-0.05	-2.33	-4.00
Ecuador	-1.21	-10.17	-16.50	-0.42	-3.56	-5.78
Mexico	1.28	-10.12	-17.62	0.33	-2.63	-4.58
Peru	3.70	-9.08	-17.03	0.67	-1.64	-3.07
Trinidad and Tobago	2.11	-9.48	-16.72	0.74	-3.32	-5.85
Venezuela	1.28	-10.12	-17.62	0.66	-5.26	-9.16

Note: p/ Projected change values.

Figure 1. Saturation Tests for the Index of Relative Prices

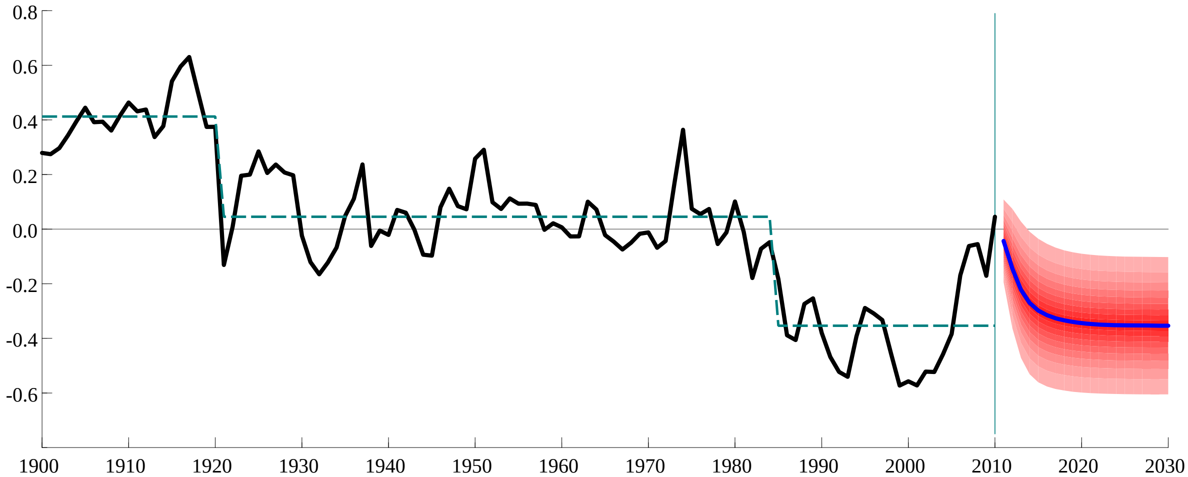


Note: Updated Grilli and Yang Index deflated by manufacturing export unit values, using data form Pfaffenzeller, Newbold, and Rayner (2007).

Impulse indicator saturation (IIS) as Hendry, Johansen, and Santos (2008) and Step-Indicator Saturation (SIS) as Doornik, Hendry, and Pretis (2013). All the analysis was using *Autometrics* in PcGive with manual grouping and a target size of 0.05, 0.025 and 0.01.

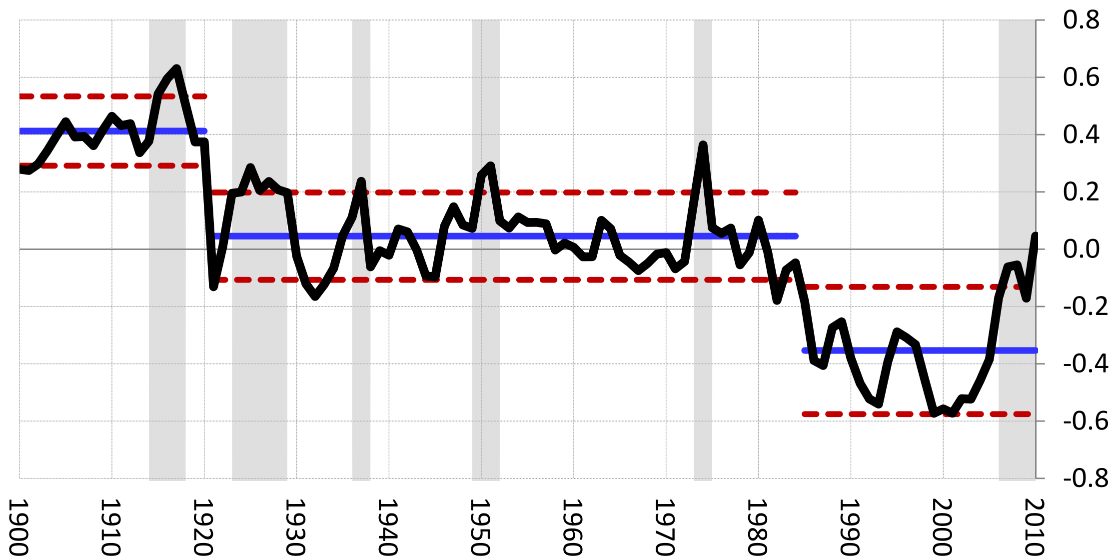
Vertical lines indicate the breaks for the Bai and Perron (1998, 2003) test with a trimming factor of 0.15 and allow up to 5 breaks. We use the BIC to conclude in favor of 2 breaks shown with vertical lines. We perform the test using the *strucchange* package in R.

Figure 2. Relative Price Index Forecast



Note: Updated Grilli and Yang Index deflated by manufacturing export unit values, using data form Pfaffenzeller, Newbold, and Rayner (2007). Dynamic forecast for 20 years ahead, confidence fan chart for 2 standard deviations.

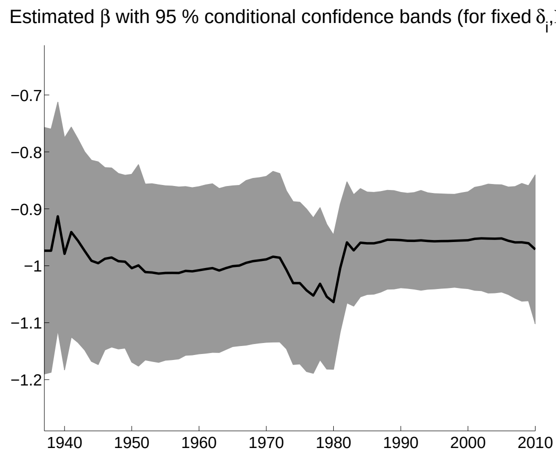
Figure 3. Index of Relative Prices, Booms and Breaks



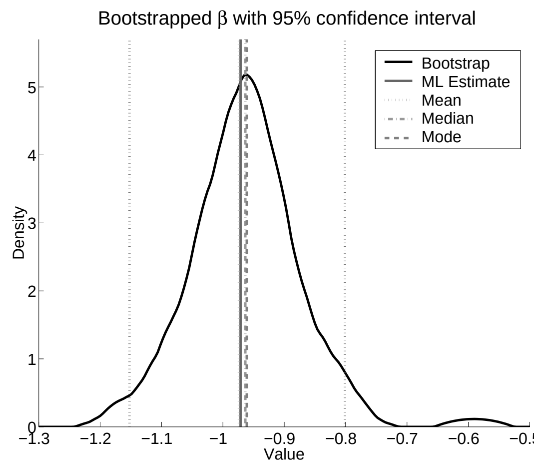
Note: Updated Grilli and Yang Index deflated by manufacturing export unit values, using data form Pfaffenzeller, Newbold, and Rayner (2007). Also potted is the mean (in blue) and 90% confidence interval (dotted red) for each regime. The booms are highlighted by gray areas.

Figure 4. Long-Run Coefficients on LMUV

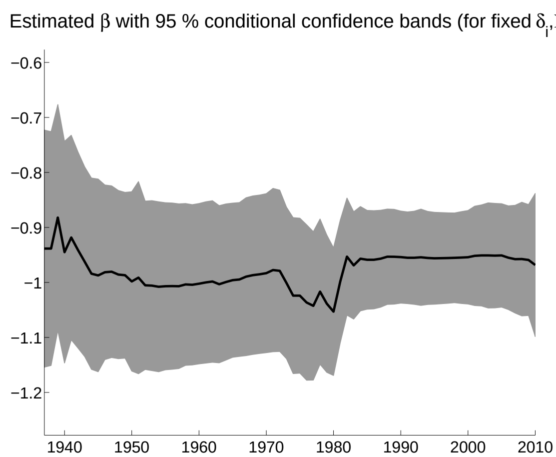
(a) Recursive Coefficient for Restricted Model



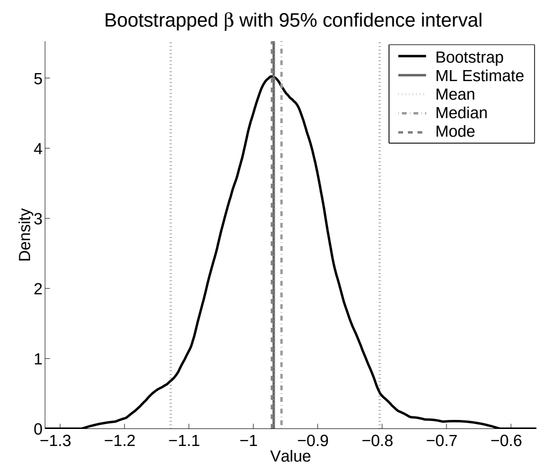
(b) Bootstraped Coefficient for Restricted Model



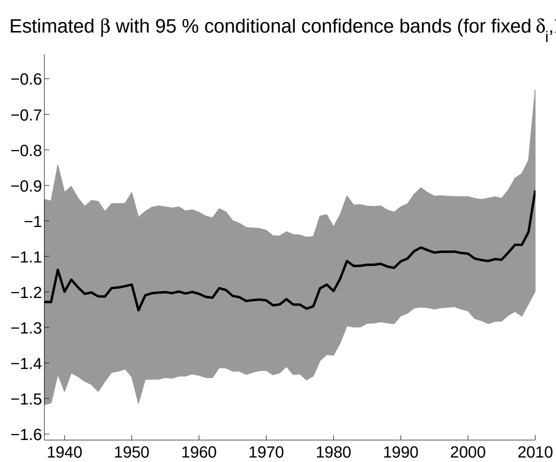
(c) Recursive Coefficient for Unrestricted Model



(d) Bootstraped Coefficient for Unrestricted Model



(e) Recursive Coefficient for Trend Model



(f) Bootstraped Coefficient for Trend Model

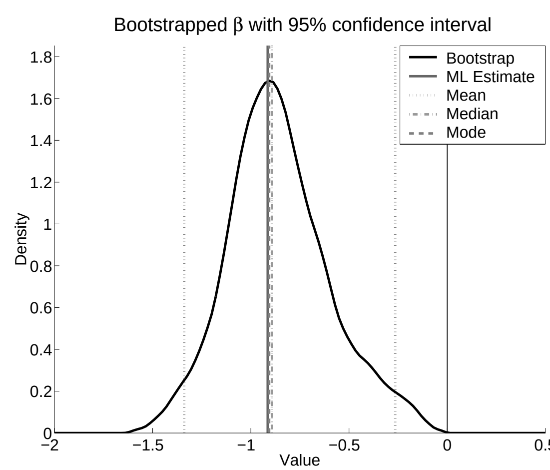
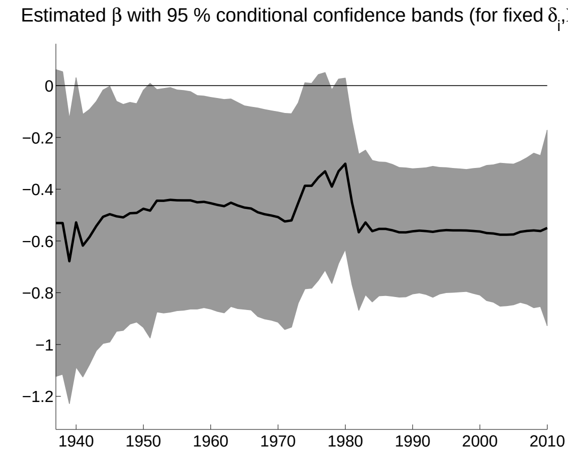
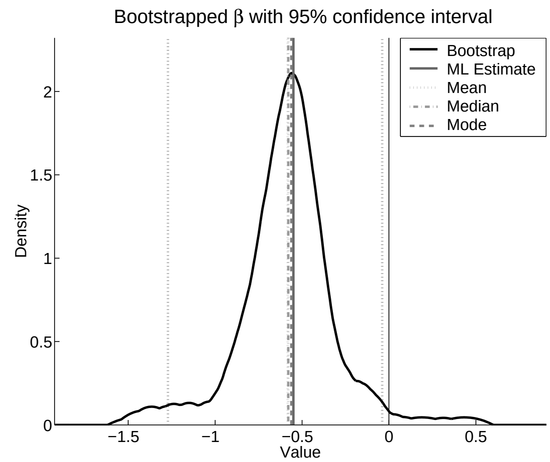


Figure 5. Long-Run Coefficients (Constant and Trend)

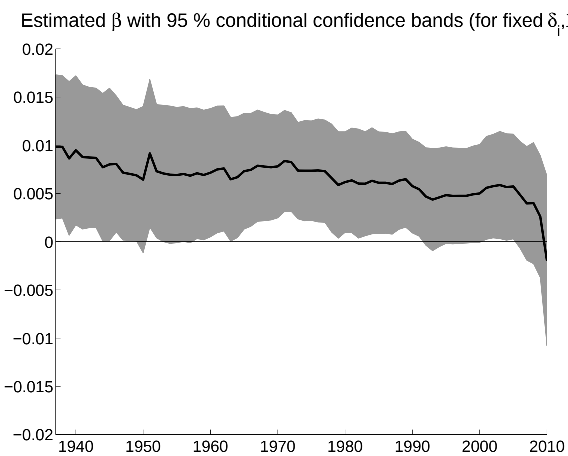
(a) Recursive Coefficient for Restricted Model



(b) Bootstraped Coefficient for Restricted Model



(c) Recursive Coefficient for Trend Model



(d) Bootstraped Coefficient for Trend Model

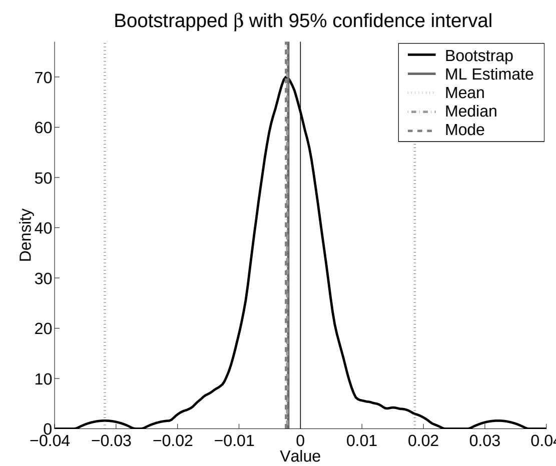
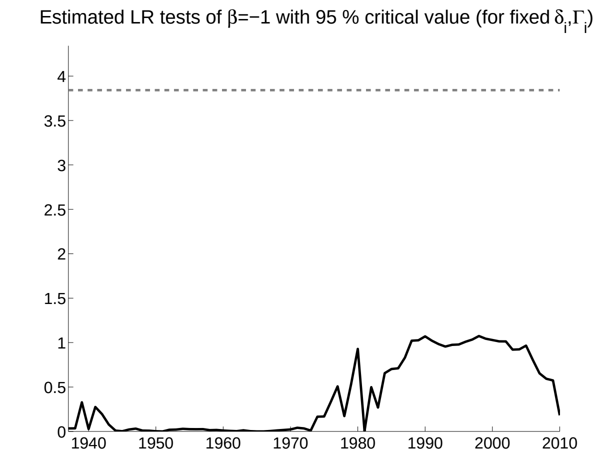
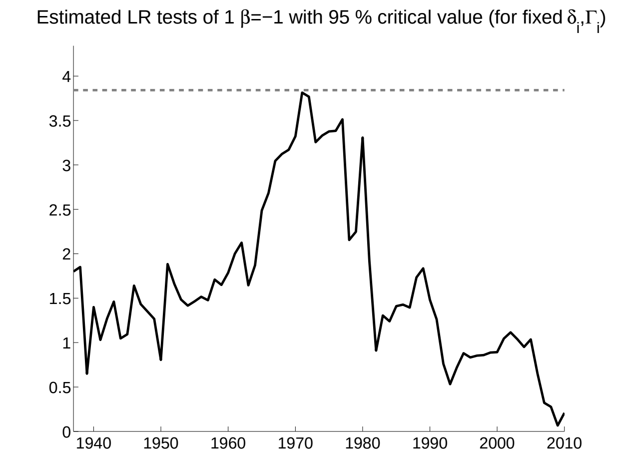


Figure 6. Recursive LR Tests for Restrictions on Long-Run Coefficients

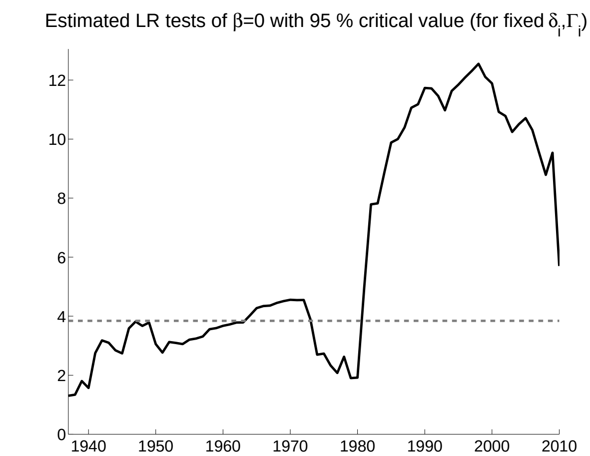
(a) Test for Restricted Model on LMUV



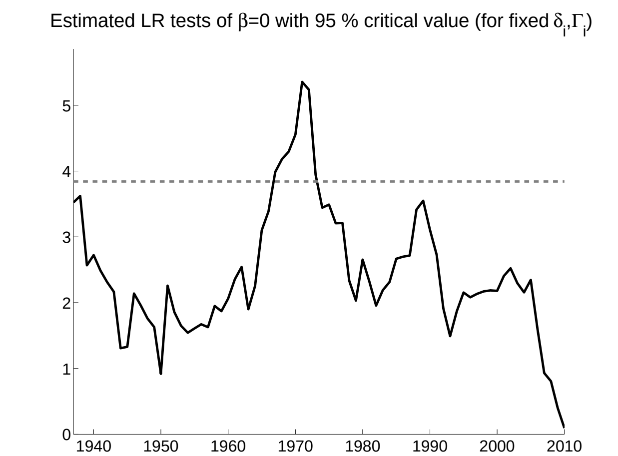
(b) Test for Trend Model on LMUV



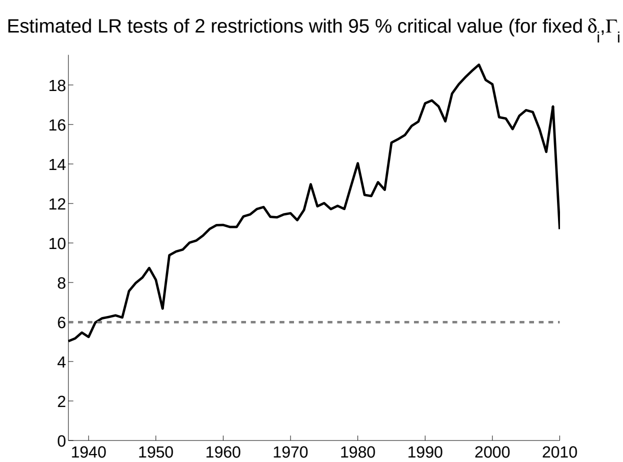
(c) Test for Restricted Model on Constant



(d) Test for Trend Model on Trend



(e) Test for Restricted Model on Both



(f) Test for Trend Model on Both

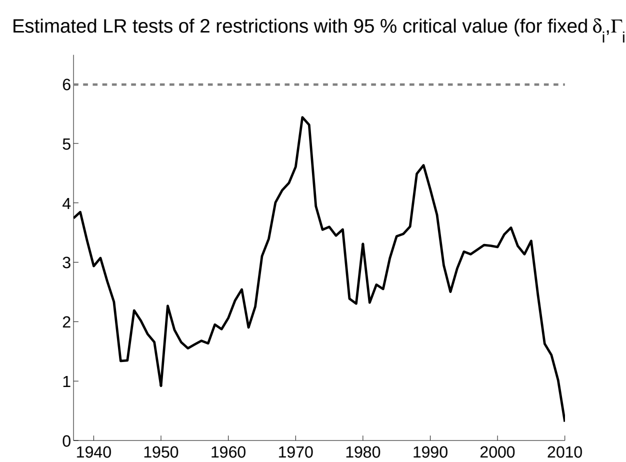
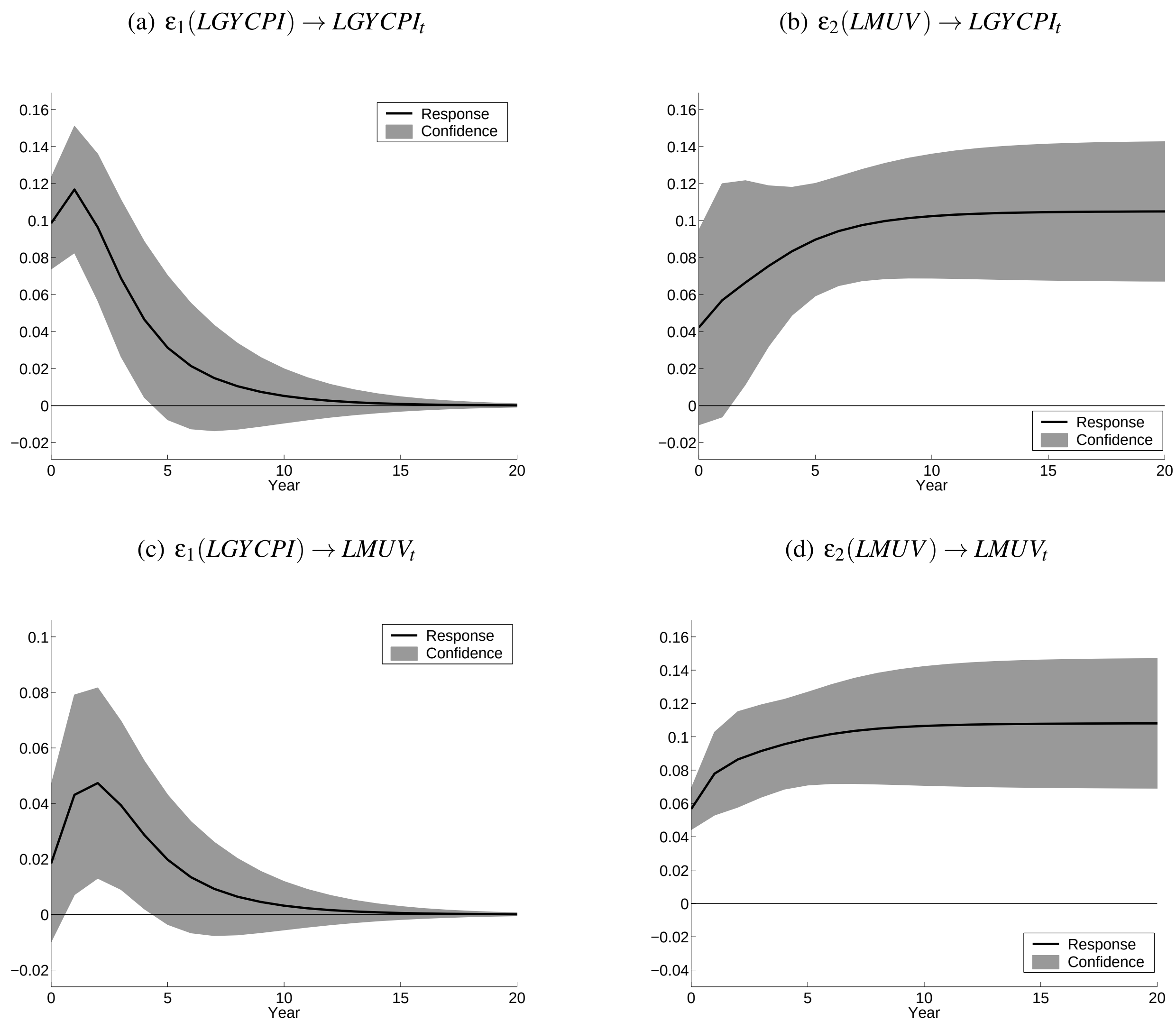


Figure 7. Impulse-Response Functions to Structural Shocks



Note: Structural impulse-responses to one unit shock to the structural error. We estimate these shock with the Restricted model, assuming that the long-run total impact of $LGYCPI_t$ on MUV_t is zero. The expression of the structural VEC is the following:
$\mathbf{\Gamma}(L)\Delta\mathbf{Y}_t = \boldsymbol{\alpha}\mu_0 + \boldsymbol{\alpha}\boldsymbol{\beta}'\,\mathbf{Y}_{t-1} + \boldsymbol{\delta}_3\mathbf{D}_t + \boldsymbol{B}\boldsymbol{\varepsilon}_t$, where $\boldsymbol{\varepsilon}_t \sim N(\,\mathbf{0}, \boldsymbol{I})$, $\boldsymbol{B}\boldsymbol{B}' = \Omega$, $\boldsymbol{A} = \boldsymbol{C}\cdot\boldsymbol{B}$, and $\boldsymbol{C} = \boldsymbol{\beta}_\perp\left(\boldsymbol{\alpha}'_\perp\mathbf{\Gamma}(1)\boldsymbol{\beta}_\perp\right)^{-1}\boldsymbol{\alpha}'_\perp$.

$$\underset{\text{long-run}}{A} = \begin{bmatrix} 0 & \underset{(0.019)}{0.105} \\ a_{21} \equiv 0 & \underset{(0.020)}{0.108} \end{bmatrix} \qquad \underset{\text{contemporaneous}}{B} = \begin{bmatrix} \underset{(0.013)}{0.099} & \underset{(0.027)}{0.042} \\ \underset{(0.014)}{0.018} & \underset{(0.006)}{0.057} \end{bmatrix}$$

Figure 8. Diagnostic Figures on the Restricted Model

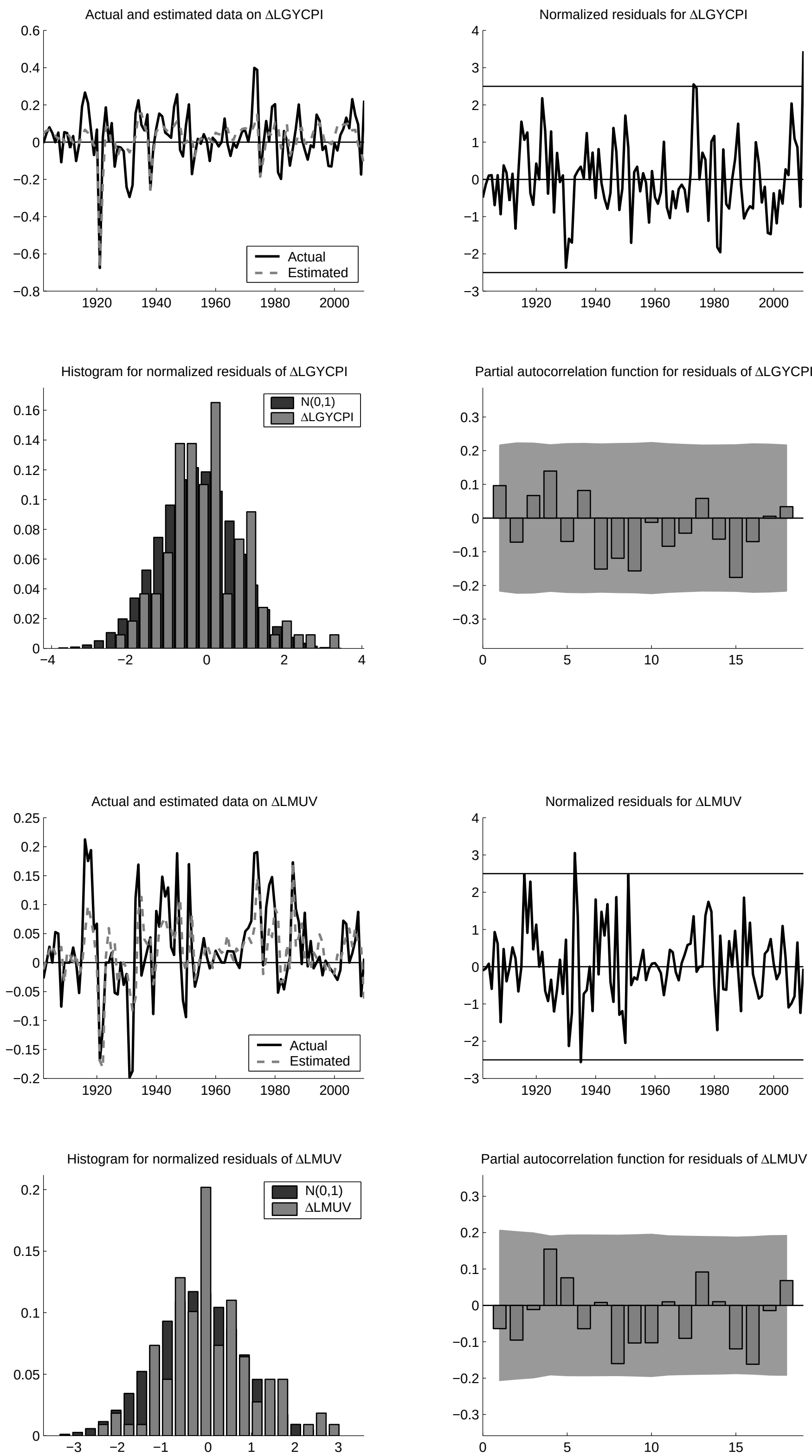


Figure 9. Diagnostic Figures on the Unrestricted Model

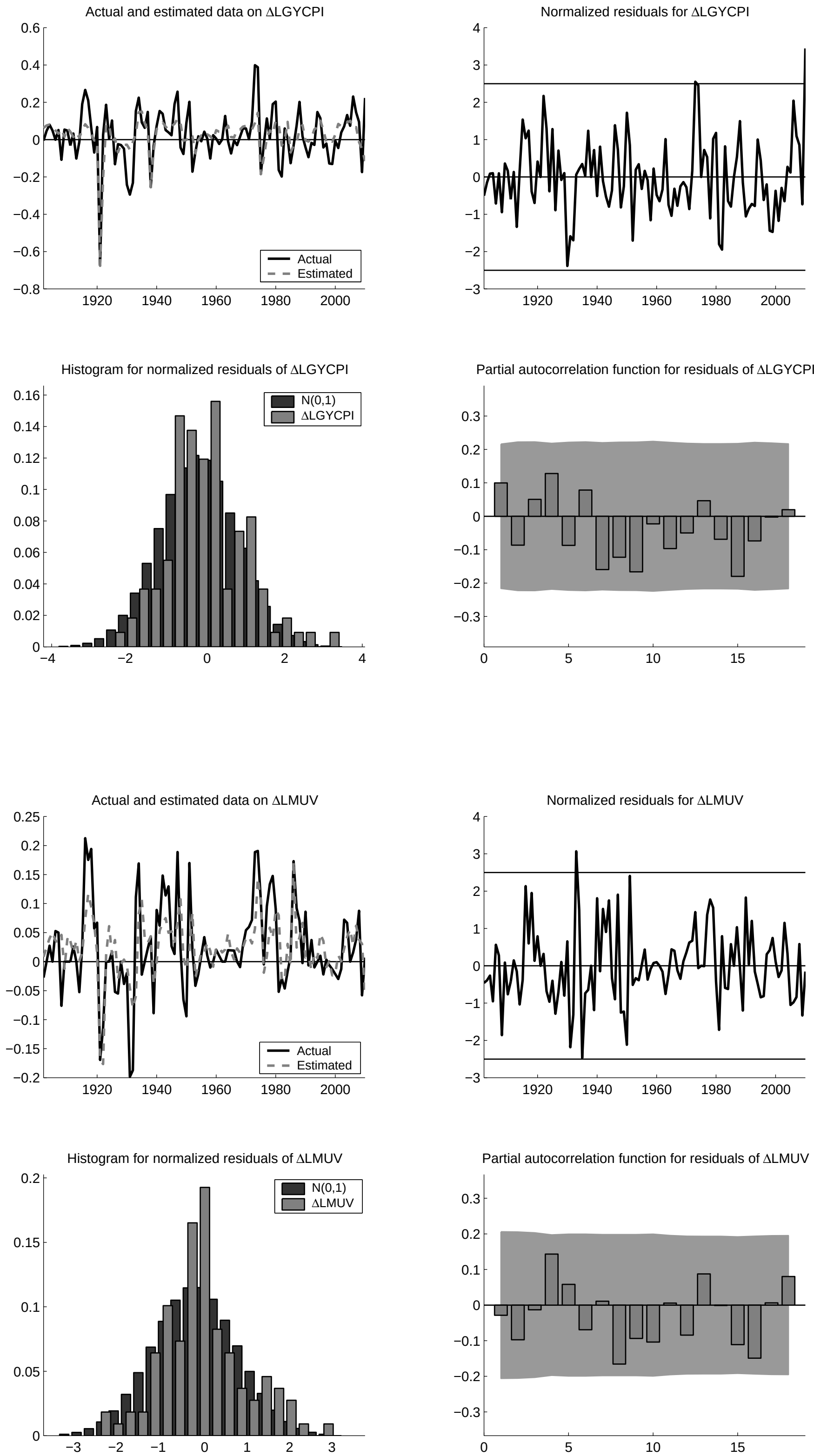
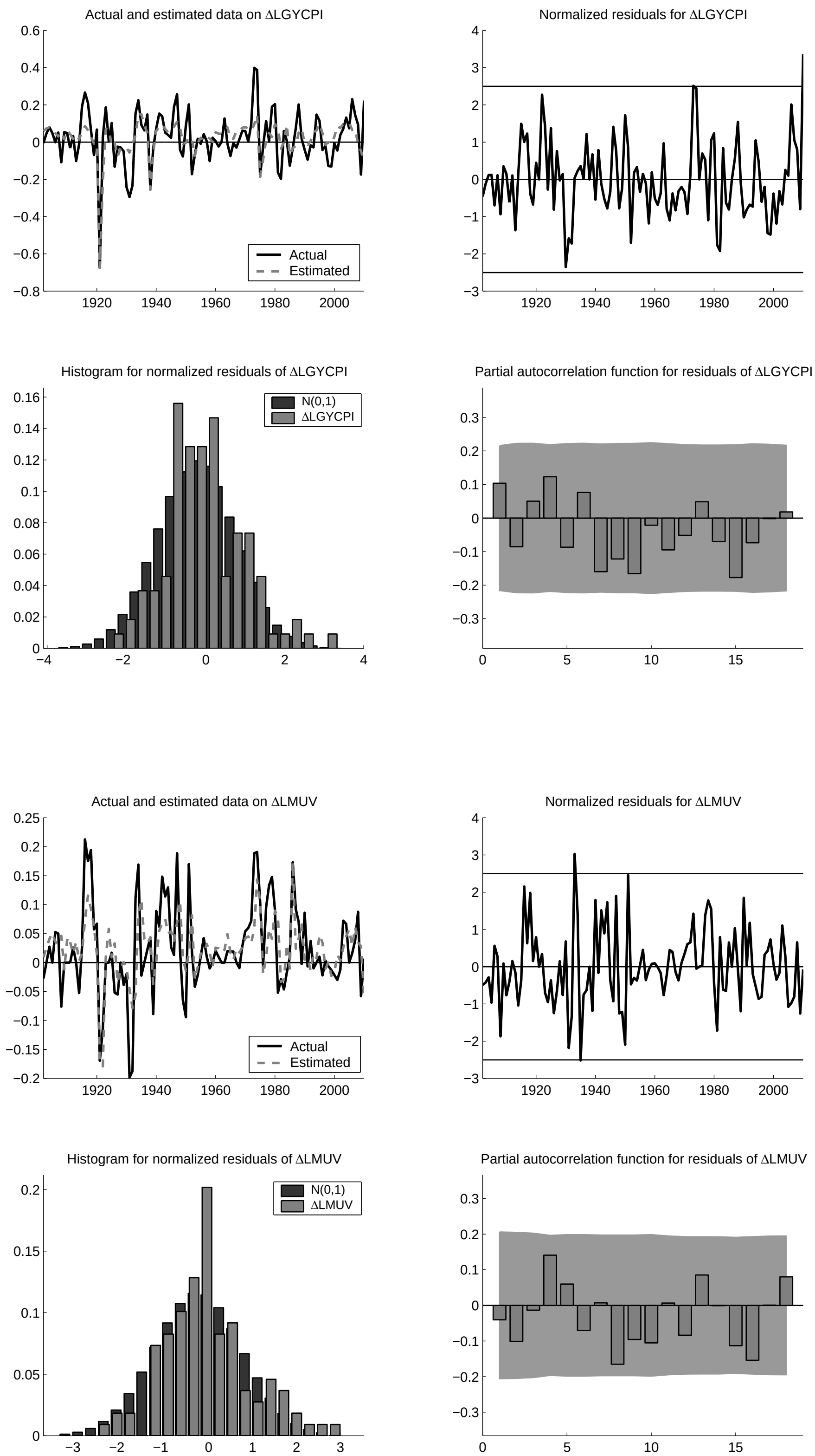


Figure 10. Diagnostic Figures on the Trend Model



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