Monetary Disorder and Financial Regimes. The Demand for Money in Argentina, 1900-2006*

Matteo Mogliani†, Giovanni Urga‡ and Carlos Winograd§

†*Paris School of Economics, France*
‡*Cass Business School, London, UK, and Bergamo University, Italy*
§*Paris School of Economics, France, and University of Evry-Val d'Essonne, France*

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Abstract

Argentina is a unique experience of protracted monetary disorder. In the framework of a long-term view, we investigate the demand for narrow money and the welfare cost of inflation in Argentina from 1900 to 2006. The paper examines the effect of monetary regimes by dealing with the presence of structural breaks in long-run equations. We estimate and test for regime changes through a sequential approach and we embed breaks in cointegrating single-equation models. Though our estimated models are comparable with those reported for industrialized countries, significant breaks appear consistent with major policy shocks occurred in Argentina during the 20th century.

Keywords: Money demand, financial regimes, structural breaks, single-equation cointegration, cointegration test, inflation, welfare, Argentina monetary history.
JEL classification number: E41, E42, C13, C22.

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† *Corresponding author*, Paris School of Economics, Paris-Jourdan Sciences Economiques, 48 Boulevard Jourdan, 75014 Paris (France). Tel: +33 (0)1 43 13 63 22. Fax: +33 (0)1 43 13 63 10. Email: mogliani@pse.ens.fr.

‡Cass Business School, Faculty of Finance, Centre for Econometric Analysis, 106 Bunhill Row, London EC1Y 8TZ (U.K.). Email: g.urga@city.ac.uk

§Paris School of Economics, Paris-Jourdan Sciences Economiques, 48 Boulevard Jourdan, 75014 Paris (France). Email: winograd@pse.ens.fr.

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1 Introduction

Since the early days of nation-building in the 19th century, Argentina has been a laboratory for monetary economics. A large variety of extreme economic episodes in the midst of political instability characterizes the history of the country: from rapid development in the *Belle Epoque* of 1880-1930 to economic crisis, from gold-standard stability to chronic monetary disorder, recurrent balance of payments crises, speculative attacks on the currency, hyperinflation episodes and debt defaults (Diaz Alejandro, 1970; Véganzones and Winograd, 1997; Della Paolera and Taylor, 2001, 2003). The striking variety of macro policy experiments and monetary events made Argentina a fascinating and puzzling case-study for economic historians and applied economists.

The extreme monetary episodes observed in Argentina over the last 30 years have led researchers to show increasing interest in the complex nature of these monetary phenomena. The research in modeling based on high-frequency data for short sample periods were used to disentangle the mechanisms driving money demand dynamics during the chronic high-inflation years of the 1970s and 1980s as well as the hyperinflation episodes of the late 1980s and early 1990s (Ahumada, 1992; Melnick, 1990; Kamin and Ericsson, 1993, 2003).

On the other hand, since the seminal works of Friedman (1956, 1959) and Friedman and Schwartz (1963, 1982) on the monetary history of the United States and the United Kingdom, low-frequency data for long sample periods have been at the heart of numerous studies in the money demand literature (Lucas, 1988, and Baba et al., 1992, for the US; Hendry and Ericsson, 1991, for the UK; Muscatelli and Spinelli, 2000, for Italy). These works showed that historical data, along with simple econometric models, are able to fit the stylized monetary facts in most industrialized countries (Bordo and Jonung, 2003).

Only recently research has started to pay attention to the historical and econometric analysis of monetary aggregates in Argentina during the 20th century (Choudhry, 1995b; Urga and Winograd, 1997; Ahumada and Garegnani, 2002; Gay, 2004). However, these contributions are often inconclusive on many issues, such as the decision making model for rational asset allocation of economic agents under conditions of extreme volatility, as well as the existence of long-run monetary equilibria and the stability of money.

This paper addresses these issues for Argentina, using annual data from 1900 to 2006. Our contribution focuses on the estimation of a long-run money demand model in a single-equation framework, providing economically interpretable opportunity cost parameters. We also consider the issue of shocks and long-run persistent financial and monetary distress in Argentina.

Drawing from the rich empirical literature on monetary instability, "missing money" and financial innovations in the industrialized countries (Judd and Scadding, 1982; Baba et al., 1992; Carlson et al., 2000), we account for financial and monetary (and, more generally, macroeconomic) regime changes. For this purpose, we date and test for structural breaks in the money demand function through a sequential bootstrap procedure. We consider the M1 aggregate as the monetary measure in our money demand analysis. From an empirical point of view, the behavior of the domestic currency allows us to explore a set of relevant economic perturbances in the monetary history of Argentina, such as currency substitution and speculative attacks, as well as persistent demonetization and the chronic loss of confidence in paper money. In addition, a study on the narrow money aggregate allows us to "isolate the numerical *constants* of monetary behavior" (Friedman, 1956) and it would straightforwardly extend to further discussions about seignorage, the welfare cost of inflation and monetary hysteresis.

Our main findings are as follows. *First*, we deal with instability in long-run relationships obtained through a standard cointegration analysis. Structural breaks analysis suggests the presence of significant regime shifts in the money demand model, consistent with historically observed macroeconomic and financial shocks in Argentina. Stability and long-run equilibrium features are then crucially related to the estimate of relevant regime shifts. *Second*, unlike the existing literature on Argentina, we report evidence suggesting that the hypothesis of income homogeneity cannot be rejected in our money demand models. Estimated interest rate and inflation coefficients are both highly significant, while they sharply differ in magnitude. This may reflect the strong financial control and the credit market regulations that affected Argentina even during periods of high inflationary pressures, leading to persistent negative real interest rate on regulated deposits. *Third*, we explore the econometric issue of long-run equilibrium with regime changes, by implementing *ad hoc* tests for cointegration with multiple breaks. Results suggest that

estimated long-run equilibria are consistent with the cointegration hypothesis. *Finally*, short-run estimates for money demand dynamics are reported.

The remainder of this paper is as follows. In Section 2, we report a brief introduction to the monetary history of Argentina during the 20th century and early 21st century. A brief review of the theoretical and empirical literature on money demand in Argentina is reported in Section 3. In Section 4, we present the dataset and we report preliminary empirical results on unit-root tests and cointegration, mainly to stress the instability features arising in long-run equilibria when potential structural breaks for monetary regime changes are neglected. In Section 5, we date and test for structural breaks in the money demand model and we discuss their economic and historical relevance. In Sections 6 and 7 we report estimates for the money demand models. Finally, in Section 8 we discuss the implication of our empirical model for the welfare cost of inflation in Argentina. Section 9 concludes.

2 A Macro-Monetary Overview of Argentina, 1900-2006

In this section we report a brief introduction to the monetary history of Argentina.

2.1 The "golden age" period (1880-1930)

The development of the financial sector in Argentina dates back to the second half of the 19th century. This period shows an unprecedented economic effervescence, the so-called *Belle Epoque* (Diaz Alejandro, 1970; Véganzones and Winograd, 1997; Della Paolera and Taylor, 2001, 2003). Nonetheless, many attempts to anchor the monetary system to the international gold-standard failed due mainly to inconsistent macroeconomic policies, or the result of the international business cycle. The gold-standard regime was restored in 1899, becoming fully operative in 1900. Monetary policy in Argentina was then conducted in accordance with the strict rule of the gold-standard regime: a Conversion Office and a national commercial bank (Banco de la Nacion Argentina) operated as independent monetary authorities, assuring a decade of remarkable financial stability until 1913. The First World War led to the collapse of the gold-standard system, but the dynamism of the economy was still the source of the high level of development that Argentina had

achieved. Nevertheless, the financial collapse of the United Kingdom in 1921, the short-lived experience of resumed convertibility in 1928-1929 and the incoming world economic crisis led to the end of the "golden age".

2.2 The crisis of the 1930s and the politics of limited public intervention

The world crisis of 1930 broke the long phase of rapid growth and virtually cut off the inflow of foreign capital to Argentina. At the same time, the rapidly deteriorating terms of trade and the collapse of economic activity led to a series of bankruptcies, which further worsened the liquidity crisis in the banking system, turning it into a solvency crisis. To contend with this adverse economic scenario, the government created the Central Bank in 1935. Its objectives were to ensure monetary stability and smooth the output fluctuations, the Central Bank was given a monopoly of issuing bank notes and was assigned to the role of lender of last resort. Over the period 1935-1945, we can identify two phases of monetary policy. The first one (1935-1939) is characterized by price stability, with the Central Bank compensating the monetary impact of the balance of payment flows. In a second phase (1940-1945), a large surplus in the balance of payments was accompanied by the generalized expansion of means of payment. This set up showed in the inflationary pressures, initially well managed by the monetary authorities, that will later turn into a chronic feature of the Argentine economy. This new regime of *limited public intervention* was effective in the objective of smoothing output fluctuations and attaining a reasonable degree of monetary stability. However, disintermediation continued to increase (M3 to GDP falling to 40% in the early 1940s) due mainly to the fragility of the local financial system, but the average inflation rate of the period 1900-1940 was a low 1.8%.

2.3 Financial repression, disintermediation and chronic inflation (1946-1973)

The institutional changes were more radical after the end of World War II, when the first Peron government (1946-1951) was elected. The public authorities decided to "nationalize" bank deposits in 1946. The compulsory reserve requirements were fixed at 100% and all deposits of the commercial banks were managed in the name of the Central Bank for

a fix fee charged by the commercial banks. In the Peron years of 1946-55, under a regime of *massive pubic intervention* in the goods and financial markets, expansionist monetary and fiscal policies were followed, and inflation (in the range of 4-40%) became a chronic feature of the monetary behavior of the economy. A first step towards the liberalization of the financial system occurred in 1957 after the coup d'état that deposed General Peron in 1955. The nationalization of deposits was abandoned, the legal minimum reserves were reduced and the credit allocation policy once again decentralized. During the 1960s, the financial sector experienced the rapid development of a large number of specialized non-banking institutions. However, monetary financial instruments continued to fare badly due to inflation that reached a peak of 80% in 1959, as the result of sharp devaluations. By the end of 1960s the government tried to integrate the newly established non-banking institutions into the official system. But the appointment of an elected government headed by General Peron in 1973 set the beginning of a new regime, marked by accelerating inflation, macroeconomic instability and political turbulence.

2.4 High inflation, financial fragility and stabilization failures (1974-1990)

The newly elected Peronist government launched in 1973 a stabilization plan based on a rigid income policy coupled with expansionary fiscal and monetary policies. In 1975, after the death of Peron in 1974, a new stabilization plan collapsed on the attempt of a sharp readjustment of relative prices to contain macroeconomic disequilibria (the so-called *Rodrigazo*). After the military coup of 1976, a financial reform took place in 1977, when the financial system was liberalized and interest rates (for the first time) were freed. In addition, the reserve requirements were decreased (and remunerated by the Central Bank) and credit controls were further relaxed for private commercial banks. A second stabilization plan, the *Tablita Plan*, was launched in December 1978, focusing mainly on a pre-announced rate of devaluation for a set period of time in view of the gradual reduction of inflation. But these policies failed to stabilize the financial system and the Argentinean economy, mainly due to the absence of public finance consistency, leading to persistent high inflation, a confidence crisis and speculative attacks. This caused the economic and,

in particular, the banking crisis of 1980. The *Tablita Plan* collapsed in 1981, a sharp devaluation took place. The implementation of the *Austral Plan*, a monetary reform implemented in mid-1985 involving wages, prices, and exchange rate freezes, and later the *Spring Plan* (a mid-1988 agreement with the private sector to limit the growth of prices and the exchange rate devaluation at 4% per month) failed to deal with exploding inflation. The collapse of public finance, rapidly rising political instability and recurrent speculative attacks on the currency, led to the hyperinflationary episodes of 1989 and 1990. The degree of demonetization then reached extreme levels, with the M1 to GDP ratio standing at 1.5% in 1990 and M3 to GDP reaching low levels of 5%, compared to 70% in the first decades of the century. Monetary stabilization was obtained in 1991, under President Menem (Peronist Party) elected in 1989, with the *Convertibility Plan* (March 1991), which fixed the parity against the dollar and led to a sharp reduction of inflation to one digit yearly levels.

2.5 The Convertibility Law of 1991: successful stabilization and monetary hysteresis (1991-2000)

The Convertibility Law of 1991 and the creation of a Currency Board regime were introduced to address the credibility problem of chronic monetary disorder and the resulting massive dollarization of Argentina's financial instruments. The system thus consisted of a simple dollar standard put into practice as a bi-monetary system with both the dollar and peso as legal tender. The monetary system was rather similar to the gold-standard rule at the beginning of the century. Furthermore, the new law banned the Central Bank from monetary financing of the public debt, proclaiming *de facto* its independence, later *de jure*. The creation of money was only allowed simultaneously with increases in foreign currency reserves. Finally, the role of the Central Bank as lender of last resort was restricted. The monetary initiative was coupled with an intensive programme of privatization and regulatory reforms. The reduction of inflation was sharp and rapid, as in other post hyperinflation episodes. In the midst of the high liquidity of international capital markets, the newly gained confidence of Argentina led to massive capital inflows and solid reserves accumulation. This provided a basis for the monetary recovery of the country,

but the pattern was probably mined by a hysteresis phenomenon affecting the amount of liquidity held by agents and recovery resulted quite slow (the M1/GDP ratio rose to 4.4% in 1992, but it only attained 6.8% in 1999). However, the sudden disinflation led to a consumption boom, reinforced by capital inflows and public finance disequilibria. A gradual currency appreciation and increasing current account deficits cum growing indebtedness will eventually set the conditions for a credibility problem on the stability of the peso parity and the convertibility regime. The debt crisis in Mexico (1994) and later the emerging market crisis in Asia, Russia and in neighboring Brazil, with the contagion effects, only reinforced the confidence crisis on the sustainability of the monetary regime of Argentina. The fundamental conditions for the speculative attacks and deep economic crisis of 2001-2002 were set.

2.6 The end of Convertibility, debt default and back... (2001-2006)

Fiscal deficits, international financial crises and overvaluation of the peso reverted the virtuous circle of the early nineties. This generated pressures over commercial banks' liquidity and interbanking rates, spreading to the whole financial system and leading to a bank panic and a bank run over the domestic commercial deposits in 2001. The convertibility system was kept unchanged until January 2002, when the parity with the dollar was abandoned. Economic policy turned into a rather chaotic exercise, with deposit freeze, a monetary reform, banking rescue packages and debt default, amid daily demonstrations of the population in a state of shock and deep social distress. The Central Bank played a crucial role in the stabilization of both the exchange rate and the inflation. In particular, this was achieved through a mix of interventions in the domestic money market and in the foreign exchange market. After a deep contraction of output of 6% in 2002, the economy started a rapid and solid recovery in 2003 sustained by very favorable external conditions of strong improvement of the terms of trade, and a recovery of public finance. Inflation fell to one digit number to rise thereafter to 8-20% range in the coming years until the end of the first decade of the 21st century. Since the recovery, real narrow aggregates grew at faster rate, with the M1/GDP ratio attaining 12.3% in 2006, in a rather stable economic environment. However, we should highlight that, post hyperinflation cum price stabilization in 1991 and later, financial disintermediation showed strong signs of persistence and

hysteresis. Argentina's financial system never returned to the high values of the *golden age*.

3 Money demand in Argentina

There is a fairly large body of research on the theoretical structure of money demand. Judd and Scadding (1982), Goldfeld and Sichel (1990) and Sriram (1999) offer, among others, an exhaustive review on the theoretical and empirical literature.

From a theoretical point of view, political economy models offer a solid support for the analysis of monetary behaviors dealing with protracted inflation regimes, hyperinflation diseases, financial adaptations (currency substitution, capital flights, demonetization) and stabilization attempts.[1] The political economy framework fits particularly well in the case of Argentina, where inflation processes and distributive struggle, financial instabilities and monetary regimes are main long-term policy elements.

Despite recent advances introduced in the monetary theory [2], the money demand function used in empirical applications is common to most models developed in the literature since the seminal works of Keynes (1936), Baumol (1952), Tobin (1956) and Friedman (1956, 1959). The standard theory of money demand posits:

$$\frac{M^d}{P} = f(Y, \mathbf{R}) \tag{1}$$

where M is the nominal stock of money demanded, P is the prices level, Y is the scale variable (*i.e.* an income variable, usually defined in real terms) and $\mathbf{R}$ is a vector of opportunity costs in holding nominal balances. The function $f(\cdot, \cdot)$ is assumed to be increasing in income and decreasing (increasing) in the vector of the opportunity costs which reference assets are outside (inside) the monetary aggregate M.

Many attempts to estimate a money demand function for Argentina can be found in the literature. Most empirical studies pay more attention to the high-inflation or

[1] These models develop a distributive conflict framework with heterogeneous agents and optimal inflation tax (Laban and Sturzenegger, 1994; Mondino et al., 1996; Ahrend et al., 2008).

[2] See, for instance, the general transaction and portfolio joint models (Ando and Shell, 1975; McCallum and Goodfriend, 1987) and models with money in a general equilibrium framework, as the models with imperfect markets (Bewley, 1983; Kehoe et al., 1990) and the class of search models (Kiyotaki and Wright, 1991).

hyperinflation episodes of the 1980s and to the monetary effects of the *Convertibility Plan* during the 1990s, by using high-frequency short samples. Ahumada (1992) and Melnick (1990) develop models of real M1 demand in Argentina between 1977 and 1988, showing the existence of non-linearities in the short-run dynamics of money demand caused by high and rising inflation rates. Choudhry (1995a) restrains the analysis on the existence of stationary long-run M1 and M2 functions between 1975 and 1987, finding significant results for both monetary aggregates. The monetary biases (non-linearities) generated by the hyperinflation episode are mainly handled by Kamin and Ericsson (1993, 2003), who develop a model of real M3 demand between 1978 and 1993. Finally, Aguirre et al. (2006) analyze the demand for money during the Convertibility period, the economic crisis and the recovery (1993-2005), finding poor results in terms of stability of the implied monetary relationships.

In sharp contrast, only a few empirical studies on low-frequency data deal with the issue of the long-run stability of money demand in Argentina. Choudhry (1995b) find a stationary long-run money demand relationship between 1935 and 1962, using quarterly values for real M1 and M2, real net national income and an inflation rate. However, results inconsistent with theoretical priors may suggest some form of misspecification in the underlying VECM. Gay (2004) analyze the demand for real M1 balances in the context of a small open economy model with nontradable goods. Evidence of endogeneity in the income variable may suggestively point out the non-neutrality of money in Argentina or, in turn, some misspecifications in the long-run equation. Finally, Urga and Winograd (1997) and Ahumada and Garegnani (2002) deal with the long-run instability issue of the money demand relationship in Argentina. The latter use annual data for real cash balances from 1935 to 2000 and introduce exogenous breaks in the long-run parameters[3], while the former show that the demand for real M1 between 1900 and 1992 is affected by few (endogenously estimated) structural breaks.

Results reported in previous works highlight the crucial issue of the long-run instability of the demand for money in Argentina. The presence of abrupt changes in monetary

[3]The income homogeneity restriction ($\beta_y = 1$) is assumed to hold from 1935 to 1955, while the Baumol-Tobin restriction ($\beta_y = 0.5$) is assumed to hold from 1956 to 2000. In addition, the interest rate is assumed to be the relevant opportunity cost from 1935 to 1973 and from 1991 through the end of sample, while the inflation is the relevant opportunity cost from 1973 to 1991.

regimes, as suggested by the evident statistical difficulty in describing monetary phenomena, may jeopardize the standard econometric approach. In this paper we address the issue of modeling instability making use of recent developments in econometric treatment of breaks.

4 Preliminary Empirical Analysis

In this section we provide a short description of the database and we report preliminary empirical analysis on unit-root features of the data and on cointegration.

4.1 Data and empirical model

Our dataset spans from 1900 to 2006, *i.e.* from the beginning of the long lasting gold-standard experiment to the aftermath of the deepest economic crisis ever experienced in Argentina (see the Data Appendix for more details). Let m be the log of the nominal M1, *i.e.* the sum of the total currency held by the public, plus small demand deposits. The annual (log) real money aggregate $(m - p)$ is computed as the annual average of monthly nominal values of M1, divided by the respective monthly value of the wholesales price index. This allows us to assign a weight to large deviations of prices and to better account for the infra-annual volatility of monthly real balances. The transaction variable is the (log) real GDP (y), computed as the annual nominal GDP divided by its implicit price deflator. The price index chosen for this study is the wholesales price index (p), rather than the CPI series, because of data quality and availability. As before, the series is computed as an annual average of monthly price index values. The inflation rate (π) is thus computed as the first difference of the annual log price series. Finally, the external opportunity cost of holding real balances is the annualized interest rate over relevant alternative assets (r), which enters the model as $r_t = \ln(1 + R_t)$, that is as a tantamount to the semi-log specification.[4] In Figure 1, we report the plots of the series, scaled in log values. The underlying empirical model we shall test below takes the following standard

[4]Given the difficulty to rebuild homogeneous series for interest rates throughout the century, we used different measures for the external opportunity cost. Rates over time deposits and saving deposits, as well as regulated and non-regulated interest rates and free market rates, were then alternated to recompose a relevant interest rate series. In addition, it is worth noticing that a double-log specification setting would not be coherent with the extreme variation of nominal variables (interest rate and inflation rate) during high-inflation episodes.

form:

$$(m-p)_t = \alpha + \beta_y y_t + \beta_r r_t + \beta_\pi \pi_t, \tag{2}$$

where β_y is the income elasticity and β_r and β_π are the interest rate and inflation semi-elasticity, respectively.

[Figure 1 about here]

4.2 Unit-root tests

The very long time span of our dataset and the high degree of volatility should allow for some non-stationarity in mean and variance of the series. We check this hypothesis using a battery of unit-root tests introduced by Ng and Perron (2001), such as the modified versions of the Elliott et al. (1996) feasible point-optimal test (MP_T^{GLS}) and GLS Dickey-Fuller test (ADF^{GLS}), as well as the GLS Phillips-Perron t-test (MZ_T^{GLS}). The power of this class of modified unit-root tests is dramatically improved by the use of the GLS detrending of the series. The optimal lag selection criterion for the computation of the Ng and Perron (2001) autoregressive spectral density function is corrected following Perron and Qu (2007). On the basis of results reported in Panel A of Table 1, the null hypothesis of unit-root cannot be rejected for all variables, even though the ADF^{GLS} and the MZ_T^{GLS} statistics are very close to the 5% rejection area for the inflation series. Our results do not conflict with findings reported elsewhere (for instances, Ahumada and Garegnani, 2002, and Choudhry, 1995b) when different sample windows are used.

[Table 1 about here]

4.3 Cointegration analysis

We employ cointegration techniques to estimate long-run money demand parameters. The multivariate system considered here involves the monetary aggregate, one transaction variable and two measures of the opportunity cost in holding real balances.

The macroeconomic theory on open economies asserts the existence of a theoretical equilibrium between prices, interest rates and the exchange rate. In the long-run, this

means that the PPP (Purchase Power Parity) and the CIRP (Covered Interest Rate Parity) should hold. International nominal equilibria lead to a domestic equilibrium between nominal interest rates and the (expected) change in prices, that is a Fisher parity in the domestic real interest rate. Assuming that in the very long-run the international equilibria hold, a multivariate cointegrated system involving money, income, inflation and nominal interest rate should be of rank order 2: a money demand equation and a real interest rate equation. However, in order to identify individual cointegrating vectors, we must impose an exclusion restriction over at least one variable for each relationship.[5]

This means that we can identify two different money demand models involving the inflation rate and the nominal interest rate. If the strong Fisher parity hypothesis is supported by data, *i.e.* the homogeneity restriction $\{1, -1\}$ over coefficients is not rejected, the two models are statistically equivalent and it is irrelevant which opportunity cost is excluded from the equilibrium specification (Hoffman et al., 1995). Otherwise, the two models are not equivalent and a different inference would be necessary.

Results from cointegration tests are reported in Panel B of Table 1. According to the Stock and Watson (1988) and the Johansen (1991) tests, the null hypothesis of no cointegration can be strongly rejected for the real interest rate equation, as well as for the money demand model. In addition, the null hypothesis of a single cointegrating vector can also be rejected for the money demand model, confirming the existence of two long-run relationships. Further, we estimate the rank 2 model by maximum likelihood, imposing an exclusion restriction over either the inflation rate (Model 1) or the interest rate (Model 2). The first cointegrating equation is then the money demand model, while the second cointegrating equation describes the real interest rate relationship. Results are presented in Panel C of Table 1. Restrictions over long-run coefficients are tested and imposed. In particular, the Baumol-Tobin restriction over the real GDP parameter ($\beta_y = 0.5$) cannot be rejected for both Models 1 and 2, while the strong Fisher parity hypothesis must be rejected.[6] The latter implies that the two money demand models are not statistically

[5]Given the existence of wide periods of fixed exchange rate regimes (convertibility under the gold-standard and the 1991 stabilization plan), we exclude the exchange rate from the long-run analysis and we assume that prices and interest rates embed enough information to capture most dynamics of the demand for real balances.

[6]Panel C of Table 1 reports results involving a large set of significant restrictions on the long-run coefficients matrix β and the adjustment coefficients matrix α. When the quantitative theory restriction on the scale variable coefficient ($\beta_y = 1$) is imposed, the LR statistic rejects both models 1 and 2, at less

equivalent and two different modeling strategies would be necessary. That is:

$$\text{MODEL 1:} \qquad (m-p)_t = \alpha + \beta_y y_t + \beta_r r_t$$

$$\text{MODEL 2:} \qquad (m-p)_t = \alpha + \beta_y y_t + \beta_\pi \pi_t$$

Further, we check the stability of the estimated money demand cointegrating relationships. We consider the *SupF* and the *MeanF* statistics developed by Hansen (1992) and based on the Phillips and Hansen (1990) Fully-Modified estimator. Results suggest a strong rejection of the null hypothesis of stable cointegration for both Models 1 and 2 (p-value$_{SupF} < 0.01$ and p-value$_{MeanF} < 0.01$ in both cases).[7] It is worth noticing that the statistical instability of cointegrating vectors is contradictory with the idea of cointegration. Quite surprisingly, our results suggest that cointegration tests spuriously detect cointegration when the underline model neglects some unknown feature of data acting as a source of long-run instability.

The monetary history of Argentina can help us to shed some light on this issue. As already mentioned in Section 2, Argentina went through many political, economic and monetary regimes since the beginning of the 20th century. We can then expect that the switch from one regime to an other, such as the collapse of a monetary system or rapid institutional changes, would be the main source of instability for money demand relationships. On the other hand, new monetary regimes can be set by long lasting persistent features affecting the demand for money in the long-run, as well as hysteresis effects over monetary behaviors and financial innovations/substitutions. Accordingly, our long span data analysis is obviously affected by the extreme economic features which took place in Argentina during the century, such as abrupt adjustments after full convertibility regimes and persistent money holding behaviors driven by chronic inflationary pressures. Neglecting these features would mean disregarding the stylized facts and reporting inaccurate econometric results.

than 1%. In addition, the adjustment coefficient $\alpha_{(m-p)}$ is not significant. When the strong Fisher parity restriction ($\beta_\pi = -\beta_r$) is imposed, the LR statistic rejects again both models 1 and 2, at 5% and 1% respectively.

[7] The *SupF* statistic tests for the null hypothesis of stable cointegration against the alternative of an unknown single change in the deterministic and stochastic long-run parameters, while the *MeanF* statistic assumes a constant hazard of parameters instability as alternative hypothesis. Estimates are computed involving a Quadratic Spectral kernel for the efficient estimate of the long-run covariance matrix and an automatic bandwidth selection method (Andrews, 1991).

We propose to embed monetary regime changes in our models by estimating structural breaks in money demand functions and then including them in the long-run equations. After a visual inspection of the velocity of money in the last panel of Figure 1, we claim that our definition of regime changes motivates the choice of deterministic structural breaks, rather than stochastic breaks. This approach should allow us to capture variations in financial equilibria, monetary trends and some institutional and economic unobserved factors, which may probably explain the instability of estimated cointegrating equations.[8]

5 Looking for structural breaks : the sequential estimation approach

We implement the sequential estimator developed by de Peretti and Urga (2004) for dating and testing break points in a multivariate framework (see also Banerjee et al., 1998, and Hansen, 2000). Following this approach, a vector of structural breaks is firstly estimated on the marginal process of a given system, and then this information is incorporated in the conditional process in order to estimate its own structural breaks.

Consider a bivariate model $w_t = (z_t, x_t)'$ with joint density function $f_w(w_t, \theta) = f_{z|x}(z_t|x_t, \lambda_1) \cdot f_x(x_t, \lambda_2)$, where the first term in the right hand side of the equation is the conditional density of z_t given x_t and the second term is the marginal density of x_t. Following de Peretti and Urga (2004), the reduced form system used to dating breaks is :

$$z_t = \mu_0 + \delta_0 t + \sum_{\tau=1}^{p_0} \rho_{0,\tau} z_{t-\tau} + \sum_{i=1}^{I_0} (\alpha_{0,i} B^M_{0,i,t} + \beta_{0,i} B^L_{0,i,t}) + \gamma x_{t-1} + \sum_{i=1}^{I_n} (\eta_i B^M_{1,i,t} + \theta_i B^L_{1,i,t}) + \varepsilon_t \quad (3)$$

$$x_t = \mu_1 + \delta_1 t + \sum_{\tau=1}^{p_1} \rho_{1,\tau} x_{t-\tau} + \sum_{i=1}^{I_1} (\alpha_{1,i} B^M_{1,i,t} + \beta_{1,i} B^L_{1,i,t}) + \nu_{n,t} \quad (4)$$

for $t = 1, 2, \ldots, T$, where Equation (3) is the conditional process and Equation (4) is the marginal process. Dummy variables $B^M_{i,t} = \mathbf{I}(t \geq b_i)$ and $B^L_{i,t} = (t - b_i + 1)\mathbf{I}(t \geq b_i)$ are designed to capture breaks b_i in the mean and in the linear trend, respectively. Marginal

[8] In other words, we depart from a simple linear specification. The aim is to replicate the actual dynamic of the monetary phenomena under analysis, which is probably characterized by small and big shocks. By handling structural breaks, we are more confident about picking up observable big shocks, the more relevant in our long span analysis.

and conditional processes also include p lags of the dependent variable.

Breaks dating is obtained through a sequential estimation which includes the breaks one by one. The structural breaks estimator is based on the conventional F-statistic estimator (Christiano, 1992), *i.e.* $\hat{b}_i$ is defined as the time t at which the F-statistic reaches is maximum ($\hat{b}_i = \text{argmax}\{F_t\}$). However, the sequential estimate of the structural breaks is showed to be biased in the presence of more than one break: the dating of the first break is biased when a second break is neglected, and so on. To overcome this problem, a backward revision of estimated breaks is proposed.

The sequential estimation approach usually poses the problem of stopping the search when further estimated breaks are no longer significant. The conventional asymptotic test for this hypothesis is showed to lead to an important bias in the rejection of structural stability (Christiano, 1992). Asymptotic alternatives proposed by Andrews (1993) and Andrews and Ploberger (1994) cannot be employed in the case of time series with unit-roots and/or for systems with one or more structural breaks in the marginal process (Hansen, 2000). A bootstrap approach to test for the null hypothesis of $i-1$ breaks against i breaks is here employed. We then stop the search when two consecutive estimated breaks are insignificant.

First, Equation (4) is estimated under the null hypothesis of $i-1$ breaks. The conventional F-statistic estimator is used to estimate breaks under the null and the statistic $\hat{\tau} = \max\{F_t\}$ is computed. Bootstrap errors are then generated (parametrically or non-parametrically), and the estimated parameters together with bootstrap errors are used to generate a recursive bootstrap sample. The resampling procedure is repeated B times and for each iteration $j = 1, \ldots, B$ the statistic $\tau_j^* = \max\{F_{t,j}\}$ is computed. The feasible bootstrap p-value is computed as:

$$p^*(\hat{\tau}) = \frac{1}{B}\sum_{j=1}^{B} I((\tau_j^*)^2 > (\hat{\tau})^2) \tag{5}$$

Finally, significant breaks are included in Equation (3), together with explanatory variables, and the sequential search is repeated for the conditional process.

From Equation (4), a preliminary analysis involves the proper selection of the functional form of the data generating process (DGP) depicting the series. Following Chris-

tiano (1992), we investigate the AR(p) form of the DGP for each marginal process under the null hypothesis of no breaks. Models are selected based on the analysis of the AR(p) fitted residuals. Results (not reported) suggest the use of $p = 4$ lags for both the inflation and the interest rate models and $p = 2$ lags for the real GDP model.[9]

We then estimate and test for structural breaks in marginal and conditional models through the bootstrap approach. For the former, each equation includes a constant, a trend and p lags of the dependent variable. For the latter, models are estimated as displayed in Equation (3), *i.e.* replacing regressors with their AR(p) marginal processes and their own structural breaks ($B^M_{n,i,t}$ and $B^L_{n,i,t}$). The lagged-dependent variable does not enter the equation, to avoid the problem of unit-roots and insignificance of regressors when the regressand is non-stationary. Results are reported in Table 2 and 3.

[Table 2 and 3 about here]

Bootstrap p-values for the null hypothesis of $i - 1$ against i breaks are computed through $B_p = 999$ random replications. Three bootstrap sampling methods are implemented in this paper: parametric, non-parametric with d.o.f. correction and non-parametric with the correction advocated by Weber (1984).

5.1 Breaks in the marginal processes

We now report and interprete results from structural breaks analysis on the determinants of money demand.

5.1.1 Interest rate

Most significant breaks in the interest rate cover the 1980s, a period marked by a high fiscal deficit and the end of the military government in 1983, many inflation and interest rates peaks (1981, 1985 and 1987) and the beginning of the hyperinflation episode in 1989. Two more significant breaks are estimated in 1976 and 1978, which coincide with the end of the third Peronist government and the military coup (1976) and the effects of

[9] Since all variables are non-stationary, autoregressive models are estimated in first differences. In addition, models showing a drift in levels are estimated without a constant in first differences, while trending variable regressions are estimated with a constant. Optimal lag-lenght selection for AR(p) models is based on the Ljung-Box test statistic and the Schwarz information criterion jointly. Results are available upon request.

the strong financial reform implemented in 1977 by authorities. Although the procedure is not able to detect accurately this financial liberalization, the effects of another financial reform are captured by a significant break in 1957. Finally, the two breaks in 1992 and 2003 are coherent with the full recover of Argentina in the aftermath of the hyperinflation and of the economic crisis (2001-2002).

5.1.2 Inflation

Two breaks are jointly significant under both the parametric and the non-parametric bootstrap. The first one (1989) coincides again with the beginning of the hyperinflation episode spanning from 1989 through 1990. The second one (1975) coincides with the abrupt adjustment in prices and tariffs set by the Minister of the Economy, Celestino Rodrigo, through a strong exchange rate devaluation. This adjustment (later called *Rodrigazo*) was decided to fix the distortion of relative prices, due to the regime of price controls in force since 1973, but caused a boom of inflation (followed by trade union strikes and slowdown of economic activity) with persistent effects until the 80s (monthly wholesales inflation registered a jump from 5% up to 44% between May and June 1975, when the adjustment was implemented). Another break is detected in 1996, but it cannot be rejected only by the parametric bootstrap-based test.

5.1.3 Real income

Although we have strong priors about regime changes in the real economic activity of Argentina, the sequential procedure is not able to identify any significant structural break even after many search stages. We can only find a significant break in 1989, *i.e.* the beginning of the hyperinflation episode. Using different model specifications did not improve results. We therefore accept the hypothesis of one single structural break.

5.2 Breaks in the money demand models

Given the set of estimated breaks in marginal models, test equations for conditional models take the following form:

$$\begin{aligned}(m-p)_t =& \mu_0 + \delta_0 t + \sum_{i=1}^{I_0}(\alpha_{0,i}B^M_{m-p,i,t} + \beta_{0,i}B^L_{m-p,i,t}) + \sum_{j=1}^{2}\gamma_{1,j}y_{t-j} \\ &+ (\eta_{1,1}B^M_{y,1,t} + \theta_{1,1}B^L_{y,1,t}) + \sum_{j=1}^{4}\gamma_{2,j}r_{t-j} + \sum_{i=1}^{10}(\eta_{2,i}B^M_{r,i,t} + \theta_{2,i}B^L_{r,i,t}) + \varepsilon_t \end{aligned} \tag{6}$$

$$\begin{aligned}(m-p)_t =& \mu_0 + \delta_0 t + \sum_{i=1}^{I_0}(\alpha_{0,i}B^M_{m-p,i,t} + \beta_{0,i}B^L_{m-p,i,t}) + \sum_{j=1}^{2}\gamma_{1,j}y_{t-j} \\ &+ (\eta_{1,1}B^M_{y,1,t} + \theta_{1,1}B^L_{y,1,t}) + \sum_{j=1}^{4}\gamma_{2,j}\pi_{t-j} + \sum_{i=1}^{2}(\eta_{2,i}B^M_{\pi,i,t} + \theta_{2,i}B^L_{\pi,i,t}) + \varepsilon_t \end{aligned} \tag{7}$$

Estimated breaks $B^M_{m-p,i,t}$ and $B^L_{m-p,i,t}$ are money demand's own breaks.

The upper panel in Table 3 reports results for the money demand model with the interest rate as opportunity cost. We have evidence of three structural breaks, in 1913, 1946 and 1959.[10] The first break (1913) coincides with the end of the gold-standard experience of Argentina, a relatively stable economic and monetary regime since the beginning of the century, that is the zenith of the so-called *Belle Epoque*. With respect to the second break (1946), the market and trade oriented economy and the financial development of the country were progressively reverted by nationalizations policies during the first Peron's mandate (1946-1951). High fiscal deficits, strong controls of the banking sector and inflationary pressures started to deteriorate the credibility in the financial sector, causing a disintermediation phenomenon which affected Argentina for most the second half of the century. The third break (1959) coincides with implementation of the Alsogaray's *Stabilization and Development Program* (July 1959), which aim was to raise economic growth (mainly through incentives to foreign investments and import substitution) and to stabilize prices (by fixing wages, prices and tariffs). The program took place during the mandate of President Frondizi (1958-1961) as a response to the explosion of prices due to

[10] Since some instability affected the estimate of the first relevant break in the output, we obtained the final sequence of breaks through the following procedure: *first*, significant breaks in 1959 and 1913 were detected; *second*, these breaks were considered as known and the first relevant break was estimated again; *finally*, the resulting relevant break (1946) was considered as known while relaxing the hypothesis over breaks in 1959 and 1913.

the second highest peak in the fiscal deficit of the central government ever experienced in Argentina (Della Paolera and Taylor, 2003).

The lower panel in Table 3 reports results for the money demand model with the inflation as opportunity cost. In addition to the breaks in 1913, 1946, 1959, we have evidence of significant breaks in 1986 and 1997. In particular, the 1997 break date may be interpreted as a signal for the retrieved monetization trend after the monetary turmoils of the previous decades, while the 1986 break date can be seen as the stabilizing effects of a macroeconomic program, the *Austral Plan* (June 1985). The program consisted in a set of stabilization policies against the stop-ad-go cycles of argentinean high-inflation, through a reduction of fiscal deficits, a gradual de-indexation of the economy and a prices-exchange rate-wages fixing commitment.

5.3 An interesting case of co-breaking

As emerges from the analysis depicted above, marginal models differ in the number and location of estimated breaks. It is worth noticing that this is not in contradiction with a structural model, because we are dealing with processes that are supposed, by the economic theory, to show common features with the whole economy, as well as idiosyncratic ones. However, the conditional model shows an interesting case of co-breaking, *i.e.* breaks which are statistically significant in marginal models become irrelevant features when money demand equation is estimated (Hendry and Massmann, 2007; Urga, 2007). In particular, this is the case of the shocks in the occurrence of the hyperinflation episodes of 1989-1990, the economic crisis of 2001-2002 or the reversion of the chronic demonetization of Argentina in act for a decade since the implementation of the *Convertibility Plan* (1991). Evidence of co-breaking can of course be detected through the Hendry and Massmann (2007) procedure, which is confirmed in this paper by searching for breaks in money considered as a marginal process. Results, reported in the lower panel of Table 2, suggest that real money balances are affected by three significant deterministic breaks, in 1959, 1975 and 1991. Note that the first break is detected in conditional models but not in marginal models, while the second one is also detected in the inflation process. The third break then vanish when linear combination of the series, such as conditional models, are considered. Insignificant estimated breaks in 1946 and 1986 appear nevertheless

significant in conditional models.

6 Estimating long-run money demand equations for Argentina

In this section we estimate money demand models and test for the presence of long-run relationships.

6.1 Estimator and tests

We estimate long-run equations through the dynamic GLS estimator (DGLS) developed by Stock and Watson (1993). The long-run equation includes k leads and lags of first-differenced regressors. We fix $k = 2$ for all models, which is sufficient to correct for potential endogeneity in a context of annual observations. Further, each money demand model includes dummies for the breaks estimated in the previous Section. Our estimator also embeds the Cochrane-Orcutt algorithm, to potentially correct for residual serial-correlation, and we set $l = 1$ the autoregressive error process. The estimated equation takes the following form:

$$(m-p)_t = \alpha + \delta t + \beta' X_t + \sum_{j=-k}^{k} \gamma_j' \Delta X_{t-j} + \sum_{i=1}^{I} (\eta_i B_{i,t}^M + \theta_i B_{i,t}^L) + \varepsilon_t, \tag{8}$$

where X_t is the matrix of regressors, k is the number of leads and lags involved in the DGLS estimation and $B_{i,t}^M$ and $B_{i,t}^L$ are the vectors of deterministic breaks. Standard mis-specification tests are applied to regression residuals. Under the hypothesis of stationarity of the error term, fast-double bootstrap p-values are computed through non-parametric (with d.o.f. correction) resampling (Davidson and MacKinnon, 2007). The number of generated bootstrap samples is $B_p = 9,999$.

6.2 Residual-based test for cointegration

To test for cointegration, we implement the KPSS residual-based test developed by Shin (1994), which assumes the null of stationarity of residuals (cointegration) against non-

stationarity. Although this test has the advantage of being designed for leads and lags regressions, it does not allow for structural breaks in its original version. Carrion-i-Silvestre and Sansò (2006) first studied the asymptotic and finite-sample performance of this test statistic when one break is included. The test statistic takes the multivariate KPSS form:

$$SC = T^{-2}\hat{\omega}^{-2}\sum_{t=1}^{T}\left(\sum_{j=1}^{t}\hat{\varepsilon}_j\right)^2 \tag{9}$$

where $\hat{\varepsilon}_t$, $\{\hat{\varepsilon}_t\}_{t=1}^{T}$, are the estimated DGLS residuals from Equation (8) and $\hat{\omega}^2$ is a consistent estimate of the long-run variance of $\{\varepsilon_t\}_{t=1}^{T}$, the latter being computed through the Bartlett kernel and a modified automatic bandwidth selection method as suggested in Kurozumi (2002).[11] Nevertheless, the distribution of this statistic is dependent on the deterministic structure and the number of regressors involved in the model (Shin, 1994), as well as the type of break (constant, trend, stochastic parameter) and its relative position in the sample. In addition, it can also be shown that the asymptotic distribution crucially depends on the absolute number of breaks. Since our model involves up to 5 deterministic breaks in the long-run equations, we cannot use critical values provided by Carrion-i-Silvestre and Sansò (2006). In this paper, we compute instead fast-double bootstrap p-values for the multiple breaks SC statistic (MSC) through non-parametric (with d.o.f. correction) resampling.[12]

6.3 Results

Results from the model including the interest rate as opportunity cost are reported in Table 4. We first run the unconstrained regression (7), including all breaks estimated in Section 5. Results are reported in column (i) of Table 4. Coefficients, except for the

[11] Kurozumi (2002) proposes the following modified Andrews (1991) automatic bandwidth selection method: $\tilde{M} = \min\left(1.1447\{4\hat{\rho}^2T/[(1+\hat{\rho})^2(1-\hat{\rho})^2]\}^{1/3}, 1.1447\{4k^2T/[(1+k)^2(1-k)^2]\}^{1/3}\right)$, where $\hat{\rho}$ is the AR(1) coefficient of $\hat{\varepsilon}_t$ and k can take an efficient value between 0.7 and 0.9. In this paper we select $k = 0.8$, as suggested by Carrion-i-Silvestre and Sansò (2006).

[12] Monte Carlo simulations show that the cointegration test under analysis, combined with the DGLS estimator, has reasonable size and power properties even in a multiple breaks framework. Results are available upon request.

trend break in 1913, are statistically significant at 1, 5 and 10%. In particular, income and interest rate parameters are both highly significant and with expected signs. The former is positive and close to unity, in contrast with the preliminary cointegration results reported in Table 1, while the latter is around -1.6. These results suggest that the long-run demand for real money balances in Argentina is driven by proportional changes in income (unit homogeneity hypothesis), while portfolio motives lead to strong adjustments of the desired level of money to variations in the yield of alternative assets. It is worth noticing that results are in line with those reported for industrialized countries (see, among others, Lucas, 1988, and Hendry and Ericsson, 1991). However, the inclusion of structural breaks is here a crucial long-run issue. Estimated residuals have good statistical properties, except for some evidence of ARCH effects at lag 2 and 4. In addition, the residual-based MSC test cannot reject the null hypothesis of cointegration.

[Table 4 about here]

In column (ii), we report results when we constrain the insignificant break coefficient to zero and we re-estimate the model. Results are very similar to those in column (i). We then test for the income homogeneity hypothesis through a Wald test. The null hypothesis being not rejected (p-value = 0.77), we estimate regression (7) including this restriction over the real GDP elasticity. Results, reported in column (iii), are very similar to those reported in (i) and (ii). As expected, the level break in 1913 is negative (the end of the gold-standard system), while the level break in 1946 is positive (a transitory effect of the redistributive peronist policies). The level break in 1959 has a negative sign, which can be associated to the fall of real balances due to sudden inflationary shocks which took place in that year. It is worth noticing that the coefficient for the trend break in 1946 has a negative sign, which identifies the beginning of the long-lasting demonetization of Argentina until the 1990s. The positive sign of the coefficient for the trend break in 1959 partially compensate the fall of money since the half of the 20th century (the sum of the coefficient is still negative). In columns (iv) and (v) we report results when we allow for a second lag in the Cochrane-Orcutt procedure, in order to check for the robustness of previous estimates. As expected, results are quantitatively similar to those reported in (ii) and (iii). Again, residual analysis does not highlight major problems in our specifications,

except for some evidence of ARCH effects at lag 4, and cointegration cannot be rejected. Error-correction terms from (iii) and (iv), *i.e.* the estimated deviations of real money balances from the long-run equilibrium, are plotted in Figure 2(a).

[Figure 2 about here]

Results from the model including the inflation rate as opportunity cost are reported in Table 5. We first run the unconstrained regression (7), and results are reported in column (i). Income and inflation coefficients are highly significant and with expected signs, while many level and trend break parameters appear insignificant. Unlike the previous specification, here the income coefficient (1.48) is significantly higher than unity and the inflation coefficient (-0.59) is quite lower than the interest rate coefficient. The latter issue confirms results showed in Table 1 about the existence of a weak Fisher parity between interest rate and inflation. Estimated residual have good statistical properties, except for some autocorrelation at lag 2, and the MSC test cannot reject the null hypothesis of cointegration.

[Table 5 about here]

In columns (ii) and (iii), we report results when we look for a more parsimonious specification by constraining insignificant breaks to zero. It is worth noticing that estimated income and inflation coefficients slightly differ from the baseline model in (i): the former tends to be lower, while the latter stays in the range of the confidence interval. We test for the income homogeneity hypothesis. Although the high income coefficient, the null hypothesis cannot be rejected in both specifications (ii) and (iii) (p-value = 0.43 and 0.28, respectively). We then estimate constrained specifications (iv) and (v). Results are similar to those reported in columns (ii) and (iii). In particular, we estimate the inflation coefficient in a range between -0.6 and -0.7. Residual tests do not point out any statistical problem and cointegration hypothesis is never rejected. Error-correction terms from specifications (iv) and (v) are plotted in Figure 2(b). As before, the level break in 1913 is negative and the level break in 1946 is positive. Again, the level break in 1959 has a negative sign, as well as the trend break in 1946. However, the trend break in 1959 is not significant in all specifications.

7 Short-run dynamics

To model dynamics, we follow Engle and Granger (1987) and we estimate the following general autoregressive specification for the growth rate of money:

$$\Delta(m-p)_t = \alpha_0 + A(L)\Delta(m-p)_t + \boldsymbol{C(L)\Delta x_t} + \gamma ECT_{t-1} + \boldsymbol{\phi' \Delta B^{M,L}} + u_t \tag{10}$$

where $A(L)$ is a polynomial vector and $\boldsymbol{B(L)}$ is a polynomial matrix, $\boldsymbol{\Delta x} = \{\Delta y, \Delta r, \Delta \pi\}'$ is the vector of regressors, $\boldsymbol{\Delta B^{M,L}}$ is the vector of structural breaks and ECT is the error-correction term.[13] We estimate Equation (10) using up to 4 lags. A parsimonious single-equation error-correction model is then obtained by dropping insignificant variables and following the automated General-to-Specific reduction approach (Krolzig and Hendry, 2001; Hendry and Krolzig, 2005; Doornik, 2007). In addition, we report standard diagnostic tests and we allow for outliers correction through the introduction of blip-dummy variables for large residuals. 1, 5 and 10% significance level (denoted by *** , ** and * , respectively) is reported, while t-statistics for coefficients significance are based on heteroscedasticity and autocorrelation consistent standard errors (Andrews, 1991). As for the long-run specification, two models involving either the interest rate or the inflation rate are estimated. Error-correction terms from specification (iii) in Table 4 and from specifications (iv) and (v) in Table 5 are used. Following this strategy, we report in Equation (11) our preferred dynamic equation for the model involving the growth rate of GDP and changes in the interest rate:

$$\begin{aligned}\Delta(m-p)_t = \; & 0.11^{*}\Delta(m-p)_{t-1} - 0.16^{***}\Delta(m-p)_{t-2} + 0.79^{***}\Delta y_t - 0.84^{***}\Delta r_t \\ & - 0.47^{***}ECT_{t-1} - 0.21^{***}\Delta 1913^{M} + 0.13^{***}\Delta 1946^{M} - 0.29^{***}\Delta 1959^{M}\end{aligned} \tag{11}$$

T=102 R^2=0.76 $\hat{\sigma}$=9% SBC=-1.69
AR(4)=1.69 ARCH(4)= 0.79 NORM= 0.66 HET=1.09

where AR(p) is the Breusch-Godfrey LM test for residual serial-correlation, ARCH(p) is the Engle-Bollerslev test for the presence of ARCH effects, NORM is the Doornik-Hansen test for normality of residuals and HET is the White test for heteroscedasticity. Although only a small set of regressors enters the final dynamic model, the equation passes all the

[13] Breaks enter the short-run model in first-difference because they were included into the long-run equation.

diagnostic tests. Our preferred model includes 3 (differenced) level breaks (1913, 1946 and 1959) and one outlier for the year 1989. We finally report our preferred dynamic equations for the model involving the growth rate of GDP and the acceleration in inflation. Equation (12) and (13) use specifications (iv) and (v) from Table 5, respectively:

$$\begin{aligned}\Delta(m-p)_t = \quad & 0.17^{***}\Delta(m-p)_{t-1} + 0.84^{***}\Delta y_t - 0.35^{***}\Delta\pi_t - 0.28^{***}ECT_{t-1} \\ & - 0.25^{***}\Delta 1913^M + 0.13^{***}\Delta 1946^M - 0.11^{***}\Delta 1959^M \\ & - 0.03^{**}\Delta 1946^L + 0.06^{***}\Delta 1997^L \end{aligned} \tag{12}$$

T=102 R^2=0.80 $\hat{\sigma}$=8% SBC=-1.82
AR(4)=2.27* ARCH(4)= 2.41* NORM= 0.62 HET=1.02

$$\begin{aligned}\Delta(m-p)_t = \quad & 0.18^{***}\Delta(m-p)_{t-1} + 0.71^{***}\Delta y_t - 0.34^{***}\Delta\pi_t - 0.35^{***}ECT_{t-1} \\ & - 0.25^{***}\Delta 1913^M + 0.12^{***}\Delta 1946^M - 0.17^{***}\Delta 1959^M \end{aligned} \tag{13}$$

T=102 R^2=0.81 $\hat{\sigma}$=8% SBC=-1.91
AR(4)=1.58 ARCH(4)= 0.93 NORM= 2.19 HET=1.61

Only one lag for the money growth term enters the dynamic equation. Level breaks in 1913, 1946 and 1959 are again significant, together with two additional trend breaks (1946 and 1997) for the specification (iv). For the latter, we have some evidence of residual autocorrelation and conditional heteroscedasticity, although at only 10% significance level. One outlier for the year 1989 enters both specifications, while an additional outlier for the year 1974 enters Equation (13).

Turning to the dynamic adjustment of real money balances, the error-correction term is significant in all preferred equations and with the expected (negative) sign. It is worth noticing that the speed of adjustment towards the long-run equilibrium is stronger in Equation (11). This is in line with results obtained in Section 6, where the estimated long-run interest rate coefficient is significantly larger than the long-run inflation coefficient. The overall high explained variance, which is close to 80%, suggests that our simple models are good enough to fit actual real money data. To further explore this issue, we plot fitted and actual values of $\Delta(m-p)_t$ from Equations (11), (12) and (13) in Figure 3(a), 3(b) and 3(c), respectively.

[Figure 3 about here]

8 Implications for the welfare cost of inflation

Following the recent theoretical and empirical contributions of Lucas (2000), Chadha *et al.* (1998), Cysne (2008, 2009) and Ireland (2009), we compute welfare cost functions for Argentina. In the spirit of the work of Bailey (1956), Lucas (2000) defines the welfare cost of inflation as the reduction in consumption needed to compensate for the higher utility due to a larger stock of money balances. Compared with the static approach of Bailey (1956), the "compensating variation approach" for the welfare cost of inflation has macroeconomic theoretical foundations arising from general-equilibrium frameworks, such as the Sidrauski (1967) and the McCallum and Goodfriend (1987) shopping-time models. Nevertheless, Lucas (2000) shows that measures issued from the Bailey's (1956) formula can be regarded as an approximation to the general-equilibrium measures emerging from these models, providing then an economic rationale for the application of the former approach. Thus, in this Section let's set the welfare cost of inflation, $w(r)$, as the area under the inverse demand function, $\psi(m)$, that could be gained by reducing the interest rate from r to zero:

$$w(r) = \int_{m(r)}^{m(0)} \psi(x)dx = \int_0^r m(x)dx - rm(r), \tag{14}$$

where $w(r)$ and $m(r)$ - the money demand function - have the dimension of a ratio to income. For the semi-log function $m(r) = B\exp(-\xi r)$ (see also Ireland (2009)), Equation (14) implies:

$$w(r) = \frac{B}{\xi}\left[1 - (1 + \xi r)\exp(-\xi r)\right], \tag{15}$$

where ξ is the estimated interest rate semi-elasticity and B is the exponential of the deterministic parameter of the model. Two differences arise in our framework with respect to the model proposed by Lucas (2000). First, we formally deal with a tantamount to the semi-log specification of the money demand by using $r_t = \ln(1 + R_t)$, a feature that we have to account for when solving the integral in Equation (14). Second, deterministic structural breaks estimated in Sections 5 and 6 have a clear impact on the shape of the welfare cost function. For the former, we can show that results do not substantially differ

from the pure semi-log specification. For the latter, the issue of time-varying parameters is nevertheless difficult to reconcile with the steady-state results advocated by Lucas (2000). We then propose to analyze the welfare cost of inflation in Argentina regime-by-regime, keeping the long-run interest rate semi-elasticity fixed and computing the deterministic term B in Equation (15) according to the sub-sampling location of estimated breaks. We do not nevertheless consider estimated regimes as independent, but we compute B_i, where i is a regime, as the interaction of past $(i-p)$ and current regimes.

For our money demand function $m(R) = C(1+R)^{-\eta}$, where η is the long-run interest rate semi-elasticity in absolute value, C is the deterministic part of the model estimated in Equation (8), and the homogeneity restriction with respect to income could not be rejected (Table 4), Equation (14) implies:

$$\begin{aligned} w(R) = \int_0^R m(x)dx - Rm(R) &= \frac{C}{1-\eta}(1+R)^{1-\eta} - \frac{C}{1-\eta} - RC(1+R)^{-\eta} \\ &= \frac{C}{1-\eta}\left[(1+R)^{1-\eta} - 1 - R(1-\eta)(1+R)^{-\eta}\right] \\ &= \frac{C}{1-\eta}\left[(1+R)^{-\eta}(1+R\eta) - 1\right]. \end{aligned} \tag{16}$$

From Table 4, column (iii), we set $\eta = 1.61$. In addition, we can compute C for 4 sub-periods, according to the number of estimated breaks. Let's rewrite C as follows:

$$C_i = \exp\left[\sum_i (\alpha_i + \delta_i t(i))\right] = \prod_i A_i \exp(\delta_i t(i)) = \prod_i A_i (1+g_i)^{t(i)},$$

where $g_i = \exp(\delta_i) - 1$, for each sub-period $i = 1, \ldots, 4$. Values for C under different estimated regimes are reported in Table 6.

The empirical exercise requires the computation of the welfare cost of inflation from Equation (16). In particular, we are interested in the standard interpretation of numerical values for $w(R)$, that is the fraction of income necessary to compensate agents from living under positive steady-state nominal interest rate policies rather than the deflationary policies (zero interest rate rule) advocated by Friedman (1969). Table 6 reports welfare cost estimates for Argentina (in percentage of income) for different steady-state real interest rates (R_{ss}), optimal annual inflation rates ($\pi^\star$) and money demand regimes.

[Table 6 about here]

As expected, results put high welfare cost of pursuing a policy of price stability or moderate inflation during the last regime (1959-2006), while for the other regimes we obtain similar measures (mainly during the second and third regime). For instances, if we assume the steady-state interest rate at 5% across regimes, we observe that the cost of policies driving zero inflation is between 0.14 and 0.18 percent of income for the period 1900-1958 and 0.25 for the last regime. However, the cost of 10% inflation tends to increase rapidly for all regimes, with a range of 0.93-1.66 percent of income. Figure 4 resumes graphically these findings, showing the path of welfare cost functions under different regimes and inflationary policies.

[Figure 4 about here]

These results are significantly higher than those reported by Lucas (2000) and Ireland (2009) for the United States, highlighting the strong welfare improvement Argentine authorities would have obtained during the 20th century by running sensible monetary policies, *i.e.*, if authorities had not recourse to seigniorage in support to high and chronic fiscal deficits. It is worth noticing that the form of the estimated money demand function (semi-log) and the quite high value of the deterministic term C across regimes have an impact on the shape of welfare cost functions. In addition, recent literature (Bali, 2000; Cysne and Tuchick, 2010) highlights the potential overestimation (underestimation) bias of the welfare cost of inflation when (alternative) interest-bearing deposits are disregarded and a specific functional form (double-log) for the demand for money is taken. Finally, interesting extensions are also represented by the multi-dimensional frameworks involving many currencies and interest-bearing deposits (Jones *et al.*, 2004; Cysne and Turchick, 2009), as well as Divisia index (Cysne, 2003).

9 Concluding remarks

In this paper, we estimated money demand in Argentina over a long time span, from 1900 to 2006. When we tested for cointegration between real M1 balances, real GDP, inflation rate and interest rate, we showed that cointegration results are not robust when structural breaks are neglected. We then tested and dated for deterministic structural

breaks in marginal (real GDP, inflation rate and interest rate) and conditional (real M1 balances) processes using a sequential search procedure. Results suggested the presence of significant breaks both in marginal and conditional processes, consistent with extreme financial episodes and macroeconomic shocks. These breaks coincides with the end of the gold-standard experience of Argentina (1913), the beginning of Peronist policies (1946), the explosion of fiscal deficits and prices (1959), the attempts of stabilization as a consequence of the *Austral Plan* (1986) and the re-monetization process during the 90s (1997). We estimated long-run money demand equations embedding structural breaks by DGLS. Estimated parameters are highly significant and correctly signed. Quite interestingly, the quantitative theory hypothesis of income homogeneity could not be rejected. Conclusive evidence on the existence of long-run money demand equilibrium in Argentina is provided by multiple breaks cointegration tests. Although the presence of extreme adjustments which took place in 1975 and in 1989-1990, estimated cointegrating vectors appear then generally stable once the breaks are taken into account. In addition, the short-run analysis highlighted the high speed at which agents re-adjusted excess money holdings towards the desired level of liquidity. These findings are in line with those reported in the literature of industrialized countries (Lucas, 1988; Hendry and Ericsson, 1991; Baba et al., 1992; Muscatelli and Spinelli, 2000), but add new elements to the understanding of the long term monetary dynamic in emerging economies. Finally, we study the implication of our empirical results for the welfare cost of inflation in Argentina. Our results suggest that welfare cost have been high, especially if compared to US figures, in particular at the end of the last century.

Some interesting issues arise. The seignorage cost of protracted monetary disorder can be explored by exploiting the results reported in this paper. In addition, the moderate re-monetization process in Argentina after decades of financial chronic disorder, as suggested by data and by a quite slower speed of adjustment in the post-1991 period, may also point out the effects of monetary hysteresis phenomena which need to be further explored. Indeed, results reported in this paper call for a comparative investigation extended to other Latin American countries. We leave these issues for future research.

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Data Appendix

Many sources were used to build the dataset. The main one is Veganzones and Winograd (1997), who collected data until 1992. From 1993 to 2006, the dataset is fully updated by the authors of the present work, while time series before 1993 are corrected and re-scaled.

- **Money** (*Nominal M1*) :
  1900-1913 : DellaPaolera (1988) and Anuario Geografico Argentino;
  1913-1935 : Anuario Geografico Argentino (various issues);
  1935-2006 : Banco Central de la Republica Argentina (*memorias* and web sources).

- **Gross Domestic Product** (*Nominal*) :
  1900-1993 : Della Paolera and Taylor (2003);
  1993-2006 : International Monetary Fund (International Financial Statistics).

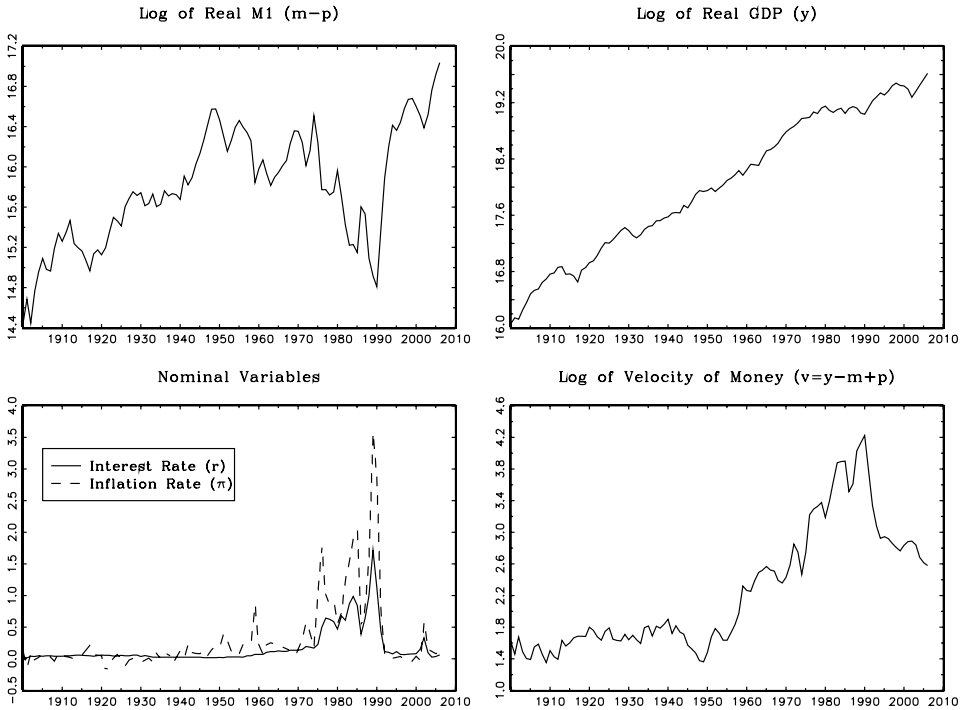
- **Interest Rate** (*Various Rates*) :
  1900-1913 : Della Paolera (1988);
  1913-1938 : Anuario Geografico Argentino;
  1938-2006 : Banco Central de la Republica Argentina (*memorias* and web sources).

- **Price Index** (*Wholesale Price Index, base-year 1993*) :
  1899-1956 : Della Paolera (1988), Economic Commission for Latin America (ECLA), United Nations and Instituto Nacional de Estadistica y Censos (INDEC).
  1956-2006 : Banco Central de la Republica Argentina (web sources).

- **Implict GDP Price Deflator** (*Base-year 1993*) :
  1900-1993 : Della Paolera and Taylor (2003);
  1993-2006 : International Monetary Fund (International Financial Statistics).

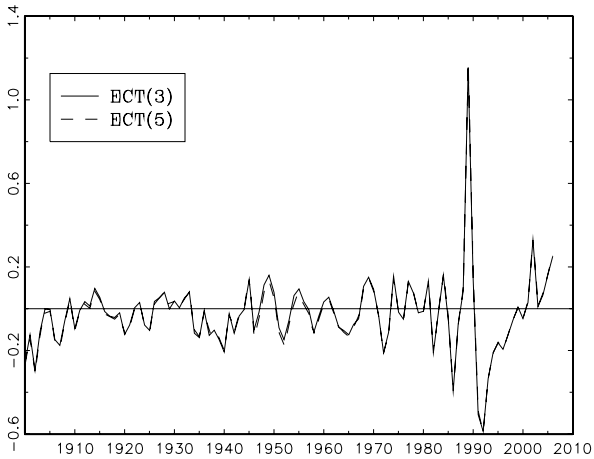
- **Gross Domestic Product** (*Real*) :
  1900-1980 : Because of rebasements in 1970 and 1986, an adjustment was needed to harmonize the real GDP series. This was done by linking the pre-1980 series to the post-1980 series. We applied growth rates of the real GDP, originally computed as the nominal GDP deflated by the implicit price index, to the 1980 value (issued from Della Paolera and Taylor, 2003) and we adjusted the series backward;
  1980-1993 : Della Paolera and Taylor (2003);
  1993-2006 : International Monetary Fund (International Financial Statistics).

Figure 1: Argentina data (1900-2006)

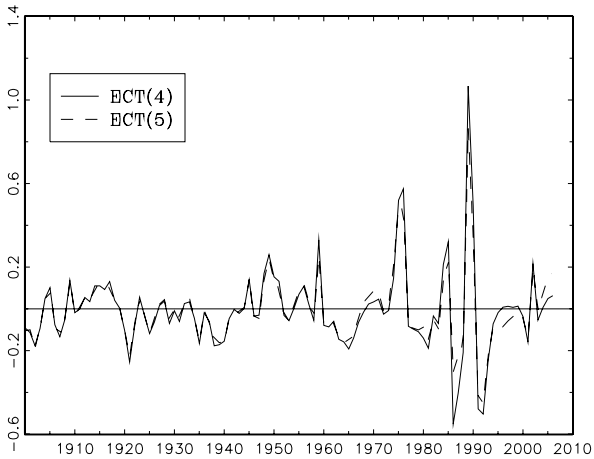


Notes: Money balances and GDP are expressed in current Argentine pesos.
Sources: see Data Appendix

Figure 2: Deviations from the long-run equilibrium for real balances.

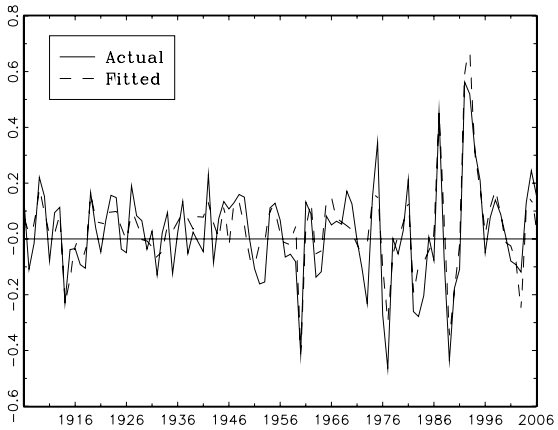


(a) Error-correction term (ECT) (3) and (5) from Table 4.

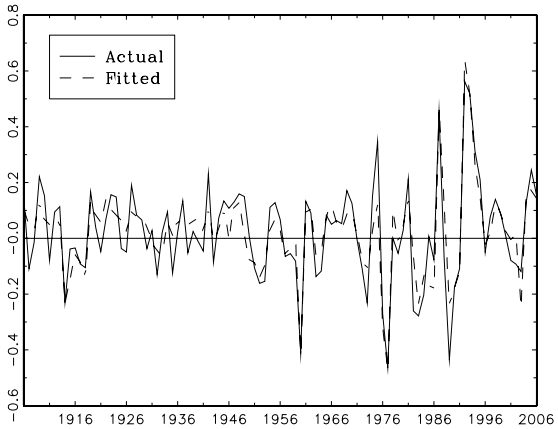


(b) Error-correction term (ECT) (4) and (5) from Table 5.

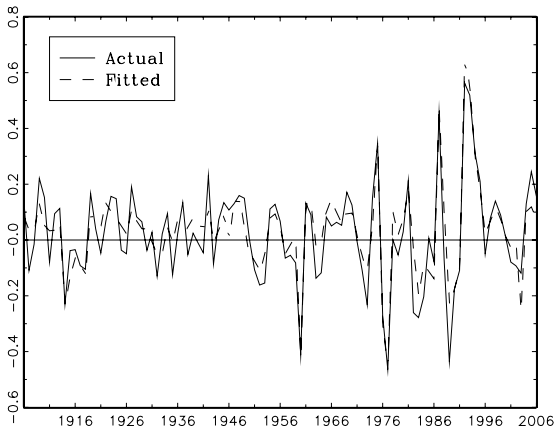
Figure 3: Actual and Fitted Values for $\Delta(m-p)_t$



(a) Equation (11).

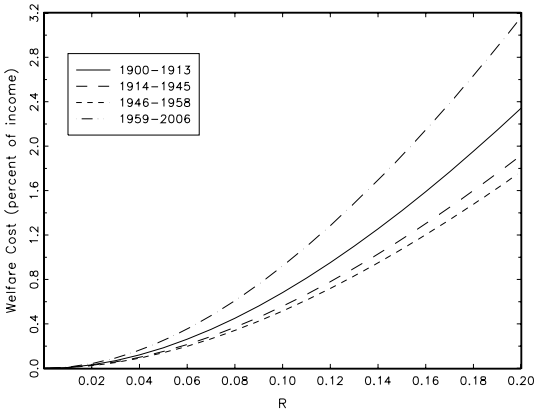


(b) Equation (12).

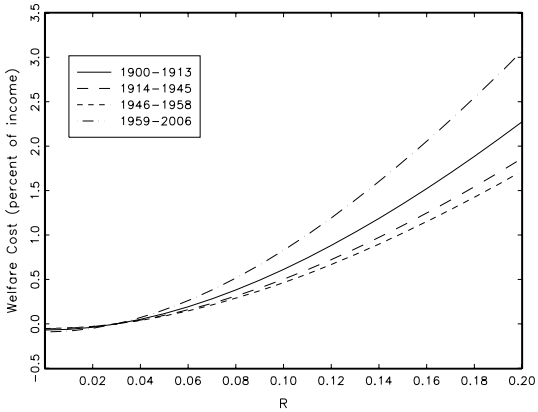


(c) Equation (13).

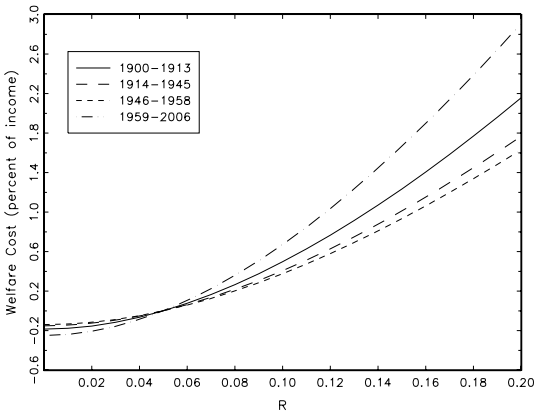
Figure 4: Welfare Cost Functions



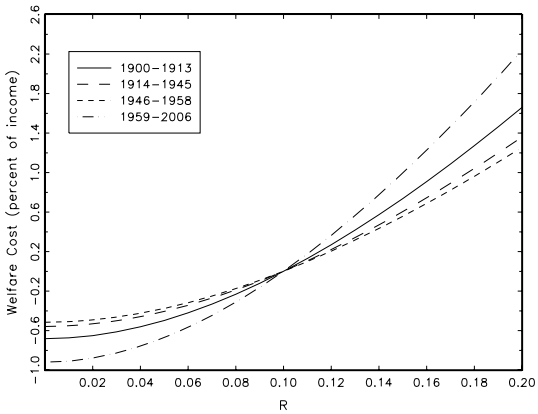
(a) $w(R)$



(b) $w(R) - w(0.03)$



(c) $w(R) - w(0.05)$



(d) $w(R) - w(0.10)$

Table 1: Unit-root and cointegration tests

A. Unit-Root Tests

Variable	ADF^{GLS}	MZ_T^{GLS}	5% cv	MP_T^{GLS}	5% cv
$(m-p)^+$	-1.96	-1.92	(-2.91)	12.09	(5.48)
y^+	-1.71	-1.60	(-2.91)	15.65	(5.48)
π	-2.04^{**}	-1.93	(-1.98)	3.28	(3.17)
r	-1.60	-1.55	(-1.98)	5.12	(3.17)

B. Cointegration Tests

	Real Interest Rate (r, π)			Money Demand $(m-p, y, r, \pi)$		
Test	Statistic	5% cv	p-value	Statistic	5% cv	p-value
$J_\mu^{Trace}(0)$	27.9	20.3	< 0.01	62.4	47.8	0.01
$J_\mu^{Max\lambda}(0)$	22.1	15.9	< 0.01	29.1	27.6	0.03
$Q_\mu^f(0)$	-68.2	-30.9	< 0.01	-59.9	-39.3	< 0.01
$J_\mu^{Trace}(1)$	-	-	-	33.4	29.8	0.02
$J_\mu^{Max\lambda}(1)$	-	-	-	25.4	21.1	0.01
$Q_\mu^f(1)$	-	-	-	-59.9	-22.5	< 0.01

C. Estimated Cointegrating Vectors

	Money Demand (Model 1)		Money Demand (Model 2)	
Variable	CE 1	CE 2	CE 1	CE 2
$(m-p)$	1	-	1	-
y	-0.5	-	-0.5	-
r	2.0^{***}	1	-	1
π	-	-0.6^{***}	1.16^{***}	-0.57^{***}

LR Test for Restrictions (Model 1): $\chi^2(3)$=5.38 (p-value=0.14)
LR Test for Restrictions (Model 2): $\chi^2(4)$=5.66 (p-value=0.22)

Note: In Panel A, $^+$ denotes the inclusion of a constant and a linear trend into the test equation. J^{Trace} and $J^{Max\lambda}$ are the Johansen (1991) trace and maximum eigenvalue tests, respectively, computed from VAR(2)'s in levels according to the Hannan-Quinn criterion. MacKinnon et al. (1999) exact p-values are computed. Models include a constant in the cointegration equation and into the VAR. Q^f is the Stock and Watson (1988) common trend statistic. Test equations include a constant term. In parentheses, the number of cointegrating vectors under the null hypothesis. Maximum likelihood estimates for cointegrating models and likelihood ratio tests for binding restrictions are reported in lower panels.
*** and ** denote 1 and 5% significance level, respectively.

Table 2: Structural breaks in marginal models

	significant breaks						
r_t	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)
Date	1989	1987	1976	1983	1981	2003	1985
PB p-value	0.00	0.00	0.00	0.00	0.00	0.00	0.00
NPB(1) p-value	0.00	0.00	0.00	0.01	0.04	0.01	0.01
NPB(2) p-value	0.00	0.00	0.01	0.01	0.02	0.00	0.01

		significant breaks					
	$\cdots$	(viii)	(ix)	(x)	(xi)	(xii)	(xiii)
Date	$\cdots$	1992	1957	2001	1978	1933	1974
PB p-value	$\cdots$	0.00	0.01	0.27	0.00	0.21	0.47
NPB(1) p-value	$\cdots$	0.00	0.02	0.63	0.00	0.21	0.51
NPB(2) p-value	$\cdots$	0.00	0.02	0.73	0.00	0.22	0.48

	significant breaks				
π_t	(i)	(ii)	(iii)	(iv)	(v)
Date	1996	1989	1975	1986	1984
PB p-value	0.02	0.00	0.00	0.07	0.06
NPB(1) p-value	0.13	0.00	0.00	0.18	0.14
NPB(2) p-value	0.14	0.00	0.00	0.13	0.13

y_t	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)
Date	1964	1913	1918	1977	1931	2002	1989
PB p-value	0.22	0.88	0.54	0.35	0.21	0.07	0.03
NPB(1) p-value	0.24	0.89	0.54	0.31	0.25	0.12	0.04
NPB(2) p-value	0.24	0.88	0.54	0.36	0.24	0.11	0.04

	significant breaks				
$(m-p)_t$	(i)	(ii)	(iii)	(iv)	(v)
Date	1975	1991	1959	1946	1986
PB p-value	0.05	0.01	0.03	0.30	0.21
NPB(1) p-value	0.05	0.00	0.05	0.44	0.32
NPB(2) p-value	0.05	0.00	0.05	0.40	0.34

Note: PB is the parametric bootstrap; NPB(1) is the non-parametric bootstrap with d.o.f. correction; NPB(2) is the non-parametric bootstrap with the correction proposed by Weber (1984). Bootstrap samples are generated through 999 random replications. The order of the estimated breaks is sorted as in outcome of the sequential procedure. The lower panel shows the output from the sequential search when the series $(m-p)_t$ is considered as a marginal process.

Table 3: Structural breaks in conditional models

	significant breaks				
$(m-p)_t, y_t, r_t$	(i)	(ii)	(iii)	(iv)	(v)
Date	1946	1959	1913	1974	1972
PB p-value	0.00	0.05	0.06	0.71	0.56
NPB(1) p-value	0.00	0.05	0.06	0.77	0.62
NPB(2) p-value	0.00	0.07	0.06	0.77	0.57

	significant breaks						
$(m-p)_t, y_t, \pi_t$	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)
Date	1959	1946	1986	1913	1997[1]	1972	2003
PB p-value	0.00	0.00	0.01	0.03	0.00	0.28	0.14
NPB(1) p-value	0.00	0.00	0.02	0.02	0.01	0.29	0.16
NPB(2) p-value	0.00	0.00	0.02	0.02	0.01	0.28	0.14

Note: See Table 2. [1] the break is detected when the 1996 date is excluded from the marginal process of inflation.

Table 4: Long-run equation for the model $(m-p, y, r)$

	(i)	(ii)	(iii)	(iv)	(v)
y	0.92^{***}	0.94^{***}	1	1.03^{***}	1
r	-1.59^{***}	-1.59^{***}	-1.61^{***}	-1.62^{***}	-1.62^{***}
1913^M	-0.23^{**}	-0.21^{***}	-0.21^{***}	-0.20^{***}	-0.20^{***}
1913^L	-0.01	-	-	-	-
1946^M	0.32^{***}	0.31^{***}	0.31^{***}	0.35^{***}	0.35^{***}
1946^L	-0.03^{***}	-0.03^{***}	-0.03^{***}	-0.03^{***}	-0.03^{***}
1959^M	-0.37^{***}	-0.37^{***}	-0.38^{***}	-0.39^{***}	-0.38^{***}
1959^L	0.02^{*}	0.02^{*}	0.02^{**}	0.02^{**}	0.02^{**}
Constant	-0.05	-0.32	-1.30	-1.85	-1.29
Trend	0.009	-0.001	-0.003	-0.005	-0.004
Obs.	102	102	102	102	102
k	2	2	2	2	2
l	1	1	1	2	2
χ^2 Restrictions	-	-	0.084	-	0.045
			[0.77]	-	[0.83]
MSC	0.0236	0.0242	0.0247	0.0277	0.0279
	[0.36]	[0.39]	[0.36]	[0.13]	[0.12]
AR (1)	[0.41]	[0.38]	[0.37]	[0.93]	[0.85]
AR (2)	[0.11]	[0.11]	[0.10]	[0.99]	[0.99]
AR (4)	[0.21]	[0.19]	[0.18]	[0.98]	[0.98]
J-B	[0.61]	[0.61]	[0.62]	[0.99]	[0.99]
B-P	[0.24]	[0.22]	[0.21]	[0.11]	[0.12]
ARCH (1)	[0.10]	[0.10]	[0.11]	[0.38]	[0.35]
ARCH (2)	[0.01]	[0.01]	[0.01]	[0.10]	[0.11]
ARCH (4)	[0.01]	[0.01]	[0.01]	[0.02]	[0.02]

Note: DGLS estimates. Upper-scripts M and L are break in the mean and in the linear trend, respectively. AR(p) is the Breusch-Godfrey statistic, B-P is the Breusch-Pagan statistic, J-B is the Jarque-Bera statistic, ARCH(p) is the Engle-Bollerslev statistic, MSC is the residual-based test for cointegration with multiple breaks. Standard errors are reported in parentheses. Fast-double bootstrap p-values in square brackets, simulated through a non-parametric sampling with 9,999 random replications. Asymptotic p-values for linear restrictions hypothesis testing in square brackets.

*** , ** and * denote 1, 5 and 10% significance level, respectively.

Table 5: Long-run equation for the model $(m-p, y, \pi)$

	(i)	(ii)	(iii)	(iv)	(v)
y	1.48***	1.23***	1.32***	1	1
π	-0.59***	-0.69***	-0.55***	-0.70***	-0.59***
1913^M	-0.21**	-0.23***	-0.22***	-0.24***	-0.23***
1913^L	0.006	-	-	-	-
1946^M	0.27***	0.29***	0.33***	0.29***	0.33***
1946^L	-0.02**	-0.03***	-0.04***	-0.03***	-0.03***
1959^M	-0.30***	-0.31***	-0.26***	-0.29***	-0.26***
1959^L	-0.01	-	-	-	-
1986^M	0.20*	0.25***	-	0.21**	-
1986^L	0.04	-	0.06***	-	0.04***
1997^M	-0.04	-	-	-	-
1997^L	0.04	0.07	-	0.06***	-
Constant	-9.25*	-5.15	-6.66	-1.44	-1.43
Trend	-0.021	-0.007	-0.010	0.001	0.001
Obs.	102	102	102	102	102
k	2	2	2	2	2
l	1	1	1	1	1
χ^2 Restrictions	-	-	-	0.62	1.16
				[0.43]	[0.28]
MSC	0.0313	0.0345	0.0361	0.0327	0.0345
	[0.07]	[0.18]	[0.13]	[0.22]	[0.15]
AR (1)	[0.24]	[0.17]	[0.15]	[0.11]	[0.07]
AR (2)	[0.04]	[0.03]	[0.02]	[0.02]	[0.01]
AR (4)	[0.17]	[0.14]	[0.08]	[0.12]	[0.06]
J-B	[0.46]	[0.45]	[0.47]	[0.45]	[0.47]
B-P	[0.24]	[0.31]	[0.06]	[0.28]	[0.05]
ARCH (1)	[0.32]	[0.25]	[0.25]	[0.15]	[0.09]
ARCH (2)	[0.16]	[0.12]	[0.22]	[0.17]	[0.19]
ARCH (4)	[0.30]	[0.26]	[0.51]	[0.31]	[0.42]

Note: See Table 4.

Table 6: Welfare cost estimates (in percentage of income)

Regimes	C	R_{ss}	$\pi^\star = 0\%$	$\pi^\star = 5\%$	$\pi^\star = 10\%$	$\bar{\pi}_i$
		3%	0.07	0.38	1.03	0.19
I: 1900-1912	1	5%	0.18	0.49	1.23	0.27
		10%	0.68	0.74	1.66	0.42
		3%	0.06	0.31	0.84	0.21
II: 1913-1945	0.8187	5%	0.15	0.41	1.01	0.32
		10%	0.56	0.60	1.36	0.50
		3%	0.05	0.29	0.78	2.90
III: 1946-1958	0.7558	5%	0.14	0.37	0.93	3.23
		10%	0.51	0.56	1.25	3.90
		3%	0.09	0.51	1.39	46.43
IV: 1959-2006	1.3498	5%	0.25	0.67	1.66	46.68
		10%	0.92	0.99	2.24	47.07

Note: Welfare cost figures are expressed in %. $\eta = 1.61$, $C_{\mathrm{I}} = \exp(0)$, $C_{\mathrm{II}} = \exp(0 - 0.21)$, $C_{\mathrm{III}} = \exp(0 - 0.21 + 0.31) \times [1 + (-0.03)]^{13}$ and $C_{\mathrm{IV}} = \exp(0 - 0.21 + 0.31 - 0.38) \times [1 + (-0.03)]^{13} \times (1 + 0.02)^{48}$. $\bar{\pi}_{\mathrm{I}} = 3\%$, $\bar{\pi}_{\mathrm{II}} = 3.5\%$, $\bar{\pi}_{\mathrm{III}} = 20\%$, $\bar{\pi}_{\mathrm{IV}} = 200\%$