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North American Journal of
Economics and Finance 14 (2003) 381–397

THE NORTH AMERICAN JOURNAL OF ECONOMICS AND FINANCE

How homogenizing are monetary unions? Evidence from the U.S. states

Georgios Karras [a,b,*]

[a] *Department of Economics (M/C 144), University of Illinois at Chicago, 601 S. Morgan Street, Chicago, IL 60607, USA*
[b] *Athens University of Economics and Business, Athens, Greece*

Received 19 October 2002; received in revised form 8 July 2003; accepted 15 July 2003

Abstract

How much should we expect monetary integration to reduce differences among participating economies? Using data for the 50 U.S. states over 1929–1999, the paper finds that large differences across states persist in both business-cycle variability and correlations of state income with the U.S. as a whole. It is also shown that about a third of a state's business-cycle volatility is explained by its income correlation with the entire U.S. Since these results hold for the U.S., a monetary union of long history and deep economic integration, the obvious implication is that volatility and correlation differences should be expected to persist to an even greater extent in more recent and less economically integrated monetary unions, such as the Eurozone or a prospective set of dollarized national economies.

JEL classification: E42; F33; F36; F42

Keywords: Monetary union; Optimum currency area; Dollarization

1. Introduction

What happens to the relationships among different economies when they form (or join) an economic or monetary union? One theoretical possibility is that they become *more alike*, primarily because deeper economic integration means a higher volume of intra-industry trade, technological transfers, and common policies (the homogenizing effect). It is also theoretically possible, however, that they become *less alike*, because a higher level of inter-industry trade means greater specialization which leaves the economies more different than before

* Tel.: +1-312-996-2321; fax: +1-312-996-3344.
E-mail address: gkarras@uic.edu (G. Karras).

1062-9408/$ – see front matter
doi:10.1016/S1062-9408(03)00034-2

(the heterogenizing effect). In practice, the homogenizing and heterogenizing effects are almost sure to coexist, so the net result depends on which one dominates.[1]

The question is important because it influences the optimality of monetary unions. Economic theory predicts that the higher the correlation among economies, the better candidates they are for monetary integration.[2] Intuitively, this is because participating in a monetary union rules out the possibility of independent monetary policy, and so it will exacerbate a member economy's business-cycle volatility if the member's output is not sufficiently correlated with that of the union as a whole, reducing the net benefit of membership for that economy.

As Frankel and Rose (1998) point out, it follows that the optimal currency area calculus is effectively endogenized. If the homogenizing effect dominates, monetary unions will be "self-validating" in the sense that participation in the union will render economies more alike, reducing everybody's ex post costs of membership. The policy implication is that the ex ante stabilization costs of joining the monetary union are overestimating the true costs, since participation by itself is going to lead to higher correlations and thus lower ex post costs. If, on the other hand, the heterogenizing effect is stronger, the monetary union will be unstable in the sense that participation will make economies less correlated, thus raising everybody's ex post cost of membership. The policy implication now is that the ex ante stabilization cost of joining the monetary union is underestimating the actual cost, since participation by itself is going to lead to lower correlations and a higher ex post cost.

To address the issue empirically, this paper looks at one of the most successful (at least in terms of longevity) monetary unions of modern times, the U.S., and asks how similar its member economies, the 50 U.S. states, have been in terms of output variability and correlation with the U.S. economy as a whole. The empirical findings show that large and economically important differences in both business-cycle volatility and correlations of state incomes with the U.S. persist across the 50 states, despite a long period of monetary integration and deep economic integration in almost every other respect (significantly centralized fiscal policy, very high factor mobility, no internal trade barriers). Furthermore, the evidence shows that, consistent with the predictions of the paper's theoretical model, there is a statistically significant negative relationship between a state's cyclical variability and its correlation with the U.S. as a whole. The implications are that the homogenizing effect, if it exists, does not dominate the heterogenizing effect in the U.S. monetary union, and that adopting the dollar as a common currency is responsible for at least some of the business-cycle volatility of U.S. states.

The rest of the paper is organized as follows. Section 2 uses a recent model of monetary policy in order to illustrate the theoretical derivation of costs and benefits from membership in a monetary union, and particularly the relationship between a member's output variability and its correlation with the union as a whole. Section 3 describes the data and the empirical

[1] This literature is growing very rapidly. See Frankel and Rose (1997, 1998) and Corsetti and Pesenti (2002) for evidence on the homogenizing effect, but Krugman (1993) and Kalemli-Ozcan, Sorensen, and Yosha (2001) for evidence on the heterogenizing effect. A third possibility, that the economies become neither more nor less alike, is usually ignored in the literature because it requires that the homogenizing and heterogenizing effects cancel out exactly.

[2] This argument, which will be formally derived in Section 2 of the paper, goes back to Mundell (1961), but has been reformulated extensively. See Alesina and Barro (2002) and Alesina, Barro, and Tenreyro (2002) for two recent contributions.

methodology. Section 4 presents the empirical results and demonstrates robustness. Section 5 summarizes and concludes by discussing implications for the U.S. and other existing or prospective monetary unions such as the Eurozone or dollarization in the Americas.

2. Theoretical background

The theoretical framework follows the "New Keynesian" monetary policy model of Clarida, Gali, and Gertler (1999).[3] Let the 50 U.S. states be indexed by i ($i = 1, 2, \ldots, 50$). Each state's loss function takes the form

$$L_i = \frac{1}{2} E_t \left\{ \sum_{j=0}^{\infty} \beta^j [a_i (y_{i,t+j} - k_i)^2 + \pi_{i,t+j}^2] \right\}, \tag{1}$$

where y denotes real income (in deviations from trend), π is inflation, a is the relative weight on output deviations ($a > 0$), β is the discount factor, E denotes mathematical expectation, and k is the output target. As usual, it is assumed that $k \geq 0$ because of distortions, such as imperfect competition or taxes.

For each state, aggregate supply is given by a "New Keynesian" expectations-augmented Phillips curve:

$$\pi_{i,t} = \lambda_i y_{i,t} + E_t \pi_{i,t+1} + u_{i,t}, \tag{2}$$

where $\lambda_i > 0$, $u_{i,t} = \phi_i u_{i,t-1} + z_{i,t}$, $0 < \phi_i < 1$, and $z_{i,t} \sim iid(0, \tau_i^2)$. Eq. (2) can also be written in aggregate-supply form as

$$y_{i,t} = \vartheta_i (\pi_{i,t} - E_t \pi_{i,t+1}) + v_{i,t}, \tag{3}$$

where $\vartheta_i = 1/\lambda_i$ and $v_{i,t} = -u_{i,t}/\lambda_i$.[4] Note that this implies $v_{i,t} = \phi_i v_{i,t-1} - z_{i,t}/\lambda_i$, and define $\sigma_i^2 \equiv \text{Var}(v_{i,t}) = \tau_i^2 [\lambda_i^2 (1 - \phi_i^2)]^{-1}$.

2.1. Independent monetary policy

Without a monetary union, when each state's monetary authority can pursue an independent monetary policy, minimizing (1) subject to (2) leads to the following equilibrium outcome (also known as "discretion"):[5]

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[3] Very similar results can be derived from the "older" monetary policy model based on the work by Kydland and Prescott (1977), Barro and Gordon (1983), and Rogoff (1985), and used to evaluate the effects of monetary integration by Alesina and Grilli (1992, 1994), De Grauwe (1994), and Alesina and Wacziarg (1999). The main differences between these models and the present formulation (see below) are a more modern aggregate supply specification and a richer dynamic structure.

[4] This specification of the Phillips curve goes back to Calvo (1983). For more recent examples, see Rotemberg (1987), Roberts (1995), Gali and Gertler (1999), and Woodford (2003).

[5] It is assumed here that the choice variable for the monetary authority is the inflation rate itself. This assumption is made only for convenience and is very common in the literature (see Alesina and Grilli, 1992; Cukierman, Kiguel, and Liviatan, 1994), because it delivers the model's basic properties without unnecessarily complicating the algebra. Adding a money demand and/or an IS curve would allow the central bank to target the interest rate or the money supply (as in Clarida et al., 1999). While more realistic, this would complicate the model without contributing to the issues examined here.

$$\pi_{i,t}^{\mathrm{IND}} = a_i q_i u_{i,t} + \frac{a_i}{\lambda_i} k_i = -a_i q_i \lambda_i v_{i,t} + \frac{a_i}{\lambda_i} k_i, \tag{4}$$

and

$$y_{i,t}^{\mathrm{IND}} = -\lambda_i q_i u_{i,t} = \lambda_i^2 q_i v_{i,t} \tag{5}$$

where an "IND" superscript denotes outcomes under independent monetary policy, and $q_i = [\lambda_i^2 + a_i(1-\beta\phi_i)]^{-1}$. The macroeconomic performance of each state will be characterized by average ("trend") inflation equal to

$$\bar{\pi}_i^{\mathrm{IND}} = \frac{a_i k_i}{\lambda_i}, \tag{6}$$

and income ("business-cycle") volatility of

$$\mathrm{Var}(y_i^{\mathrm{IND}}) = \lambda_i^4 [\lambda_i^2 + a_i(1-\beta\phi_i)]^{-2} \sigma_i^2. \tag{7}$$

As expected, the inflation bias is increasing with (i) the weight on output (a), (ii) the output target (k), and (iii) the slope of aggregate supply ($\vartheta = 1/\lambda$). It is also apparent that there is a trade-off between average inflation and income variability:[6] if the value of a is very low (so that the central bank is very "conservative" in the sense of assigning a much higher relative weight to inflation than to income), average inflation will be very low, but income will be very unstable.[7]

2.2. A monetary union

Next, consider the formation of a monetary union: the 50 states adopt a common currency, the U.S. dollar, and monetary authority is delegated to a common central bank, the Federal Reserve System. Generalizing Eqs. (1) and (2), and using an "f" subscript to indicate U.S.-wide ("federal") values for variables and parameters, we can write the Fed's loss function as

$$L_f = \frac{1}{2} E_t \left\{ \sum_{j=0}^{\infty} \beta^j [a_f (y_{f,t+j} - k_f)^2 + \pi_{f,t+j}^2] \right\},$$

and the U.S.-wide aggregate supply as $\pi_{f,t} = \lambda_f y_{f,t} + E_t \pi_{f,t+1} + u_{f,t}$.

Then, for state i at equilibrium, as suggested by (4),

$$\pi_{i,t}^{\mathrm{UNION}} = \pi_{f,t} = a_f q_f u_{f,t} + \frac{a_f}{\lambda_f} k_f = -a_f q_f \lambda_f v_{f,t} + \frac{a_f}{\lambda_f} k_f, \quad \forall i,$$

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[6] As pointed out first by Taylor (1979), there is also a trade-off between output variability and inflation variability. The latter is given here by

$$\mathrm{Var}(\pi_i^{\mathrm{IND}}) = a_i^2 \lambda_i^2 [\lambda_i^2 + a_i(1-\beta\phi_i)]^{-2} \sigma_i^2;$$

that is a low a reduces the volatility of inflation but raises that of income. See Fuhrer (1997).

[7] Rogoff (1985) examines the optimal value for a. Fischer and Summers (1989) show that a similar trade-off exists if the source of uncertainty is the central bank's inability to determine the inflation rate without error.

and substituting into (3)

$$y_{i,t}^{\text{UNION}} = -a_f q_f (1 - \phi_f) v_{f,t} + v_{i,t}, \tag{8}$$

where the "UNION" superscript refers to outcomes under the monetary union.[8] Note that, under the Fed, state *i*'s income is affected not just by its own income shock, $v_{i,t}$, but also by the U.S.-wide shock, $v_{f,t}$. The reason, of course, is that the U.S.-wide shock is (negatively) "exported" to all the states via the conduct of monetary policy by the Fed.

How does each state's macroeconomy perform under a monetary union? Average inflation will be given by

$$\bar{\pi}_i^{\text{UNION}} = \frac{a_f k_f}{\lambda_f}, \tag{9}$$

and business-cycle volatility by

$$\text{Var}(y_i^{\text{UNION}}) = a_f^2 q_f^2 (1 - \phi_f)^2 \sigma_f^2 + \sigma_i^2 - 2a_f q_f (1 - \phi_f) \rho_{i,f} \sigma_i \sigma_f, \tag{10}$$

where $\rho_{i,f} \equiv \text{corr}(v_{i,t}, v_{f,t})$.

Comparing Eqs. (6) and (9), it is apparent that the monetary union will reduce a state's average inflation rate only if the Fed is a more "conservative" monetary authority ($a_f = a_i$ and $k_f < k_i$) and faces a less tempting aggregate supply ($\vartheta_f = \vartheta_i$). Under these conditions, $\bar{\pi}_i^{\text{UNION}} < \bar{\pi}_i^{\text{IND}}$.

More relevant to the point of the present paper, however, note that Eq. (10) implies that each state's income variability under the monetary union will be negatively related to the correlation between that state's income and U.S.-wide income. In fact, comparing Eqs. (7) and (10) shows that joining the monetary union may very well increase income volatility: this can be thought of as the macroeconomic *cost* of union membership (while the inflation effect discussed in the previous paragraph is the main *benefit*). From (10), this stabilization cost will be smaller, the closer $\rho_{i,j}$ is to unity. Intuitively, if the business cycle of, say, Illinois is very highly correlated with U.S.-wide cyclical income, countercyclical monetary policy conducted by the Fed will be a very close substitute for monetary policy conducted by a hypothetical Illinois Central Bank. In this case, Illinois membership in the monetary union, even though it means giving up independent monetary policy, will not be very costly. If, on the other hand, Illinois income is negatively correlated with the U.S., so that downturns in Illinois tend to coincide with expansions in the U.S. and vice versa, surrendering monetary policy to the Fed will *de*stabilize Illinois by amplifying its business cycle.

3. Data and empirical methodology

All state data are from the Bureau of Economic Analysis. The series used is personal income per capita, which is available annually from 1929 to 1999 for each of the 48 con-

[8] Once more, the equality of inflation rates under the Fed follows from the assumption that the central bank uses the inflation rate itself as the instrument of monetary policy. If the instrument is the interest rate or the money growth rate, then these variables would be common across the states and inflation rates may differ according to money demand and IS curves. This would not change any of the conclusions drawn below, however, so the simpler specification is preferred here, as in much of the literature.

tiguous U.S. states and the District of Columbia (*Data Set I*), and annually from 1950 to 1999 for each of the 50 U.S. states and the District of Columbia (*Data Set II*).[9] The data are deflated by the Consumer Price Index to construct a series for real personal income, Y.

Three different methods are used to detrend the output series of each state and estimate its cyclical component. The first is simple differencing, calculating the growth rate of real income as $(Y_{i,t} - Y_{i,t-1})/Y_{i,t-1}$.

The second method is the Hodrick–Prescott (HP) filter, proposed by Hodrick and Prescott (1980) and extensively used in the business-cycle literature. Letting $y_{i,t} = \ln(Y_{i,t})$, the HP filter defines the trend component, $\bar{y}_{i,t}$, as the one that minimizes $\sum_{t=1}^{T}(y_{i,t} - \bar{y}_{i,t})^2 + \ell\sum_{t=2}^{T-1}[(\bar{y}_{i,t+1} - \bar{y}_{i,t}) - (\bar{y}_{i,t} - \bar{y}_{i,t-1})]$ for $\ell > 0$. The cyclical component is simply $y_{i,t} - \bar{y}_{i,t}$. Here, we select $\ell = 100$, the value recommended by Kydland and Prescott (1989) for annual data.

The third method makes use of the recently very popular band-pass (BP) filter proposed by Baxter and King (1995) and evaluated by Stock and Watson (1998) and Christiano and Fitzgerald (1999), who also compare its properties to those of the HP filter. The main objective of the BP filter as implemented by Baxter and King (1995) is to remove both the high-frequency and low-frequency component of a series, leaving the business-cycle frequencies. Here, we remove all fluctuations shorter than two or longer than 8 years.[10]

While differences among the results obtained by the three filters are not difficult to detect (for example, differencing generally produces the most volatile series, while the BP filter yields the smoothest), the main characteristics are remarkably similar. This robustness will be formally confirmed by the findings of the next section.

4. Empirical results

4.1. State income variability

We start by comparing the volatility of cyclical income across the states. Table 1 reports the standard deviation of detrended real personal income for each state and for the entire U.S. over the 1950–1999 period (*Data Set II*) for the three methods outlined in the last section (differencing, the HP filter, and the BP filter). While, not surprisingly, the estimates vary by the detrending method chosen, the relative positions of the states in volatility rankings are very similar across the three methods.[11] For example, New Mexico has the smoothest

9 Gross State Product (GSP) data would arguably be a better alternative. Unfortunately, the GSP data are available for too short a period of time to be of use in this study.

10 The low-pass (LP) filter $\alpha(L)$, which forms the basis for the band-pass filter, selects a finite number of moving average weights, α_h, to minimize $Q = \int_{-\pi}^{\pi}|\delta(\omega)|^2\,d\omega$, where $\alpha(L) = \sum_{h=-K}^{K}\alpha_h L^h$ and $\alpha_K(\omega) = \sum_{h=-K}^{K}\alpha_h\,e^{-i\omega h}$. The LP filter uses $\alpha_K(\omega)$ to approximate the infinite MA filter $\beta(\omega)$. Define $\delta(\omega) \equiv \beta(\omega) - \alpha(\omega)$. Minimizing Q minimizes the discrepancy between the ideal LP filter $\beta(\omega)$ and its finite representation, $\alpha_K(\omega)$, at frequency ω. We set ω_L and ω_H, the lower and upper frequencies of two low-pass filters, at 8 and 2, respectively. The frequency representation of the band-pass weights becomes $\alpha_K(\omega_H) - \alpha_K(\omega_L)$, and forms the basis of the Baxter–King filter, which provides an alternative estimate of the trend component $\bar{y}_{i,t}$, and the cyclical component $y_{i,t} - \bar{y}_{i,t}$.

11 Spearman's rank correlation coefficients are 0.74 for the (DIFF, HP) pair, 0.95 for (DIFF, BP), and 0.70 for (HP, BP).

Table 1
Variability of cyclical income by state: 1950–1999

i	σ_i^{DIFF}	σ_i^{HP}	σ_i^{BP}	Rank
New England				
CT	2.609	2.693	1.614	26
ME	2.796	2.824	1.608	34
MA	2.275	2.738	1.368	14
NH	2.694	3.143	1.648	30
RI	2.343	2.635	1.392	16
VT	2.618	2.681	1.587	29
Mideast				
DE	2.922	2.790	1.794	37
DC	2.615	2.567	1.655	27
MD	2.236	2.569	1.388	13
NJ	2.145	2.353	1.341	8
NY	1.931	2.234	1.161	2
PA	2.048	2.032	1.313	6
Great Lakes				
IL	2.148	2.073	1.384	9
IN	3.205	2.812	2.032	44
MI	3.715	3.708	2.402	45
OH	2.598	2.497	1.668	25
WI	2.360	2.037	1.530	18
Plains				
IA	4.026	2.977	2.440	47
KS	3.021	2.337	1.824	38
MN	2.423	2.304	1.530	20
MO	2.214	2.155	1.469	12
NE	4.075	2.913	2.428	48
ND	9.923	7.136	5.849	51
SD	6.690	4.693	4.168	50
Southeast				
AL	2.548	2.347	1.538	24
AR	3.032	2.474	1.843	39
FL	2.455	2.753	1.484	21
GA	2.616	2.774	1.639	28
KY	2.517	2.191	1.502	23
LA	2.124	2.009	1.283	7
MS	3.171	2.284	1.907	43
NC	2.515	2.543	1.557	22
SC	3.065	3.048	1.765	40
TN	2.371	2.323	1.490	19
VA	2.172	2.334	1.303	10
WV	2.745	2.370	1.785	33
Southwest				
AZ	2.875	3.089	1.763	36
NM	1.874	1.626	1.183	1
OK	2.713	1.873	1.385	31
TX	2.023	1.837	1.192	4

Table 1 (*Continued*)

i	σ_i^{DIFF}	σ_i^{HP}	σ_i^{BP}	Rank
Rocky Mountain				
CO	2.342	2.005	1.577	15
ID	3.155	2.536	1.914	42
MT	3.924	2.619	2.225	46
UT	2.186	1.963	1.506	11
WY	2.872	2.597	1.736	35
Far West				
AK	5.775	5.030	3.611	49
CA	2.036	2.145	1.251	5
HI	3.091	3.206	1.604	41
NV	2.728	2.722	1.750	32
OR	2.354	2.263	1.507	17
WA	2.017	2.281	1.288	3
USA	1.887	1.957	1.206	

Notes: σ_i is the standard deviation ($\times$100) of state i's detrended income series. DIFF refers to differencing; HP to the Hodrick–Prescott filter, using $\ell = 100$; BP to the band-pass filter, implemented as in Baxter and King (1995) using $K = 2$ lags. "Rank" ranks the states in ascending variability based on the differenced series.

cyclical component when differencing and the HP filter are used (standard deviations of 0.019 and 0.016, respectively), while New York has the smoothest cycle with the BP filter (standard deviation of 0.016). At the other extreme, North Dakota has the most volatile cyclical component (with a standard deviation of 0.099 for the differenced series, 0.071 with the HP filter, and 0.058 under the BP filter), regardless of the detrending method chosen.[12]

But the most striking feature of Table 1 is the markedly wide range of these measures of income variability across the states. Thus, the standard deviation of cyclical income in the most volatile state (ND) exceeds that of the least volatile (NM or NY) by a factor of 4 (for the HP filter) or 5 (for differencing and the BP method). There are states, such as North Dakota, South Dakota, and Alaska, for which standard deviations are close to the maximum values; other states, like Michigan, Arizona, and South Carolina, display medium-sized volatility; and there are also states, such as New York, Pennsylvania, and New Mexico, for which standard deviations are quite low.

It is obvious that large, economically meaningful differences in business-cycle performance across the 50 U.S. states have survived monetary integration. This finding is robust to the detrending approach used and the choice of data set.[13]

[12] It is interesting that the U.S. as a whole has a less volatile cyclical component than the overwhelming majority of the states. Indeed, the only exceptions are the states of New York and New Mexico.

[13] These estimates of cyclical volatility were also estimated using the longer series of *Data Set I*, but they lead to very similar conclusions and so are not reported in order to preserve space. The only substantive difference is that the *Data Set I* standard deviations are consistently higher for all states and the entire U.S. economy, but this was to be expected as the longer series include the Great Depression, its aftermath, and World War II. Despite that, the estimated volatility measures are highly correlated between the two data sets (correlation coefficients are 0.81 for the differenced series, 0.59 for the HP filtered, and 0.81 for the BP filtered).

4.2. State correlation with the U.S.

Table 2 reports the correlation coefficients of each state's cyclical component of real personal income with that of the entire U.S. over the 1950–1999 period, using the three detrending methods outlined in Section 3. Though, again as expected, the estimated coefficients vary with the detrending method used, once more the implied rankings are very similar across the three approaches.[14] Georgia is the most highly correlated state with the U.S. when differencing and the HP filter are used (correlation coefficients of 0.912 and 0.939, respectively), while Illinois has the highest correlation with the BP filter (coefficient of 0.932). At the other end, the least correlated state with the entire U.S. is Alaska under differencing and the BP filter (correlation coefficients are 0.084 and 0.035, respectively) and Wyoming when the HP filter is used (coefficient of −0.046).

For the purpose of the present study, however, the most telling aspect of Table 2 is that states differ significantly in their correlations with the entire U.S. There are states, such as Illinois, Ohio, and Georgia, for which all coefficients exceed 0.9, which are extremely strongly correlated with the U.S. as a whole; other states, like Louisiana, Colorado, and Idaho, with coefficients around 0.5, which are moderately correlated with the entire economy; and there are also states, such as North Dakota, South Dakota, and Alaska, for which all coefficients are below 0.3, and which are very poorly correlated with the entire U.S.

It would be also interesting to compare these estimated coefficients with the correlations within other groups of countries. Starting with such cyclical correlations for some of the current and prospective members of the Eurozone, Karras (2002a) has estimated that, over the 1960–1999 period, such correlations have ranged from 0.02 (Turkey–EU) to 0.93 (France–EU) when differencing is used, and from −0.24 (Turkey–EU) to 0.85 (France–EU) for the BP-filtered series.

Karras (2002b) has also estimated cyclical correlations for several North, Central, and South American countries with the U.S. over 1968–1997. These have ranged from −0.25 (Panama–U.S.) to 0.74 (Canada–U.S.) when differencing is used, and from −0.24 (Panama–U.S.) to 0.80 (Canada–U.S.) for the BP-detrended series. An important difference, of course, that must be kept in mind is that, unlike the U.S. states over 1950–1999 (or 1929–1999), neither the European economies in 1960–1999 nor the American countries (except Panama) in 1968–1997 had formed a monetary union.

Overall, it is apparent that, similar to the case of cyclical variability, large and economically important differences in correlations of state incomes with the U.S. persist, despite a long period of monetary integration. This result is also robust to the detrending method employed and the data set used.[15]

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[14] Spearman's rank correlation coefficients are 0.93 for the (DIFF, HP) pair, 0.99 for (DIFF, BP), and 0.92 for (HP, BP).

[15] Cyclical correlations were also estimated using the longer series of *Data Set I*, but again they result in conclusions so similar to those of *Data Set II* that they are not reported because of space considerations. Virtually the only difference is that the *Data Set I* correlations are higher for all states, but this should not be interpreted as evidence that the correlations have decreased over time. Instead, it follows from the fact that the Great Depression, its aftermath, and World War II (included in *Data Set I* but not *Data Set II*) were highly synchronized across the states. Once more, the estimated correlation measures are highly correlated themselves between the two data sets (with correlation coefficients of 0.82 for the differenced series, 0.65 for the HP filtered, and 0.78 for the BP filtered).

Table 2
Cyclical correlations of the states with the U.S.: 1950–1999

i	$\rho_{i,\mathrm{US}}^{\mathrm{DIFF}}$	$\rho_{i,\mathrm{US}}^{\mathrm{HP}}$	$\rho_{i,\mathrm{US}}^{\mathrm{BP}}$	Rank
New England				
CT	0.788	0.872	0.799	23
ME	0.751	0.865	0.749	27
MA	0.826	0.865	0.868	16
NH	0.792	0.846	0.818	22
RI	0.801	0.863	0.790	21
VT	0.850	0.901	0.851	14
Mideast				
DE	0.692	0.846	0.753	32
DC	0.343	0.453	0.522	46
MD	0.881	0.914	0.868	7
NJ	0.881	0.928	0.896	6
NY	0.815	0.873	0.847	19
PA	0.902	0.918	0.901	4
Great Lakes				
IL	0.909	0.920	0.932	3
IN	0.867	0.847	0.901	10
MI	0.845	0.844	0.868	15
OH	0.910	0.933	0.923	2
WI	0.878	0.870	0.891	8
Plains				
IA	0.364	0.523	0.432	45
KS	0.382	0.523	0.431	44
MN	0.734	0.873	0.748	28
MO	0.852	0.913	0.869	13
NE	0.278	0.449	0.334	48
ND	0.238	0.132	0.281	49
SD	0.216	0.237	0.258	50
Southeast				
AL	0.859	0.848	0.867	12
AR	0.773	0.788	0.784	25
FL	0.764	0.754	0.770	26
GA	0.912	0.939	0.915	1
KY	0.819	0.853	0.841	18
LA	0.519	0.220	0.566	40
MS	0.648	0.699	0.673	35
NC	0.821	0.884	0.836	17
SC	0.806	0.852	0.789	20
TN	0.878	0.887	0.885	9
VA	0.866	0.932	0.871	11
WV	0.675	0.589	0.643	34
Southwest				
AZ	0.715	0.705	0.705	30
NM	0.637	0.533	0.686	37
OK	0.537	0.296	0.517	39
TX	0.683	0.617	0.734	33

Table 2 (*Continued*)

i	$\rho_{i,US}^{DIFF}$	$\rho_{i,US}^{HP}$	$\rho_{i,US}^{BP}$	Rank
Rocky Mountain				
CO	0.562	0.420	0.579	38
ID	0.455	0.504	0.494	41
MT	0.436	0.348	0.472	42
UT	0.727	0.609	0.754	29
WY	0.322	−0.069	0.392	47
Far West				
AK	0.084	−0.046	0.035	51
CA	0.896	0.908	0.903	5
HI	0.398	0.398	0.281	43
NV	0.646	0.508	0.661	36
OR	0.786	0.746	0.826	24
WA	0.710	0.674	0.694	31
USA	1.000	1.000	1.000	

Notes: $\rho_{i,US}$ is the correlation of state i's cyclical component with the U.S. cyclical component. DIFF refers to differencing; HP to the Hodrick–Prescott filter, using $\ell = 100$; BP to the band-pass filter, implemented as in Baxter and King (1995) using $K = 2$ lags. "Rank" ranks the states in descending correlation with the U.S. based on the differenced series.

4.3. Volatilities and correlations over time

Another interesting question is how the estimated volatilities and correlations have evolved over time. In addition to providing some predictive information, such an exercise may shed light on the endogeneity issue discussed earlier. While a comparison of the pre-WW II and post-WW II numbers would be very difficult to interpret and thus is of limited value (see footnotes 12 and 13), this section considers an alternative by splitting the 1950–1999 period into two approximately equal subperiods: 1950–1975 and 1976–1999.[16]

Tables 3 and 4 present the results for the two subperiods. Tables 3 and 4 report the standard deviations of only the differenced and BP-filtered series (as Table 1 shows, the differenced and HP-filtered series give very similar estimates) and correlations of the differenced series only. The rest of the results, not reported to preserve space, lead to the same conclusions discussed below (all results are available on request).

Comparing the standard deviations first, the evidence indicates that income volatility has generally decreased, and particularly in the high-variance states. While there are exceptions (such as Texas, where the standard deviation increased from 1.844 to 2.196, according to the differenced series), most of the state economies were less volatile in 1976–1999 than in 1950–1975. Note that the volatility of the entire U.S. economy declined in the most recent period, as well. Interestingly, however, for many of the high-variance states (such as the Dakotas and Alaska), the decline in variability has been more pronounced than that of the entire economy, so that their *relative* volatility has fallen as well.

Moving on to the estimated correlations, most of the numbers of Table 4 are higher than those of Table 3. Once more there are exceptions (such as Texas, again, and Oklahoma,

[16] I am grateful to an anonymous referee for suggesting this.

Table 3
Variability and correlations: 1950–1975

i	σ_i^{DIFF}	σ_i^{BP}	$\rho_{i,\text{US}}^{\text{DIFF}}$
New England			
CT	2.931	1.733	0.814
ME	2.993	1.658	0.676
MA	2.184	1.309	0.896
NH	2.552	1.494	0.824
RI	2.465	1.345	0.819
VT	2.815	1.641	0.826
Mideast			
DE	3.528	2.181	0.696
DC	2.681	1.652	0.137
MD	2.380	1.534	0.875
NJ	2.221	1.341	0.912
NY	1.887	1.149	0.810
PA	2.253	1.420	0.884
Great Lakes			
IL	2.144	1.406	0.885
IN	3.722	2.300	0.850
MI	4.262	2.585	0.864
OH	2.944	1.807	0.903
WI	2.639	1.700	0.867
Plains			
IA	4.587	2.802	0.127
KS	3.727	2.303	0.201
MN	2.326	1.465	0.603
MO	2.280	1.455	0.801
NE	4.878	2.894	0.064
ND	11.807	6.820	0.197
SD	8.300	5.159	0.101
Southeast			
AL	2.776	1.733	0.832
AR	3.199	1.920	0.693
FL	2.916	1.729	0.773
GA	2.860	1.768	0.897
KY	2.554	1.450	0.785
LA	2.978	1.247	0.622
MS	3.675	2.266	0.567
NC	2.590	1.596	0.752
SC	3.724	2.118	0.793
TN	2.420	1.507	0.853
VA	2.289	1.394	0.838
WV	3.244	2.204	0.632
Southwest			
AZ	3.427	1.964	0.670
NM	1.998	1.233	0.529
OK	2.995	1.404	0.614
TX	1.844	0.973	0.746

Table 3 (*Continued*)

i	σ_i^{DIFF}	σ_i^{BP}	$\rho_{i,\text{US}}^{\text{DIFF}}$
Rocky Mountain			
CO	2.875	1.969	0.496
ID	3.778	2.352	0.351
MT	4.813	2.723	0.428
UT	2.312	1.570	0.664
WY	2.636	1.555	0.347
Far West			
AK	6.898	3.874	−0.045
CA	2.007	1.283	0.890
HI	3.483	2.025	0.407
NV	3.156	1.943	0.574
OR	2.459	1.541	0.766
WA	2.135	1.383	0.664
USA	1.932	1.212	1.000

Notes: See *Notes* to Table 1.

where the correlation coefficients declined), but the overwhelming majority of states were more highly correlated with the entire U.S. over the 1976–1999 period than over 1950–1975.

In terms of their monetary union, the states' lower variance and their higher correlation with the U.S., point in the same direction: a common currency was more appropriate (i.e., less costly in stabilization terms) during 1976–1999 than during 1950–1975. The extent to which this result is the direct consequence of the monetary union (as predicted by the endogeneity argument) is one of the most promising areas of future research.

4.4. Are volatilities and correlations related?

The last three subsections have established that there are sizable differences in income volatility across the states, and in their correlations with the entire U.S. economy. An interesting and important question is whether these measures of variability and correlation are inversely related, as suggested by the theoretical model of Section 2. Answering this question is important not only for deciding whether the theory is supported or refuted, but also because it can help us determine the extent to which a common currency has contributed to the variability of cyclical income in the U.S. states.

Table 5 examines the relationship between the states' income volatility and their correlation with the entire U.S. for both data sets and each of the three detrending methods. Focusing on *Data Set I* first, the top panel of Table 5 shows that the relationship is found to be negative for all three detrending approaches, and highly statistically significant for two of them (differencing and the BP filter). The bottom panel repeats the same exercise for *Data Set II*, finding very similar results. The relationship is negative in all three specifications, and again highly statistically significant when differencing and the BP filter are used. These results are consistent with the predictions of the theoretical model: the more (less) correlated a state's income is with that of the U.S. as a whole, the lower (higher) its cyclical variability.

Table 4
Variability and correlations: 1976–1999

i	σ_i^{DIFF}	σ_i^{BP}	$\rho_{i,US}^{DIFF}$
New England			
CT	2.273	1.388	0.821
ME	2.623	1.262	0.821
MA	2.415	1.385	0.841
NH	2.895	1.619	0.778
RI	2.254	1.347	0.799
VT	2.415	1.388	0.860
Mideast			
DE	2.160	1.213	0.744
DC	2.596	1.646	0.601
MD	2.047	1.140	0.876
NJ	2.100	1.252	0.861
NY	2.018	1.167	0.877
PA	1.785	1.069	0.923
Great Lakes			
IL	2.158	1.312	0.936
IN	2.559	1.533	0.912
MI	3.079	1.963	0.823
OH	2.178	1.357	0.930
WI	2.000	1.228	0.886
Plains			
IA	3.271	2.039	0.750
KS	1.854	1.119	0.757
MN	2.535	1.564	0.891
MO	2.142	1.368	0.900
NE	2.973	1.889	0.734
ND	7.379	4.522	0.381
SD	4.358	2.638	0.644
Southeast			
AL	2.043	1.146	0.913
AR	2.666	1.725	0.852
FL	1.742	1.064	0.774
GA	2.304	1.380	0.930
KY	2.344	1.399	0.834
LA	2.126	1.188	0.263
MS	2.299	1.232	0.771
NC	2.404	1.427	0.887
SC	2.041	1.111	0.873
TN	2.260	1.367	0.887
VA	1.951	1.110	0.885
WV	1.852	1.081	0.733
Southwest			
AZ	2.156	1.281	0.794
NM	1.629	1.059	0.715
OK	2.222	1.357	0.355
TX	2.196	1.328	0.571

Table 4 (*Continued*)

i	σ_i^{DIFF}	σ_i^{BP}	$\rho_{i,US}^{DIFF}$
Rocky Mountain			
CO	1.623	0.938	0.704
ID	2.222	1.297	0.579
MT	2.593	1.592	0.449
UT	2.063	1.226	0.775
WY	3.076	1.944	0.236
Far West			
AK	3.703	2.208	0.056
CA	2.040	1.102	0.884
HI	1.915	0.979	0.295
NV	2.225	1.345	0.758
OR	2.271	1.315	0.780
WA	1.909	1.136	0.740
USA	1.810	1.094	1.000

Notes: See *Notes* to Table 1.

Table 5
$\sigma_i = b_0 + b_1\rho_{i,US} + e_i$

	DIFF	HP	BP
	Data Set I: 1929–1999		
b_0	0.156** (0.036)	0.103* (0.047)	0.094** (0.019)
b_1	−0.105** (0.040)	−0.042 (0.051)	−0.061** (0.021)
R^2	0.368	0.067	0.351
N	49	49	49
	Data Set II: 1950–1999		
b_0	0.056** (0.009)	0.036** (0.006)	0.034** (0.006)
b_1	−0.039** (0.011)	−0.014 (0.007)	−0.023** (0.007)
R^2	0.418	0.176	0.394
N	51	51	51

Notes: DIFF refers to differencing; HP to the Hodrick–Prescott filter; BP to the band-pass filter. See the text for details. N is the number of observations. Heteroskedasticity-consistent (White, 1980) estimated standard errors in parentheses. Statistical significance indicated by asterisks (** at the 1% level and * at the 5% level).

Moreover, the estimated R^2s are substantial. On the basis of the differenced or BP-filtered series, they imply that more than one third of a state's cyclical variability is explained by how strongly (or weakly) its income is correlated with that of the entire U.S. In addition to being consistent with the predictions of the model, this gives us an estimate of the business-cycle cost of monetary union.

5. Conclusions

Economic theory suggests that the stabilization cost of joining a monetary union depends on the cyclical correlation between the member country and the union as a whole. It has been argued persuasively, however, that this correlation itself may also endogenously depend on membership because joining the union could make an economy more similar to other members (raising the correlation) or less similar to them (reducing the correlation) depending on the nature of trade and specialization that takes place.

To shed some light on this question, this paper examined the monetary union of the 50 U.S. states, one of the most successful monetary unions of modern times. Using data for the 1929–1999 period for the 48 contiguous states and the District of Columbia, and for the 1949–1999 period for the 50 states and the District of Columbia, the paper investigates the extent to which the economies of the U.S. states are similar in terms of business-cycle volatility and output correlation to the U.S. as a whole.

The results show that large and economically important differences across states persist in both business-cycle variability and correlations of state income with the entire U.S., despite a long period of monetary integration. These results are robust to the three detrending methods employed and the two data sets used.

It is worth emphasizing that these results are obtained for a monetary union which has been characterized by virtually no internal impediments to trade, high degree of financial integration, substantially integrated fiscal policy, and very high labor and capital mobility. The obvious implication is that volatility and correlation differences should be expected to persist to an even greater degree in other, less developed, monetary unions, such as the Eurozone or a prospective set of dollarized national economies, where the homogenizing effect is expected to be even weaker.

How big is the stabilization cost of these differences? Recall from Section 2 that economic theory predicts a negative relationship between cyclical variability and correlation with the union as a whole. This prediction is statistically significantly confirmed for our data sets: states that are highly correlated with the U.S. have smoother cyclical incomes, whereas states that are poorly correlated with the U.S. are characterized by significantly wider cyclical income fluctuations. Additionally, the effect appears to be economically large: at least a third of a state's business-cycle volatility is explained by how strongly (or weakly) its income is correlated with that of the entire U.S.

Acknowledgements

I am grateful to Pierre Siklos and three anonymous referees for helpful comments and suggestions. Remaining errors and omissions are mine.

References

Alesina, A., & Barro, R. J. (2002). Currency unions. *Quarterly Journal of Economics*, *117*, 409–436.

Alesina, A., & Grilli, V. (1992). The European Central Bank: Reshaping monetary politics in Europe. In M. B. Canzoneri, V. Grilli, & P. R. Masson (Eds.), *Establishing a central bank: Issues in Europe and lessons from the US*. Cambridge: Cambridge University Press.

Alesina, A., & Grilli, V. (1994). On the feasibility of a one-speed or multispeed European monetary union. In B. Eichengreen & J. Frieden (Eds.), *The political economy of European monetary unification*. Boulder: Westview Press.

Alesina, A., & Wacziarg, R. (1999, January). *Is Europe going too far?* (NBER Working Paper No. 6883).

Alesina, A., Barro, R. J., & Tenreyro, S. (2002, July). *Optimal currency areas* (NBER Working Paper No. 9072).

Barro, R., & Gordon, D. (1983). Rules, discretion, and reputation in a model of monetary policy. *Journal of Monetary Economics, 12,* 101–122.

Baxter, M., & King, R. G. (1995, February). *Measuring business cycles: Approximate band-pass filters for economic time series* (NBER Working Paper No. 5022).

Calvo, G. (1983). Staggered prices in a utility maximizing framework. *Journal of Monetary Economics*, *12*, 383–398.

Christiano, L. J., & Fitzgerald, T. J. (1999, July). *The band pass filter* (NBER Working Paper No. 7257).

Clarida, R., Gali, J., & Gertler, M. (1999). The science of monetary policy: A new Keynesian perspective. *Journal of Economic Literature, 37,* 1661–1707.

Corsetti, G., & Pesenti, P. (2002, February). *Self-validating optimum currency areas* (NBER Working Paper No. 8783).

Cukierman, A., Kiguel, M. A., & Liviatan, N. (1994). How much to commit to an exchange-rate rule: Balancing credibility and flexibility. In P. L. Siklos (Ed.), *Varieties of monetary reforms: Lessons and experiences on the road to monetary union*. Dordrecht: Kluwer Academic Publishers.

De Grauwe, P. (1994). Towards European monetary union without the EMS. *Economic Policy*, *18*, 149–185.

Fischer, S., & Summers, L. H. (1989, May). Should governments learn to live with inflation? *American Economic Review*, *79*, 382–387.

Frankel, J. A., & Rose, A. K. (1997). Is EMU more justifiable ex post than ex ante? *European Economic Review*, *41*, 753–760.

Frankel, J. A., & Rose, A. K. (1998, July). The endogeneity of the optimum currency area criteria. *Economic Journal*, *108*, 1009–1025.

Fuhrer, J. C. (1997). Inflation/output variance trade-offs and optimal monetary policy. *Journal of Money, Credit, and Banking*, *29*, 214–234.

Gali, J., & Gertler, M. (1999). Inflation dynamics: A structural econometric analysis. *Journal of Monetary Economics*, *44*, 195–222.

Hodrick, R. J., & Prescott, E. C. (1980). *Postwar US business cycles: An empirical investigation* (Discussion Paper 451). Carnegie Mellon University.

Kalemli-Ozcan, S., Sorensen, B. E., & Yosha, O. (2001). Economic integration, industrial specialization, and the asymmetry of economic fluctuations. *Journal of International Economics*, *55*, 107–137.

Karras, G. (2002a). Costs and benefits of Euro membership: Evidence from nineteen European countries. *National Business and Economics Society, Papers and Proceedings* (on CD-ROM).

Karras, G. (2002b, September). Costs and benefits of dollarization: Evidence from North, Central, and South America. *Journal of Economic Integration*, *17*(3), 502–516.

Krugman, P. (1993). Lessons of Massachusetts for EMU. In F. Giavazzi & F. Torres (Eds.), *The transition to economic and monetary union in Europe* (pp. 241–261). Cambridge: Cambridge University Press.

Kydland, F. E., & Prescott, E. C. (1977). Rules rather than discretion: The inconsistency of optimal plans. *Journal of Political Economy*, *85*, 473–490.

Kydland, F. E., & Prescott, E. C. (1989, April). *A Fortran subroutine for efficiently computing HP-filtered time series* [Research Memorandum]. Federal Reserve Bank of Minneapolis.

Mundell, R. (1961). A theory of optimum currency areas. *American Economic Review*, *51*, 657–665.

Roberts, J. M. (1995). New Keynesian economics and the Phillips curve. *Journal of Money, Credit, and Banking*, *27*, 975–998.

Rogoff, K. (1985). The optimal degree of commitment to an intermediate monetary target. *Quarterly Journal of Economics*, *100*, 1169–1190.

Rotemberg, J. (1987). The new Keynesian microfoundations. In S. Fischer (Ed.), *NBER macroeconomics annual*. Cambridge, MA: MIT Press.

Stock, J. H., & Watson, M. W. (1998, April). *Business cycle fluctuations in US macroeconomic time series* (NBER Working Paper No. 6528).

Taylor, J. B. (1979). Estimation and control of a macroeconomic model with rational expectations. *Econometrica*, *47*, 1267–1286.

White, H. (1980). A heteroskedasticity-consistent covariace matrix estimator and a direct test for heteroskedasticity. *Econometrica*, *48*, 817–838.

Woodford, M. (2003). *Interest and prices: Foundations of a theory of monetary policy*. Princeton, NJ: Princeton University Press.