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Bilateral Trade and Business Cycle Synchronization: Evidence from Mexico and United States Manufacturing Industries¹

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Abstract

We provide evidence that production-side links between Mexico and U.S. manufacturing sectors became stronger after NAFTA was enacted and, as a consequence, business cycles in these countries became more synchronized. This suggests that the positive effect of bilateral trade on business cycle synchronization found in previous studies for the case of industrial countries may also hold for industrial and less developed country pairs. The recent entry of other unskilled labor-abundant countries into global trade, however, seems to be affecting Mexico's competitiveness in some industries and causing Mexico to be losing market share in the U.S. import market. As a consequence, this event could lead to a permanent negative shift in Mexico's manufacturing output levels, relative to the U.S., and could possibly weaken the degree of business cycle synchronization between these countries. A related effect is shown to be that, in some industries where strong Mexico-U.S. production-sharing links persist, overall North American output is apparently being affected by the global movement of these activities towards the Asian block.

Keywords: Business Cycle Synchronization, Trade Integration, NAFTA.

JEL Classification: E32, F15, F32.

Resumen

En este trabajo se proporciona evidencia que sugiere que la integración productiva entre las manufacturas mexicanas y las estadounidenses se intensificó a partir de la entrada en marcha del Tratado de Libre Comercio de América del Norte. Como una consecuencia de ello, el grado de sincronización de los ciclos económicos de ambos países aparentemente se acrecentó. La entrada reciente de países abundantes en mano de obra poco calificada a los flujos de comercio internacional, no obstante, parece estar afectando la competitividad de las exportaciones mexicanas en algunos sectores y provocando que México pierda participación en el mercado de las importaciones estadounidenses. Esto podría llegar a debilitar el grado de sincronización cíclica entre estos dos países.

Palabras Clave: Sincronización de los Ciclos Económicos, Integración Comercial, TLCAN.

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Introduction

Ever since Mundell's (1961) seminal work on optimum currency areas, the net benefits of forming a monetary union have been argued to be larger to the extent that the countries involved exhibit stronger links through trade, a higher synchronization in their business cycles, a larger cross-country mobility of labor, and the possibility of sharing risks through, for example, fiscal transfers. Frankel and Rose (1998), however, point out that at least two of these criteria –the degrees of trade intensity and of business cycle synchronization- are jointly endogenous to the decision of forming a currency area. In particular, forming a currency union may reduce transaction costs of trade and, as a consequence, may strengthen the commercial links between the member countries. Furthermore, deeper trade integration may itself modify the extent to which business cycles of the trading countries are synchronized. As a consequence of these interactions, the *ex post* benefits and costs of forming a monetary union may be different from those that can be computed from data on trade and business cycle correlations observed before the currency area is formed. In effect, if an increase in bilateral trade leads to a higher correlation in the trading countries' business cycles, the costs of giving up autonomous monetary policy to face idiosyncratic shocks may become smaller once the currency union starts operating. In this case, the two criteria interact to strengthen the *ex post* optimality of a currency union. In contrast, if trade integration reduces the countries' cyclical synchronization, the costs of not being able to counteract country-specific shocks through monetary policy may become larger and, thus, the net benefits of forming a currency union are reduced.

Theoretically, the effect of an increase in bilateral trade on the degree of business cycle synchronization could be positive or negative. In particular, this effect depends on the kind of shocks that dominate business cycle fluctuations and on the changes that trade integration may induce on the production structure of the trading countries. If demand shocks are dominant, we should expect trade integration to strengthen the transmission of shocks from one country to another, especially through the impact of these shocks on import demand. The presence of important industry-specific shocks, however, may strengthen or offset this effect. If, as a result of different comparative advantages, the increase in bilateral trade leads each country to specialize in different industries, the net effect of trade on business cycle correlation could become negative (Kenen, 1969; Eichengreen, 1992; Krugman, 1993). In contrast, if trade is mostly of an intra-industry nature, it should lead to a higher degree of business cycle synchronization (Frankel and Rose, 1998).

Since theory alone cannot determine the direction of the effect of trade integration on business cycle synchronization, some authors have tried to identify the nature of this effect empirically. Frankel and Rose (1998), Artis and Zhang (1999), Anderson, Kwark and Vahid (1999), Clark and van Wincoop (2001) and Gruben, Koo and Millis (2002) provide evidence supporting a positive effect of trade intensity on business cycle synchronization. With the exception of Anderson, Kwark and Vahid (1999), however, these studies are based on data that includes only industrialized countries, where most trade is of an intra-industry nature. In this context, Imbs (1999 and 2000) suggests that it may be the similitude in the production structure across countries, more than the volume of trade, what may be driving business cycle synchronization. In this same vein, Firdmuc (2001) provides evidence that, within the group of OECD countries, only trade that is of an intra-industry nature leads significantly to a higher cyclical synchronization.

If, as the previous two authors suggest, the findings of a positive effect of trade on business cycle synchronization reflect the fact that most trade between industrial countries is of an intra-industry nature, then we should not expect such an effect for industrial-developing country pairs that engage in trade integration, unless their trade is also predominantly intra-industry in nature. Empirical evidence concerning the possible links between trade and business cycle synchronization for developing countries that become commercially integrated with industrial countries, however, is mostly absent in the literature.¹ Analyzing these links would yield relevant insights concerning the mechanisms through which trade integration may affect business cycle fluctuations and, thus, would provide important elements to evaluate the welfare effects of the increasing economic integration observed around the globe.

In this paper, we try to fill this gap by analyzing the effect of the North American Free Trade Agreement (NAFTA) on the degree of business cycle synchronization between Mexico and the United States (U.S.).² NAFTA is an excellent case study to analyze the implications of trade integration between a developing and an industrial country. Several authors have already documented an important increase in the correlations of Mexico and U.S. economic activity aggregates after NAFTA was enacted (Cuevas, Messmacher and Werner, 2003; Torres and

¹ An exception is Calderón, Chong and Stein (2003), who extend Frankel and Rose's (1998) analysis to include bilateral trade between industrial-developing and developing country pairs in the sample. They find that the overall effect of trade integration on business cycle synchronization is positive. However, this effect is apparently smaller for industrial-developing and developing country pairs than for industrial country pairs. Furthermore, they find that this effect is also decreasing in the degree of production structure asymmetry between the trading countries.

² We could also have analyzed if NAFTA may have had an effect on the degree of business cycle synchronization between Mexico and Canada. However, bilateral trade between these two countries represents a very small fraction of the trade that each of these countries conducts with the U.S.

Vela, 2003). These authors, however, base their conclusions on simple correlations and regression equations and do not conduct a more formal analysis concerning the extent to which the evidence suggests significant changes in the specific correlation patterns between the business cycle and low frequency components of Mexico and U.S. economic activity levels.³

Another related issue we address in this paper is the possible fragility that the business cycle synchronization between Mexico and the U.S. may be currently exhibiting. Given its initial comparative advantages, Mexico responded to trade integration mostly by specializing in unskilled labor-intensive processes conducted in footloose assembly manufacturing plants. As a consequence, the links this country achieved with the U.S. through trade could be easily weakened by the entry into global trade and investment flows of other unskilled labor-abundant countries that could reap off Mexico's initial comparative advantages. In contrast, if Mexico had specialized more extensively in higher value-added industries that entailed larger, dynamic economies of scale, it may have enjoyed a head start with respect to developing countries that entered the globalization process afterwards (see Krugman, 1987a and 1987b and Grossman and Helpman, 1995).

These concerns have recently acquired a new dimension with China's increasing role in the world's trade flows and, especially, after its accession to the World Trade Organization (WTO) on December of 2001 (Shafaeddin, 2002; Goldman Sachs, 2003; Cervera, 2004). Mexico stands out within the group of countries that may be most affected by China's competition since there is a large overlap in the kinds of products both countries have specialized in and, thus, their export mixes are very similar. It has been argued that, as China reaps off Mexico's comparative advantage in terms of low wages for unskilled labor, Mexico's disadvantages in terms of a lack of human capital and poor institutions, utilities and infrastructure may become more apparent (Rosen, 2003). This could cause some production facilities in industries in which Mexico previously enjoyed marginal comparative advantages to move from Mexico towards other unskilled labor-abundant countries, such as China.

The previous concerns are mostly related to the specific loss of Mexico's comparative advantages against China. The entry of China into international trade flows, however, may entail a larger, global dimension. As its less developed countries have liberalized to trade, the Asian block has gradually achieved significant production complementarities through the creation of its own regional production-sharing networks (Gereffi, 1999; Ng and Yeats, 1999;

³ Herrera (forthcoming) distinguishes the degree of correlation between Mexico and U.S. business cycles at both cyclical and long term frequencies. He finds that both economies share a common trend and a common cycle. His findings, however, correspond only to the 1993-2001 period, so he is unable to assess if these correlation patterns were different before NAFTA was enacted.

Shafaeddin, 2002). This has apparently affected the relative competitiveness of several North American regional production chains and, as a consequence, has caused a gradual movement of all stages of the production processes of some industries, such as apparel, towards Asia. China's entry into world trade flows may strengthen this process. Formal tests on the relevance of these events for Mexico and U.S. manufacturing sectors and their degree of synchronization, however, are lacking in the literature. In part, this may be a result of the difficulty of detecting a structural break in the long-run relationship between these sectors when such break may have occurred relatively recently.

To analyze the issues discussed above, in this paper we study the changes in the degree and the nature of the synchronization between Mexico and U.S. manufacturing production levels. In particular, we apply spectral analysis and cointegration tests to assess the correlation of the business cycle and low frequency components of the manufacturing production series of the two countries. We find that, before NAFTA started operating, a significant, but weak correlation between these series apparently existed only for business cycle frequencies. Moreover, during this period Mexican manufacturing production seemed to have followed its U.S. counterpart with a long lag. In contrast, after NAFTA was enacted, the correlation between Mexico and U.S. manufacturing sectors in business cycle frequencies became stronger, a significant long-run link between them seems to have evolved, and their cyclical movements tended to become contemporaneous. We therefore conclude that, even before NAFTA was enacted, temporary demand shocks in the U.S. may have been transmitted to Mexico through trade. However, a long-run relationship between the manufacturing sectors of these two countries seems to have been a consequence of NAFTA and, thus, is a relatively recent phenomenon. This suggests that the current synchronization of Mexico and U.S. business cycles is not driven only by the transmission of transitory demand shocks, but also by supply-side links derived from production-sharing schemes induced by NAFTA.

Once these results are established, we proceed to test if there is evidence to the effect that the cointegrating vector linking Mexico and U.S. manufacturing production levels has suffered a recent structural break. We show that standard tests for structural break in cointegrating vectors fail to reject the null of stability. This failure, however, can be a result of the fact that the break may have occurred near the end of the sample period. We therefore apply some recently developed tests that have greater power for the alternative of a structural break at the end of the sample period. The results reject overwhelmingly the null of stability, suggesting that the links between Mexico and U.S. manufacturing production levels derived from NAFTA may have recently become weaker. With the data currently available, however,

we cannot determine if this weakening took the form of a permanent negative level shift in Mexico's production levels or of a decrease in the elasticity with which Mexican production responds in the long run to its counterpart in the U.S. That is, we still do not have enough data under the new regime to test if the structural change in the cointegrating vector affected its constant or its slope coefficient.

We also provide disaggregated evidence suggesting that the production-sharing links may have become weaker precisely in the industries where Mexico has concentrated most of its trade with U.S. (metal products and machinery). We also show that, in some other important sectors where strong links seem to persist (textiles and apparel), the full North American region seems to have been affected by the global movement of these industries towards other regions of the world that have achieved important competitiveness gains.

The remainder of the paper is as follows. In section 1 we briefly describe the recent evolution of manufacturing production in Mexico and the U.S. and discuss the reasons why their cycles are thought to have become more synchronized after NAFTA was enacted. Section 2 conducts a spectral analysis to assess the nature and extent of business cycle synchronization between these countries. In Section 3 we apply some cointegration tests for Mexico and U.S. manufacturing production levels. Section 4 describes the results of structural break tests undertaken to assess if the long-run relation between Mexico and U.S. manufacturing sectors has been recently affected. Section 5 complements the findings of the previous parts by extending the analysis to the eight main manufacturing divisions.⁴ Finally, Section 6 summarizes our findings.

1. The synchronization between Mexico and U.S. business cycles

As a consequence of their proximity, Mexico and the U.S. already exhibited strong trade links much before NAFTA was enacted. Since the mid-sixties, Mexico allowed the creation of foreign owned *maquiladora* assembly plants with a duty-free treatment through its "Border Industrialization Program". These manufacturing plants allowed U.S. firms to take advantage of Mexico's proximity and lower wages to conduct routine, unskilled labor-intensive operations within their production processes. Indeed, these plants import virtually all materials from the U.S., use Mexican labor to conduct assembly activities, and re-export the final product. As a consequence of this program, some manufacturing-related trade was already

⁴ In Mexico's national accounting system, the manufacturing sector is divided into 9 broad industries: i) food, beverage and tobacco products; ii) textile, apparel and leather products; iii) wood products; iv) paper and printing; v) chemical products; vi) nonmetallic mineral products; vii) primary metal; viii) metal products and machinery; and ix) other manufacturing industries. In Section 5, we focus on the first eight of these industries.

taking place between Mexico and U.S. during the seventies and the first half of the eighties. However, the operation of *maquiladoras* was initially allowed only in a 20 kilometer zone along international borders and coastlines. Therefore, the creation of this program did not represent a true shift in the import substitution scheme that characterized Mexico's trade policy at the time and did not induce Mexico's overall manufacturing sector to become more export-oriented.

During the mid-eighties, however, Mexico signed on to the General Agreement on Tariffs and Trade (GATT) and opened up unilaterally to trade. The fundamental effect of this policy shift seems to have been that it enhanced the competitiveness of the Mexican manufacturing industry by allowing it to import capital goods and inputs at international prices (see Sánchez, 1992; Tybout and Westbrook, 1995; and Banco de México, 1998). This induced higher and more diversified export flows after 1985: while in 1982 non-oil exports accounted for roughly 30% of total exports of goods, by 1993, just before NAFTA started operating, these exports already represented close to 90% of the total. This export diversification was mainly driven by the steady increase in the export activity of the manufacturing sector. Furthermore, most international trade operations of Mexico were undertaken with the U.S. In 1993, manufacturing exports to the U.S. already represented 60% of all Mexican exports. Similarly, the U.S. contributed with 64% of overall Mexican imports.

The enactment of NAFTA strengthened the trade integration between Mexico and U.S. further. From 1994 to 2003, Mexican manufacturing exports to the U.S. increased by a factor of 3.6. As a consequence, these exports increased their share in total Mexican exports to 70% in 2000-2003. Furthermore, during this period Mexico's share in U.S. trade flows also increased significantly: the share of U.S. manufacturing exports going to Mexico rose from 9.6% to 14.2%. Similarly, the share of U.S. manufacturing imports coming from Mexico rose from 6.5% to 11%. Vargas (2000) provides examples of several important product categories where the specific tariff reductions implied by NAFTA led to large, immediate increases in bilateral trade between Mexico and the U.S.

While the presence of a positive effect of NAFTA on the volume of bilateral Mexico-U.S. trade seems undisputable, its alleged effect on the degree of business cycle synchronization faces some theoretical challenges. As discussed in the introduction, the effect of an increase in Mexico-U.S. bilateral trade on the degree of synchronization of their business cycles should be positive only if demand shocks are dominant or if trade is of an intra-industry nature. In this context, the transmission of U.S. demand shocks to Mexico may have been present even before NAFTA since, as already discussed, U.S. was by far the most important

market for Mexican exports before this treaty was enacted. However, Mexico and the U.S. exhibit large differences in their relative factor endowments. We should then expect trade between these two countries to be mainly driven by comparative advantage. If we rely on a traditional Heckscher-Ohlin framework, an increase in the degree of trade integration between Mexico and U.S. should tend to induce a higher degree of specialization across industries and, as a consequence, could diminish their business cycle synchronization.

This argument, however, misses the worldwide trend towards a vertical specialization in production chains, in which each country in a trade relationship has tended to specialize in particular stages of each good's production process (Feenstra, 1998; Hummels, Ishii and Yi, 2001). As less-developed countries have liberalized to trade and global transport costs have decreased, firms have become increasingly able to spread out geographically the different processes for the production of goods, in order to take advantage of differences in relative input prices across countries.⁵ In particular, firms in industrial countries have increasingly set up affiliate assembly plants or outsourced low-skill activities to local firms in unskilled labor-abundant countries, while more skill-intensive activities and the distribution of finished products have mostly remained in the skill-abundant home country headquarters. This has induced a specific "vertical" type of intra-industry trade, in which the same product may cross borders of countries that exhibit large differences in factor endowments several times during the manufacturing process. In this context, specialization occurs across processes within each industry, and not across industries. Under this framework, therefore, it is possible to reconcile intra-industry trade with a comparative advantage motive for trade (see Falvey, 1981; Falvey and Kierzkowski, 1985; Greenaway, Hine and Milner, 1995). Kose and Yi (2001a and 2001b), in fact, show that the empirical evidence concerning a positive effect of trade integration on business cycle correlation can be matched with the results of standard international trade-business cycle models only if this kind of intra-industry specialization pattern is introduced.

The specialization pattern described above has characterized Mexico-U.S. trade even before NAFTA was enacted. The *maquiladora* industry is, in fact, a clear example of this kind of regional production-sharing arrangements (Feenstra and Hanson, 1997). The increasing specialization of Mexican firms in unskilled labor-intensive assembly activities undertaken for American firms seems to have become a more general phenomenon within Mexico's

⁵ Feenstra (1998) treats trade liberalization, the reduction of transport costs and vertical specialization as conceptually distinct factors that may explain the increase in worldwide trade and, based on Baier and Bergstrand (1997), suggests that the combined effect of trade liberalization and transport cost reductions may account for two fifths of the increase in bilateral trade within OECD countries from 1958 to 1988. It is important to note, however,

manufacturing industry after this country opened up to trade in the mid-eighties (Hanson, 1996). Therefore, Mexico-U.S. trade has become predominantly intra-industry in nature, even in the face of the large differences in factor endowments that these countries exhibit.⁶ NAFTA seems to have reinforced this trend and, in particular, it boosted the formation of regional production-sharing arrangements between Mexico and the U.S. that, in fact, tended to induce not only more intra-industry, but also intra-firm trade between these countries.⁷ This reflects the fact that the signature of this treaty induced important changes in Mexico's foreign investment laws that reduced significantly the restrictions and disincentives for investors to create foreign-owned plants in most sectors of the economy. In this sense, NAFTA not only seems to have had even more important effects on the nature than on the volume of Mexico-U.S. trade, but it also boosted investment projects for the creation of manufacturing facilities destined to serve the whole North American market from Mexico (Graham and Wada, 2000). Once we look at Mexico-U.S. experience under this framework, it is clear why NAFTA may have indeed enhanced the business cycle synchronization between these countries.

The behavior of manufacturing production in Mexico and the U.S. seems to be consistent with the hypothesis that the business cycles of these countries have become more synchronized after NAFTA. Figures 1 and 2, respectively, exhibit the levels and annual growth rates of Mexico and U.S. monthly seasonally adjusted manufacturing production indexes from January, 1980 to February, 2004.⁸ These series did not exhibit an especially similar behavior from 1980 to 1993. In contrast, after NAFTA was enacted and the 1995 peso crisis was over, the series exhibit a fairly similar evolution.

From 2003 on, however, we observe that, while U.S. manufacturing production started to recover gradually from the 2001 recession, Mexico's manufacturing kept on exhibiting negative growth rates. The distinct evolution of Mexico and U.S. manufacturing activities seems consistent with the concerns discussed previously about a possible weakening of Mexico-U.S. production-sharing links. As mentioned before, the entry of other unskilled labor-

that the outsourcing of unskilled labor-intensive processes to less developed countries may well have been itself, at least in part, a consequence of trade liberalization and the decrease in transport costs.

⁶ According to data from the OECD (2002), Mexico exhibited one of the largest shares and fastest growth rates of intra-industry trade within OECD countries during the nineties. See also Ruffin (1999), Vargas (2000), Ekanayake (2001) and Moreno and Palerm (2001) for evidence on the large magnitude of Mexico-U.S. intra-industry trade.

⁷ During NAFTA negotiations U.S. firms explicitly argued that they needed a low-wage partner for routine, low-skill operations such as assembly, in order to compete with suppliers (or branches) of Japanese multinationals in less developed countries (Markusen and Zahnsier, 1997).

⁸ We use manufacturing production to assess the effects of trade on business cycle synchronization between Mexico and the U.S. for two reasons. First, most trade between these countries corresponds to manufactured goods. Second, a broader measure of economic activity, such as aggregate GDP, may exhibit a smaller cross-country correlation, due to the inclusion of non-tradable sectors in its measurement.

abundant countries into international trade flows and the reduction in international transport costs could make Mexico lose some of its original comparative advantages related to its proximity to the U.S. market and its abundance of unskilled labor. This has apparently been reflected in a loss of market share of Mexican exports in the U.S. market. Figure 3 exhibits the shares of Mexico and China in U.S. non-oil imports. A visual inspection seems to suggest that imports from China accelerated their pace from 2001 on, while Mexico started losing market share beginning in 2002. As a consequence, by 2003 China had become a more important supplier to this market. The loss of market share of Mexico against China has been especially strong in the industries of metal products and machinery (particularly in the case of computers and computer accessories, as well as VCRs and TV sets) and in apparel (Cervera, 2004).⁹ These two industries account for most of Mexico's manufacturing exports to the U.S.

These events seem to have had a permanent effect on Mexico's export-oriented sector. For instance, from 2001 to 2003, the number of active *maquiladora* plants in the country was cut by 25%. Most of this decrease reflects the closure of plants in the textile and apparel industry. We can observe in Figure 4 that most of these lock-outs occurred in a fairly short period of time, going from October, 2001 to March, 2002. This suggests that the closure of export-oriented plants in Mexico may have in part responded to specific events that took place during these months (as China's entry into the WTO).

In the following sections of the paper, we address these issues more formally. We study the correlation between both the business cycle and the low frequency components of Mexico and U.S. manufacturing production series. This distinction may allow determining if the synchronization between Mexico and U.S. production levels is driven fundamentally by temporary demand shocks or by longer-run links in the production side. Indeed, a demand-side link between these sectors would tend to show up in the data only as a high correlation in the business cycle frequency components of the series. In contrast, a long-run link based on production complementarities would tend to show up also in terms of cointegration between them. In this context, we first use spectral techniques to estimate the business cycle correlations that may be present in the data. We then test for cointegration between Mexico and U.S. manufacturing production indexes and assess if NAFTA may have led to a deeper, longer-term link between these sectors. Finally, we test, both from an aggregate and a disaggregated point of view, if these long-run links have recently become weaker.

⁹ Mexico's exports to the U.S. of transportation equipment have also been recently losing market share. However, this seems to be reflecting idiosyncratic factors in the industry and changes in the worldwide share of specific firms and not a result of Chinese competition (Cervera, 2004).

2. Spectral analysis

To analyze whether manufacturing production in Mexico has become more synchronized with its counterpart in the U.S. after NAFTA was enacted, we first focus on the behavior of manufacturing production within these countries at business cycle frequencies. If NAFTA strengthened the links between Mexico's and the U.S. manufacturing sectors, we should expect to observe an increase in the correlation of the cyclical components of manufacturing production of both countries after this treaty was enacted.

To study this issue, it is natural to rely on spectral analysis (see Fuller, 1976; Hamilton, 1994; Koopmans, 1995 or Warner, 1998). The basic idea underlying this approach is that a time series may be decomposed into a finite number of orthogonal components, each representing cycles of a particular frequency. In this context, we may analyze the degree of association exhibited by different cyclical components of two time series, where these components differ in terms of the periodicity of their corresponding cycles. In particular, the squared coherence of two series measures the proportion of the variance of either series that can be explained linearly by the other for cycles corresponding to each particular frequency. Formally, if $g_{x,y}(\omega)$ is the estimated cross-spectrum (i.e. the smoothed cross-periodogram) and $g_{x,x}(\omega)$, $g_{y,y}(\omega)$ are the estimated spectra for X and Y at frequency ω , respectively, the estimated squared coherence between these series is:

$$s_{x,y}(\omega)^2 = g_{x,y}(\omega)^2 / g_{x,x}(\omega)g_{y,y}(\omega) \quad (1)$$

Thus, a plot of the coherence between Mexico and U.S. manufacturing indexes identifies to what extent, and within which frequency bands, manufacturing production fluctuations in Mexico have been correlated with fluctuations in manufacturing production in the U.S.

Figure 5 exhibits estimates of the coherence between the detrended logs of manufacturing production indexes in Mexico and in the U.S. for the periods 1980-1993 and 1996-2003.¹⁰ Coherences exhibited in the figure are roughly significant at a 5% (1%) level if

¹⁰ As we will see below, we are unable to reject the hypothesis that the manufacturing production series in Mexico and the U.S. are I(1). Therefore, to conduct this analysis we detrended the series with the Hodrick-Prescott filter. We used quarterly data for this analysis. If the coherence was instead computed with monthly data, all business cycle frequencies would be bunched within the leftmost sixth segment of the graph, making it difficult to interpret the results. We dropped 1994-1995 from the post-NAFTA sample to prevent the 1995 peso crisis from affecting the results. Gerlach (1998) conducts a very similar analysis to the one undertaken in this section to study the

they are higher than 0.51 (0.61). These cutoffs were computed using Fuller's (1976) suggested F test for the null hypothesis that the coherence is zero.

According to these results, coherence increased significantly in 1996-2003 with respect to the preceding period for most frequencies. This suggests that the correlation between Mexico and U.S. manufacturing cycles became significantly stronger after NAFTA was enacted. It is also important to note that, while coherence estimates for 1996-2003 are statistically significant for most frequencies, those for the 1980-1993 period seem to be statistically significant only within a band of cycles with periodicities from around 2 to 12 years. Moreover, if we focus on low frequencies we observe that, while for 1996-2003 coherence is consistently high, during the preceding period it tended to become smaller for cycles closer to the zero-frequency (long-run) line. While we will analyze this issue more formally in the next section, these results suggest that the long-run association between Mexico and U.S. manufacturing industries strengthened considerably only after NAFTA was enacted.

Another important result is that, for low-frequency cycles, variations in Mexican manufacturing production seemed to exhibit important lags with respect to movements in U.S. manufacturing production from 1980 to 1993. In contrast, for 1996-2003, cyclical movements in manufacturing production in both countries seemed to be roughly contemporaneous. This can be observed in Figure 6, which exhibits the estimates of the phase lead of Mexico's manufacturing production with respect to the U.S. The graph includes only the estimates for frequencies corresponding to cycles of 1 year length or more, where the estimated coherence is relatively high. The reason we did not include higher frequencies is that the phase lead is poorly estimated and becomes meaningless when the coherence is small. We can observe that, for 1980-1993, Mexican manufacturing production exhibited an important, negative phase lead with respect to U.S. manufacturing production for low frequencies. In contrast, for 1996-2003, the phase lead estimate is very close to zero for all frequencies, suggesting that cycles across both countries were roughly synchronized in time during this period.

The results described above are consistent with the idea that, before NAFTA, the cyclical correlation between Mexico and U.S. manufacturing production was driven mainly by the transmission of temporary demand shocks from the U.S. to Mexico. While this transmission mechanism seems to have still been present after NAFTA was enacted, it appears that a more permanent, long-run association between Mexico and U.S. manufacturing sectors arose as a consequence of this treaty. This apparently led to a stronger synchronization in

changes in the correlation of cyclical components of industrial production indexes of several industrial countries from the Bretton-Woods to the flexible exchange rates regimes.

Mexico-U.S. manufacturing cycles and may be a reflection of the production-sharing schemes induced by NAFTA.

We now follow Stock and Watson's (1999) approach and use Baxter and King's (1999) band-pass filter to extract from the manufacturing production series for Mexico and the U.S. a business cycle component, which includes cycles that have periodicities between 6 quarters and 8 years, and a trend component, including all cycles with lower frequencies. Figures 7 and 8 exhibit the filtered series. The behavior of these series confirms the arguments made above. With respect to the business cycle component we observe that, before NAFTA, cyclical movements in U.S. tended to lead Mexico's cycle. However, U.S. and Mexico's cycles were not closely related. In contrast, after NAFTA was enacted and the 1995 peso crisis was over, we observe a close, contemporaneous correlation between Mexico and U.S. manufacturing production cycles. The trend component also yields some important insights. There appears to be a closer relationship between Mexico and U.S. at low frequencies after NAFTA was enacted than in the 1980-1993 period.

The results of this section suggest that Mexico and U.S. manufacturing production levels may have become cointegrated only after NAFTA started operating. Indeed, Levy (2002) shows that if two series are cointegrated, then at zero frequency their coherence should be one and their phase shift should be zero. According to the results of this section, these features appear to hold for the case of Mexico and U.S. manufacturing production levels only after NAFTA was enacted. We therefore test the cointegration hypothesis formally in the next section.

3. Cointegration tests

If the high synchronization between Mexico and U.S. manufacturing production cycles found above for the post-NAFTA period reflects links derived from production-sharing arrangements and not only the transmission of transitory demand shocks, we should expect these series to be linked in terms of their low frequency, long-run behavior. To address this question, in this section we apply a battery of cointegration tests. In all cases, the null hypothesis is an absence of cointegration between Mexico and U.S. manufacturing production levels.¹¹ The tests include Engle and Granger's (1987) two-step residual-based *ADF* test, Campos, Ericsson and Hendry's (1996) ECM-based test and Johansen's (1991, 1995)

¹¹ We conducted a set of unit root tests for the logs of monthly seasonally adjusted Mexico and U.S. manufacturing production indexes, using a sample going from January, 1980 to February, 2004. The results consistently failed to reject the null of a unit root, suggesting that these series are $I(1)$.

maximum eigenvalue test. These tests do not consider regime shifts under the alternative hypothesis and, therefore, could suffer from low power if the alternative corresponding to cointegration with a structural change is true. Therefore, we follow Gregory and Hansen (1996) and complement the analysis with their extensions to the conventional *ADF* and to Phillips' (1987) Z_α and Z_t tests, which allow for the possibility of a regime shift in the cointegrating vector in a single unknown point of time during the sample period.¹² The tests are applied to monthly data of the logs of seasonally adjusted manufacturing production indexes in Mexico and United States. The sample covers the period from January, 1980 to February, 2004. We also test cointegration for some sub-periods within this sample. In particular, we apply the tests listed above for the full sample, for the pre-NAFTA years (1980-1993), for the pre-NAFTA period when Mexico had already opened up unilaterally to trade (1986-1993) and for the post-NAFTA years (1996-2004).¹³

Table 1 summarizes the results. A clear pattern emerges from these tests. The null hypothesis of no cointegration cannot be rejected neither for the full sample, nor for the periods 1980-1993 or 1986-1993, with any of the tests applied. In contrast, all tests, with the exception of Engle and Granger's (1987) procedure, reject the null hypothesis for the sample covering 1996-2004. These results suggest that manufacturing production in Mexico became cointegrated with its U.S. counterpart only after NAFTA was enacted and support the hypothesis that the higher degree of synchronization between Mexico and U.S. cycles observed during this period may be a reflection of the production-sharing schemes that NAFTA promoted.¹⁴ In this context, it is interesting to note that the long-run elasticity of manufacturing

¹² The procedure behind Gregory and Hansen's (1996) approach is the following. For each possible regime shift date within a period comprising 15% and 85% of the sample, the conventional *ADF*, Z_α and Z_t tests are computed from an OLS regression including a dummy variable that controls for the regime shift. The smallest values of these tests correspond to their ADF^* , Z_α^* and Z_t^* tests. A similar approach is taken by Zivot and Andrews (1992) to handle the possibility of regime shifts in unit root tests.

¹³ We dropped the observations for 1994 and 1995 from this last sub-sample to prevent the peso crisis of 1995 from affecting the results.

¹⁴ A caveat concerning this interpretation must be kept in mind. NAFTA was not the only relevant change in the Mexican economy that took place since 1994 and that could affect business cycle synchronization with the U.S. First, since 1995 Mexico adopted a flexible exchange rate regime. This could theoretically also induce a larger synchronization with the U.S. business cycle (see Flood and Marion, 1982). Second, after the 1995 peso crisis was over, Mexico stabilized its economy and achieved a successful disinflation. This may have allowed this country to become decoupled from other emerging markets and, as a consequence, to be more resilient to the contagion of shocks from these economies. Therefore, the relative relevance of shocks derived from Mexico's own macroeconomic instability or originated in other emerging markets may have decreased, allowing Mexico's business cycle to reflect to a larger extent shocks originated in the U.S. It is important to note, however, that at least part of this macroeconomic stability may have itself been a result of the improvement in business expectations, the larger foreign direct investment flows and the reduced vulnerability to terms of trade shocks that NAFTA may have promoted.

production in Mexico with respect to manufacturing production in the U.S. during this period is found to be very close to one.

Other results obtained for the 1996-2004 period are also worthwhile to mention. The adjustment coefficient for the error correction mechanism of the VAR system estimated with Johansen's (1991, 1995) procedure is only significant for the case of Mexico's manufacturing production. Moreover, unreported Granger causality tests suggest that, during this period, causation was unidirectional, going from U.S. manufacturing production to Mexico's manufacturing production. These results suggest that U.S. manufacturing is strongly exogenous with respect to Mexico's manufacturing production and imply that Mexico's manufacturing production tends to adjust to shocks in U.S. manufacturing production, and not vice-versa. In terms of Gonzalo and Granger's (1995) framework, the results suggest that the common stochastic trend that drove Mexico and U.S. manufacturing production levels during this period was completely originated by U.S. economic activity.

It is important to note that Engle and Granger's (1987) *ADF* test is unable to reject the null hypothesis for 1996-2004, while Gregory and Hansen's (1996) *ADF** test rejects the null for that same period. As these authors argue, this combination of results suggests that a structural break in the cointegrating vector may have occurred during this period.¹⁵ This could be consistent with a recent weakening of the cointegrating relationship between Mexico and U.S. manufacturing production. Indeed, Gregory and Hansen's (1996) tests for this period consistently identify the timing of a possible structural change as November of 2002, which is close to the period when the links between Mexico and U.S. manufacturing production levels apparently became weaker. We test this hypothesis more formally in the next section.

4. Structural break tests

Given the results found in the previous section, we now focus on the 1996-2004 period, in which cointegration between Mexico and U.S. manufacturing production levels seems to hold. We first explore to what extent the data supports the null hypothesis of no structural break in the cointegrating vector linking manufacturing production in Mexico and in the U.S., without specifying the form of the alternative hypothesis. To do this, we use the diagnostic test proposed by Hao and Inder (1996). This test is an extension for the case of cointegrated regression models of the CUSUM test with OLS residuals originally proposed by Ploberger

¹⁵ Campos, Ericsson and Hendry's (1996) ECM-based test also rejects the null for this period. These authors show that their test is more powerful than the residual-based test of Engle and Granger (1987) under an alternative of cointegration with a regime shift.

and Krämer (1992) for stationary regressors. The test uses the residuals from a cointegrating vector estimate based on the fully modified (FM) OLS estimator due to Phillips and Hansen (1990).¹⁶ In particular, the FM-OLS based CUSUM test statistic is:

$$B^{(T)}(\tau) = \frac{1}{\hat{\omega}_{0,1}\sqrt{T}} \sum_{t=1}^{[T\tau]} \hat{u}_t^{+(T)} \quad (2)$$

where T is the sample size, τ identifies the fraction of the sample for which the statistic is computed, $\hat{\omega}_{0,1}^2$ is the estimate of the long run variance of the error, $[T\tau]$ denotes the integer part of $T\tau$, and $\hat{u}_t^{+(T)}$ is the FM-OLS residual for observation t . Basically, this statistic corresponds to a standardized partial sum of FM-OLS residuals. We reject the null for large values of $\sup_{0 < \tau < 1} |B^{(T)}(\tau)|$.

Table 2 summarizes OLS and FM-OLS estimates of the cointegrating vector relating seasonally adjusted logs of Mexico and U.S. manufacturing production indexes from January, 1996 to February, 2004. We normalize Mexico's coefficient to be one, so that the slope coefficient corresponds to the estimate of the long-run elasticity of Mexico's manufacturing production with respect to its U.S. counterpart. Consistently with the previous results, the estimated elasticity of Mexico's manufacturing production with respect to its U.S. counterpart for this period is found to be close to one. We observe that the FM-OLS estimate for this elasticity turned out to be similar to the OLS estimate for this parameter.

The additional set of statistics included in the table will be discussed below. For now, we focus on Hao and Inder's (1996) test. Figure 9 plots the FM-OLS based CUSUM test statistics for this estimate, along with 10%, 5% and 1% critical values for rejection. The test tends to exhibit high values around the period between August, 2002 and February, 2003. This suggests that, if it exists, the structural break may have occurred around that period. Note that the estimated breakpoint from Gregory and Hansen's (1996) tests conducted above corresponds precisely to the midpoint of this period. The CUSUM test statistic, however, falls slightly short of reaching a 10% significance level.

We next test for a shift in the cointegrating vector using Hansen's (1992) L_c , $MeanF$ and $SupF$ tests.¹⁷ All of these test the null of a stable cointegrating vector. However, they differ

¹⁶ When a group of I(1) time series is cointegrated, OLS yields super-consistent estimates of the cointegrating vector (Stock, 1987). However, the asymptotic distribution of the estimates depends on nuisance parameters arising from the possible endogeneity of the regressors and serial correlation of the residuals. The FM-OLS

in terms of the alternative hypothesis. The *SupF* test has as alternative hypothesis a swift shift in the cointegrating vector at an unknown point in time. In contrast, the alternative for the other two tests corresponds to the case when the coefficients of the cointegrating vector follow a martingale process and, therefore, reflects an unstable, gradually evolving cointegrating relationship. Indeed, the L_c test may be interpreted as testing the null of cointegration against the alternative of no cointegration. In practice, however, the three tests tend to have power in similar directions. As before, the tests are based on the FM-OLS estimate of the cointegrating vector.

The results are summarized in Table 2. In all cases, the tests are unable to reject the null of a stable cointegrating vector. In fact, none of the tests is significant even at a 20% level. These tests, however, may suffer from low power when the structural break occurs close to the end of the sample.¹⁸ This is apparently the case. Figure 10 exhibits the sequence of F tests for structural change used for the computation of the test statistics. We observe that the values for these tests are especially large at the end of the sample. In particular, the *SupF* value is attained in the November 2002 observation, the same breakpoint date estimated by Gregory and Hansen's (1996) tests discussed previously. This date corresponds to the next-to-last observation used to compute the *SupF* and *MeanF* test statistics (the last 15% observations in the sample correspond to the period going from January 2003 to February 2004).

Given the discussion above, we now apply the end-of-sample cointegration breakdown tests recently developed by Andrews and Kim (2003). These tests are specifically designed to test the null of no structural break when the period for which the structural change may have occurred is relatively short.¹⁹ Therefore, they may be more powerful to detect a recent structural change in the cointegrating relationship between Mexico and U.S. manufacturing production than the tests applied before. Indeed, Andrews and Kim's (2003) tests are asymptotically valid when the length of the post-breakdown period, m , is fixed and finite as the total sample size $T + m$ goes to infinity. In contrast, the tests applied previously assume that the

estimator deals with this problem by applying a non-parametric correction to the OLS estimator of the cointegrating vector.

¹⁷ The *SupF* and *MeanF* tests are, respectively, the maximum and the mean values of the sequence of Chow structural break tests (using the variance estimate for the full sample) for each possible breakpoint in the period going from 15% to 85% of the sample. Critical values differ from the case when the timing of the structural break is known.

¹⁸ Other structural change tests for cointegration regressions, such as Quintos and Phillips (1993), suffer from this same problem.

¹⁹ The structural change may take the form of either a shift in the cointegrating vector or a shift in the error process from being $I(0)$ to $I(1)$ or both. That is, under the alternative the cointegrating relationship may have suffered a structural break or a full breakdown. All tests developed by Andrews and Kim (2003) have power for both alternatives.

cointegration breakdown period is relatively long and rely on asymptotics in which the length of this period increases to infinity with the sample size.

The first set of tests proposed by Andrews and Kim (2003) is motivated by the F statistic for parameter change in a regression with *iid* normal errors and strictly exogenous regressors. In particular, test P_a is the sum of squared post-break residuals $\{\hat{u}_s : s = T + 1, \dots, T + m\}$, where these residuals are evaluated with the use of an estimate of the cointegration relationship based on the first T (pre-break) observations:

$$P_a = \sum_{s=T+1}^{T+m} (\hat{u}_s)^2 \quad (3)$$

As the authors show, this test tends to over-reject the null. Thus, they propose some adjustments to the test statistic and its critical values to yield better finite-sample properties. These adjustments give rise to the tests P_b and P_c . In effect, the residuals used to compute P_b are based on an estimator of the cointegration relationship that uses the first half of the post-breakdown period, in addition to the first T observations. P_c has the same form as P_a and P_b , but the residuals used for its computation are based on an estimate of the cointegrating relationship that uses the full $T + m$ sized sample.

The second group of tests is derived from the locally best invariant test for a shift in the error distribution from being *iid* normal for all observations to being *iid* normal for the first T observations and then a normal unit root process for the last m observations. In particular, the R_a test statistic is given by the sum of squared reverse partial sums of the post-break residuals, evaluated using a pre-break estimate of the cointegrating vector:

$$R_a = \sum_{t=T+1}^{T+m} \left(\sum_{s=t}^{T+m} \hat{u}_s \right)^2 \quad (4)$$

In parallel to the case of P_a , this test tends to over-reject the null when it is true. Thus, similar adjustments are made to this test to obtain better finite-sample properties. These adjustments yield the tests R_b and R_c .

Critical values for the tests described above are computed with the use of a parametric subsampling method introduced by Andrews (2003). Briefly, the first T observations in the sample are used to compute a set of $T - m + 1$ versions of the test statistic of interest, but

where each corresponds to the test for cointegration breakdown over some group of m observations that fall within the time period where no breakdown has occurred under both the null and the alternative. The distribution of these $T - m + 1$ versions of the test statistic allows computing p -values for the test statistic of interest. In particular, the $1 - \alpha$ sample quantile of these statistics is the α critical value for the end of sample breakdown test statistic.²⁰ According to Andrews and Kim (2003), from all the tests described above, the P_c and R_c tests have better finite-sample properties in terms of power and lack of size distortion.

Table 3 summarizes the p -values for Andrews and Kim (2003) tests applied to the residuals from OLS estimates of the cointegrating vector for the logs of monthly seasonally adjusted manufacturing production indexes in Mexico and the U.S. The sample starts on January of 1996. We chose January, 2002 and January, 2003 as possible structural break dates. The first choice corresponds to the first full month after China became a member of the WTO. The second date was chosen to test if the evolution of Mexico's manufacturing production during 2003 was sufficiently distinct from that observed in the U.S. to support the hypothesis of a recent weakening of Mexico-U.S. business cycle synchronization. It is important to recall that most of the tests conducted previously suggest that the structural break, if it exists, may have occurred around the end of 2002.

The results suggest that a structural change in the cointegrating relationship between manufacturing production in Mexico and in the U.S. may have taken place at the end of the sample period. In particular, all tests reject the null of no structural change after January of 2003. In contrast, the null is rejected only with some of the tests and, in general, with larger p values, when the breakpoint under the alternative is assumed to be January, 2002.

The structural change that the cointegrating relationship between Mexico and U.S. manufacturing production indexes seems to have suffered apparently takes the form of a decrease in the elasticity of the former with respect to the latter. In particular, as can be observed in Figure 11, the estimate of this elasticity for the full sample is smaller than the estimates obtained from each of the subsamples from 1996 to 2002 used to compute P_c for the case when the structural break is assumed to have occurred in January, 2003. That is, including the information contained in the period from January, 2003 to February, 2004 to estimate this elasticity yields a smaller value than the one obtained if this period is dropped from the

²⁰ To avoid under-rejection in the case of statistics P_c and R_c , an adjustment to the subsample statistics used to compute critical values is made. In effect, for these two statistics, critical values are based on the distribution of subsample statistics in which, for each m -sized period, only $m/2$ observations are dropped to compute the cointegrating vector. See Andrews and Kim (2003) for details.

estimation.²¹ It is important to note, however, that with the data currently available, we cannot determine if this structural break represents a true, permanent change in the elasticity with which Mexico's production responds in the long run to its U.S. counterpart. The data could as well be consistent with a decrease in the relative level of Mexico's manufacturing production. That is, we still do not have enough data to determine if the structural break that we have apparently detected took the form of a change in the constant or in the slope of the cointegrating vector.

Another approach that can be taken to assess the nature and significance of this possible structural change is to compare the recent behavior exhibited by Mexico's manufacturing output with the behavior that would be predicted if its relationship with U.S. production levels had remained unchanged. As already discussed, the results of Johansen's (1991, 1995) procedure conducted in Section 3 suggest that U.S. manufacturing production is strongly exogenous with respect to Mexico's manufacturing output. This, along with the evidence supporting cointegration between these series for the post-NAFTA period, implies that it is possible to specify a dynamic equation describing Mexico's manufacturing production as a function of the behavior of its counterpart in U.S. If no structural change had occurred, this equation could be used to forecast the changes in Mexico's manufacturing production as a function of lagged changes of Mexico and U.S. manufacturing outputs and of the lagged disequilibrium with respect to their long-run relationship (i.e. of an error correction mechanism).

In this context, we followed a general-to-specific specification methodology to estimate a dynamic equation describing Mexico's monthly manufacturing production for the period going from January, 1996 to December, 2002 (see Hendry (1995) for a description of this modeling approach). Note that we are leaving the observations going from January, 2003 to February, 2004 out of the estimation period in order to conduct out-of-sample forecasts. The estimation entailed several steps. We first ran an OLS regression of the log of the seasonally adjusted manufacturing production index in Mexico against the log of the seasonally adjusted manufacturing production index in U.S. The first lag of the residuals from this levels regression was then included as an additional regressor in a regression of the monthly change of the manufacturing production index in Mexico against 12 lags of the changes in Mexico and

²¹ An OLS estimate of this elasticity using the data from January, 1996 to December, 2002 yields a highly significant value of 1.129. In contrast, the estimate of this elasticity using a sample going from January, 2003 to February, 2004 yields a statistically insignificant value of 0.33. Nevertheless, it is evident that not much should be read into this last figure, as the post-breakpoint estimate of this elasticity is based on a sample covering only 14 months.

U.S. manufacturing production indexes. This equation was then reduced to a parsimonious representation by sequentially eliminating insignificant regressors, until only those with a p value of .10 or less remained in the specification. Finally, the resulting equation was re-estimated with non-linear least squares to identify the long-run elasticity and the short-term adjustment coefficients. The results are summarized in Table 4. Note that, as found before, the equation suggests that the long-run elasticity of Mexico's manufacturing production with respect to its U.S. counterpart is very close to one.

If the relationship between Mexico and U.S. manufacturing production levels had not suffered a structural break after 2002, the out-of-sample forecasts from this equation should not differ significantly or systematically from Mexico's manufacturing production observed behavior. Figure 12, however, shows that the forecasts from this equation consistently over-estimated Mexico's manufacturing production levels for 2003 and the first two months of 2004. Furthermore, note that the observed manufacturing output levels fell below the lower limit of the 95% interval for the forecasts in several months. This is consistent with the fact that, as can be observed in Table 4, the Chow forecast stability test for the estimated equation overwhelmingly rejects the null of no structural change. As discussed previously, these results suggest that the recent structural change in the link between Mexico and U.S. manufacturing activities may have caused a smaller-than-expected response of Mexico's manufacturing output to the recovery of the U.S. manufacturing sector observed in 2003 and is consistent with the idea that the links between Mexico and U.S. manufacturing activities may have recently become weaker.

5. Disaggregated evidence

Another possible approach to test if the links between Mexico and U.S. manufacturing sectors have recently become weaker is to look at disaggregated evidence. To the extent that the increasing competition from other unskilled labor-abundant countries may have affected the Mexico-U.S. links in some particular industries earlier or in a larger magnitude than in others in which Mexico may still exhibit comparative advantages, this analysis may yield more powerful evidence concerning this hypothesis. Furthermore, this analysis may also be helpful to identify in which industries a larger competition from Asian production networks may be affecting the production levels of North American product chains that may remain strongly integrated.

In this section we undertake this analysis. We first summarize the relative importance of each of eight broadly defined manufacturing industries on Mexico's overall manufacturing

output and exports to the U.S. We also compare the recent behavior of Mexico's and China's shares in U.S. imports of products corresponding to each industry. We then proceed to analyze to what extent the evidence supports the hypothesis that, within each specific division, the link between Mexican exports to the U.S. and the output level in the U.S. may have recently become weaker. This entails two steps. First, we test for cointegration between U.S. imports from Mexico and output levels for each manufacturing industry in the period from January, 1996 to February, 2004.²² We analyze the link between imports from Mexico and U.S. output levels, instead of the possible links between output levels in both countries, to focus specifically on long-run links in the production side and to avoid domestic shocks in Mexico to influence the results. This distinction is especially important in some divisions where a large fraction of Mexican output is not exported. Once establishing the evidence concerning the existence of long-run links between imports from Mexico and U.S. output levels, in the second step we test for structural instability in the corresponding cointegrating vectors.

Table 5 summarizes the breakup of Mexico's manufacturing output and exports to the U.S. by division, as well as the ratio of exports to the U.S. to Mexican value added for each industry.²³ In decreasing order of importance, the divisions that contribute most to Mexico's manufacturing production are metallic products and machinery (including computer, electrical, electronic and transportation equipment and parts), food, beverage and tobacco products, chemical products, textiles and apparel and nonmetallic mineral products. From these industries, the most export-oriented ones seem to be metallic products and machinery, textiles and apparel and, to a smaller extent, chemical products. In contrast, the food, beverage and tobacco sector, although an important contributor to local output, seems to be mostly oriented to the domestic market. Thus, the main links between Mexico and U.S. manufacturing sectors

²² Using a broad set of different unit root tests, we could not reject the null of a unit root for any of the U.S. output series. Most of these tests suggested that the series measuring imports from Mexico are also I(1). Interestingly, however, some tests provided evidence that the imports of food, beverage and tobacco, wood, chemical and primary metal products could be trend-stationary instead. This evidence was stronger when we used tests that allow for a structural break under the alternative (Zivot and Andrews, 1992). We will nonetheless test for cointegration in these industries too. As will be shown, for the cases of imports of food, beverage and tobacco and of wood products, the evidence of cointegration with U.S. output levels will be rather weak. This is consistent with the possibility that the imports series may be trend-stationary while the output series are difference-stationary. In contrast, we will find stronger evidence of cointegration for the cases of chemical and primary metal products, which suggests that the imports series of these products may not be truly trend-stationary. With a finite sample, it may be difficult to distinguish trend against difference stationarity, and this appears to be the case for these industries. It is important to note, however, that all product categories in which the cointegration results are dubious are not very relevant for the analysis we conduct in the paper. The most important manufacturing divisions, in terms of their large share in Mexico's exports to the U.S., are the textile and apparel and metal products and machinery industries. For these two cases, we never found any evidence suggesting the imports series could be trend-stationary.

seem to arise mostly from the activities of the metallic products and machinery and the textiles and apparel divisions. These two industries account for almost 90% of Mexican manufacturing exports to the U.S.

Table 6 exhibits the recent evolution of Mexico's and China's shares in U.S. imports of each manufacturing division. China has been increasing its market share continuously in all product categories. In contrast, Mexico seems to be losing market share in most divisions. In the case of textile and apparel products, the downward trend of Mexico's share is observed long before China entered the WTO. In contrast, Mexico's share in U.S. imports of metallic products and machinery exhibited an upward trend until 2001. After that year, these exports have been gradually losing market share too.

We now test for cointegration between Mexico's exports to the U.S. and U.S. output levels of each manufacturing division. In order to capture secular trends in U.S. imports from Mexico or the influence of omitted factors, we include a deterministic trend in the cointegrating vectors under the alternative hypothesis. That is, in this section we are allowing for the weaker concept of stochastic cointegration under the alternative.²⁴ The inclusion of a trend seems especially relevant in the case of textiles and apparel, where imports from Mexico were increasing at a relatively fast rate up to year 2000, possibly reflecting the increasingly common production-sharing schemes developed after NAFTA was enacted (see Figure 13, panel (b)).

We conduct the same battery of cointegration tests that we applied for aggregate data in Section 3. All tests are applied to seasonally adjusted monthly data going from January, 1996 to February, 2004 on the logs of constant-dollar U.S. imports from Mexico and U.S. production indexes by industry group, aggregated to match Mexico's industrial classification.²⁵ The results are summarized in Table 7. In most cases, the evidence supports the presence of strong long-run links between Mexican exports and U.S. output levels. In fact, for some industries the null of no cointegration tends to be rejected even without considering structural change under the alternative (chemical products, paper and printing and primary metals). Note, however, that for some other industries the null of no cointegration is not rejected with most tests unless

²³ The ratio of exports to value added may be higher than 100% since exports include the full value of the finished product, while the denominator accounts only for value added generated in Mexico. Thus, this ratio may exceed 100% whenever the imported content of exports is sufficiently high, as in the *maquiladora* industry.

²⁴ We also tested for the significance of a deterministic trend in the cointegrating relationship for the aggregate data used in the previous sections. In contrast with the tests for disaggregated data applied here, in that case the trend never appeared to be significant. See Campbell and Perron (1991), Ogaki and Park (1997) or Harris, McCabe and Leybourne (2002) for the distinction between the concepts of deterministic and stochastic cointegration.

structural change is allowed for. This suggests that structural breaks in the links between some of Mexico and U.S. manufacturing divisions may have been important during this period. This possibility seems to be especially relevant in the case of metal products and machinery, in which no test rejects the null unless structural change is allowed for in the alternative.

The exception to the results described above seems to be the wood products industry, in which none of the tests rejects the null of no cointegration at a 5% significance level, even after allowing for structural change. In fact, even when Johansen's (1991, 1995) procedure provides weak evidence of cointegration (at a 10% level), the estimate for the elasticity of U.S. imports of wood with respect to this country's output level derived from this procedure is not statistically significant. These results are consistent with the hypothesis that U.S. imports of wood products from Mexico may be trend-stationary (see footnote 22). Thus, the results for this particular industry should be taken with caution. As we will see below, further evidence will suggest that this same situation appears to characterize the case of food, beverage and tobacco products.

We now test for structural breaks in the cointegrating vectors relating Mexican exports and U.S. output levels. We apply Hansen's (1992) instability tests, along with Andrews and Kim's (2003) end-of-sample cointegration breakdown tests. Table 8 summarizes the FM-OLS estimates of the cointegrating vectors and Hansen's (1992) instability tests. The estimates for the slope coefficients are positive for all industries. Furthermore, with the exception of the food, beverage and tobacco and wood products industries, all these estimates are statistically significant at a 1% level.²⁶ The values of these estimates seem rather high for the textile and apparel and chemical industries. These high elasticities, however, may be reflecting a relatively high response of Mexico's exports to the U.S. as new production-sharing schemes were formalized in the first years of the sample. For most of the other industries, the estimates for these elasticities are found to be somewhat larger than one.

Hansen's (1992) tests suggest instability in the cointegrating vectors only for the cases of wood products, chemical products, primary metals and metal products and machinery. As mentioned before, the case of wood products must be taken with care, since these tests are strictly valid only for I(1) variables. Concerning the case of chemical products, the break detected by the *SupF* test corresponds to a one-time level shift in imports in December, 1997 (see Figure 13, panel (e)). In unreported results we found that, once accounting for this shift,

²⁵ U.S. imports were converted to constant-dollar values at the product level, using specific producer price indexes for each product category.

²⁶ This is further evidence that, for these two industries, the results should be taken with care. In particular, in these cases the imports series seem to be trend-stationary (see footnote 22).

the tests are unable to reject the null of stability for this industry. In contrast, the rejections of the *MeanF* tests for the case of primary metals and metal products and machinery suggest some more general form of instability. In the first case, this may be related to the especially high volatility exhibited by the imports series (see Figure 13, panel (g)). In the second case, however, the rejection of the stability null seems to be more related to an apparently lower response of imports to U.S. output variations after 2000 (see Figure 13, panel (h)).

As discussed before, Hansen's (1992) tests may have low power when the breakpoint is near the end of the sample. We therefore complement the analysis conducted above with Andrews and Kim's (2003) tests, summarized in Table 9. We use the same breakpoint dates as in the aggregate analysis and we focus on the P_c and R_c tests which, as mentioned before, seem to be the ones with better finite-sample properties. As found before, there is no evidence of recent structural changes in the cointegrating vectors linking Mexico's textile and apparel and chemical exports with U.S. production levels of these industries. In contrast, for the cases of nonmetallic mineral products and primary metals, the evidence of structural instability seems to be stronger. Note that, for the former case, this evidence was not found with Hansen's (1992) tests. It is important to mention, however, that the rejections of a stable cointegrating relationship for these two cases do not correspond to evidence concerning the hypothesis we are testing in this paper. Indeed, as can be observed in Figure 13, panels (f) and (g), in both cases the recent breaks appear to be related to sudden decreases in U.S. production levels that were not completely matched by the evolution of imports from Mexico. Similar events seem to be behind the weak evidence of structural break in the paper and printing industry cointegrating vector (Figure 13, panel (d)).

Finally, the tests overwhelmingly reject the null of no structural change in the case of metal products and machinery, suggesting that the links of this industry between Mexico and the U.S. may be weakening. The overall results from this section suggest that this structural break may have occurred even before the end-of-sample breakdown periods tested in Table 9. Indeed, we can observe that, in Table 7, the timing of a possible structural change in this relation is identified by Gregory and Hansen's (1996) tests as April, 2000. This roughly corresponds to the period when the imports from Mexico of this kind of products seem to have reduced their response to movements in U.S. production levels (see Figure 13, panel (h)).

The results of this section suggest that the production-sharing links between Mexico and U.S. manufacturing production have recently weakened only for the case of metal products and machinery. This, however, is an extremely relevant event: this industry contributes with almost one third of Mexico's manufacturing GDP and accounts for roughly 80% of Mexican

exports to the U.S. The recent weakening of the Mexico-U.S. link in this particular sector may very well be behind the differences in the behavior that was observed during 2003 in the aggregate manufacturing production levels of these countries. The fact that, in many cases, China has achieved significant increases in its exports to the U.S. in precisely the kinds of products within this industry in which Mexico has lost market share suggests that the weakening in the cointegrating relationship between Mexico's exports of metal products and machinery to the U.S. and the output levels of this industry may be reflecting a shift of U.S. outsourcing from Mexico to other countries.

Concerning the other industry in which Mexico and the U.S. have achieved important production-sharing schemes (textiles and apparel), the evidence does not suggest a recent breakdown in these links. This, however, masks the negative trend that this industry is exhibiting in both countries (see Figure 13, panel (b)). In this case the relevant situation does not seem to be that U.S. is outsourcing less processes to Mexico, but that the full North American production chain is being affected by the fact that Asian countries have achieved a higher degree of integration in the production of apparel and are exporting finished products to the U.S. without relying on the supply of American-made inputs (see Gereffi, 2000). This may reduce both Mexico's share in U.S. apparel imports and U.S. textile inputs production. In this context, the extension of preferential tariff treatment for apparel produced in countries in the Caribbean Basin that use U.S. textile inputs seems to have been offered by the U.S. in 2000 to give some temporary relief to its own textile industry (Gitli and Arce, 2001). The constraints that China accepted when entering the WTO may also cause this process to become slower.²⁷ In any case, Mexico and the U.S. seem to be strongly attached in the production of an industry that is apparently being fully transferred out of North America as a consequence of higher competition from integrated Asian region production networks.

6. Conclusions

There are several conclusions that can be drawn from the results presented in this paper. First, to the extent that many developing countries have become integrated to the world's capital and trade flows and, as a consequence, have specialized in specific processes within

²⁷ China accepted several conditions on its accession to the WTO that limit the growth of its exports, especially concerning textile and apparel products. These conditions will affect its market access, as compared to that of other exporting countries, especially after the Agreement on Textiles and Clothing comes to an end in January of 2005. Furthermore, up to 2013, WTO members will be able to use special safeguards to face large increases in imports from China. Finally, up to 2016, China will face a special treatment in terms of price comparability to assess the presence of dumping in its exports. The importing country will have the option to use price levels of a

each manufacturing industry, it is possible to expect their business cycles to become synchronized with those of their main trading partners. Mexico-U.S. trade integration experience supports this conclusion. Even before NAFTA was enacted, cyclical movements in the U.S. seem to have been transmitted to Mexico through trade. This treaty, however, increased the incentives to form production-sharing schemes between Mexico and the U.S. and, therefore, induced their bilateral trade to become fundamentally intra-industry in nature. As a consequence, this treaty not only seems to have strengthened the demand-side link between these countries, but appears to have created a more permanent link based on supply-side complementarities.

Second, our results are consistent with the view that China's entry into the international product markets may have important, long-term consequences on the structure of international trade flows. China's accession to the WTO has not only affected Mexico's competitiveness and, therefore, may be having a permanent negative effect on its share in the U.S. import market. It has also provided other industrialized Asian countries with a large supply of unskilled labor that can be exploited to achieve large production complementarities in the region and, through this venue, compete more fiercely with North American product chains.

The previous point leads us to our last conclusion. The impulse that NAFTA gave to Mexico's exports during the mid-nineties seems to be exhausting as a result of the greater competition implied by the entry of other less developed countries into the globalization process. The production-sharing links that this treaty induced seem to be weakening in some industries in which other suppliers are reaping off Mexico's initial comparative advantages. Furthermore, some industries where the strength of these links seems to remain are activities that are apparently being transferred out of the North American region. This may suggest that, as compared to the mid-nineties, Mexico is not currently in such a favorable position to take advantage of the U.S. economy's upswing to attain important increases in its exports and, through this venue, enhance its own economic growth rates.

We must emphasize, however, that most of the evidence supporting a recent weakening of the synchronization between Mexico and U.S. business cycles is based on newly-developed end-of-sample structural break tests. These tests do not allow us to distinguish if the structural break took the form of a downward level shift in Mexico's relative output levels or of a decrease in their elasticity with respect to U.S. output. Furthermore, with the data currently available, traditional tests do not support the structural change hypothesis as strongly. We

third country as benchmark in the evaluation. This will very possibly lead to cases in which dumping will appear to exist when, in fact, Chinese production costs are legitimately lower (see Li, 2002).

therefore cannot discard the possibility that the apparent recent weakening of the links between Mexico and U.S. manufacturing is driven by some high-frequency phenomena as could be, for example, an extraordinarily long lag in Mexico's response to the upturn in U.S. manufacturing, and not necessarily by a true structural break in the long-run links between these sectors. Such a long lag in the response of Mexico's manufacturing to changes in its U.S. counterpart, however, has not been observed since 1996. After that year, Mexico's response to movements in U.S. manufacturing output had been roughly contemporaneous. Obviously, future data will allow us to distinguish more clearly between these possibilities and to have a more complete view of the factors that have driven the differences between Mexico and U.S. manufacturing industries' recent growth patterns.

We must finally note that we purposefully did not make any assessments concerning the possible welfare effects of a weakening of the Mexico-U.S. business cycle synchronization. In order to assess these effects fully, an analysis of the flexibility with which resources within Mexico can be shifted from one sector to another and of the existence of profits that could be shifting from Mexican firms to other parts of the world, is needed. This analysis is beyond the scope of this paper and is therefore left for future research.

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Table 1. Cointegration tests for the logs of manufacturing production in Mexico and U.S.

	Sample			
	Jan 1980-Feb 2004	Jan 1980-Dec 1993	Jan 1986-Dec 1993	Jan 1996-Feb 2004
Engle and Granger (1987)	-2.6713	-1.4448	-1.8371	-2.4946
Campos, Ericsson and Hendry (1996)	-2.7605	-1.7714	-2.7773	-5.4757 ***
Johansen (1991, 1995)				
Test statistic	9.7313	6.8371	6.5417	18.8455 **
Long-run elasticity	-----	-----	-----	0.9743 **
(s.e.)	-----	-----	-----	(0.0524)
Adjustment Coefficients				
Mexico	-----	-----	-----	-0.2638 ***
(s.e.)	-----	-----	-----	(0.0699)
US	-----	-----	-----	-0.0422
(s.e.)	-----	-----	-----	(0.0386)
Gregory and Hansen (1996)				
ADF^*	-4.4279	-4.2527	-3.7103	-5.6125 ***
Estimated breakpoint	Jul-1989	May-1989	Oct-1989	Nov-2002
Z_t^*	-3.6121	-3.1797	-3.4736	-5.5566 ***
Estimated breakpoint	Jan-1989	Jan-1989	Jan-1990	Nov-2002
Z_α^*	-24.1645	-18.3535	-18.7771	-45.0226 *
Estimated breakpoint	Jan-1989	Jan-1989	Oct-1989	Nov-2002

Notes: For the cointegration tests, one, two and three asterisks denote rejection of the null hypothesis at 10%, 5% and 1%, respectively. For coefficient estimates, one, two and three asterisks denote rejection of the null that the coefficient equals zero at 10%, 5% and 1% significance level, respectively. With the exception of Johansen's (1991, 1995) tests, in all cases the number of lags included in the test regressions were chosen with a downward testing procedure, starting with a maximum of 12 lags and eliminating the largest lags until the last lag was significant at the 5% level using t (or F) tests. The results were similar to those obtained when the lag order was chosen with Akaike's Information Criterion. For Johansen's (1991, 1995) tests, linear deterministic trends were allowed in the data, and the Akaike and Schwartz Information Criteria were used to choose the maximum lag order. Long run elasticities and adjustment coefficients estimated with Johansen's (1991, 1995) procedure are only reported for the cases when the test rejects the null of no cointegration at least at a 10% significance level. Engle and Granger's (1987) and Campos, Ericsson and Hendry's (1996) test statistics were compared to critical values computed using MacKinnon's (1991) response surface regressions. Gregory and Hansen's (1996) tests use their model 4, which allows for both a level shift and a slope change at the time of the structural break, as alternative hypothesis. The results are very similar to those obtained if we used model 3, which allows for a deterministic trend and a level shift, but not a change in the slope, in the cointegrating vector.

Table 2. OLS and FM-OLS estimates of the cointegrating vector for manufacturing in Mexico and U.S. and stability tests

Cointegrating vector estimates				
	OLS	FM-OLS		
Constant	-0.3080	-0.1083		
(s.e.)	(0.0846)	(0.1644)		
Slope	1.1069	1.0642 ***		
(s.e.)	(0.0181)	(0.0351)		
Stability tests				
		Significance levels		
		10%	5%	1%
<hr/>				
Hansen (1992) stability tests				
L_c	0.171	0.36	0.47	0.72
$MeanF$	1.510	3.73	4.48	6.83
$SupF$	4.558	11.20	12.90	16.40
Hao and Inder (1996)				
CUSUM diagnostic test	0.738	0.77	0.83	0.97

Notes: The sample goes from January, 1996 to February, 2004. For the stability tests, one, two and three asterisks denote rejection of the null hypothesis at 10%, 5% and 1%, respectively. For the coefficient estimates from the FM-OLS regression, one, two and three asterisks denote rejection of the null that the coefficient equals zero at 10%, 5% and 1% significance level, respectively. Given the dependence of their distribution on nuisance parameters, no significance levels are reported for the OLS estimates. See main text for details on the construction of the stability tests.

Table 3. Andrews-Kim (2003) end-of-sample cointegration breakdown tests
(*p* values)

	Breakpoint	
	Jan-2002	Jan-2003
P_a	0.1489	0.0000
P_b	0.0426	0.0000
P_c	0.0000	0.0000
R_a	0.3617	0.0000
R_b	0.2128	0.0000
R_c	0.0851	0.0000

Notes: The sample is January, 1996 to February, 2004. All tests are based on OLS residuals. See main text for details on the construction of the cointegration breakdown tests.

Table 4. Dynamic equation for Mexico's manufacturing production index

Sample: 1996:01 2002:12		
Equation: $\Delta MEX_{t-1} = \beta(1)\Delta MEX_{t-4} + \beta(2)\Delta US_{t-12} + \beta(3)ECM_{t-1}$		
$ECM_t = MEX_t - \beta(4)US_t$		
	Coefficient	Std. Error
$\beta(1)$	0.1433 *	0.0851
$\beta(2)$	-0.2936 *	0.1502
$\beta(3)$	-0.4189 ***	0.0676
$\beta(4)$	1.0431 ***	0.0006
Diagnostic tests		
R^2	0.3836	
Durbin-Watson	1.8958	
F (constraints on general model)	1.1252	($p = 0.3489$)
Chow Forecast X^2 (2003:1-2004:2)	35.9414 ***	($p = 0.0011$)
Jarque-Bera	3.1056	($p = 0.2117$)
LM (12)	7.1882	($p = 0.8449$)
ARCH (1)	3.2311 *	($p = 0.0722$)
White heteroskedasticity test	15.9458	($p = 0.1937$)

Notes: The sample goes from January, 1996 to December, 2002. MEX_t and US_t denote logs of Mexico and U.S. manufacturing production levels, respectively. For coefficient estimates, one, two and three asterisks denote rejection of the null that the coefficient equals zero at 10%, 5% and 1% significance level, respectively. For the diagnostic tests, one, two and three asterisks denote rejection of the relevant null hypothesis at 10%, 5% and 1%, respectively. p -values for these tests are also provided. See main text for details on the specification strategy for this equation.

Table 5. Summary Statistics of Mexico's manufacturing divisions

	Share in Mexico's manufacturing GDP (percent) ^{3/}				
	1999	2000	2001	2002	2003
Food, beverage and tobacco products	25.17	24.48	26.05	26.68	27.49
Textile, apparel and leather products	8.67	8.55	8.13	7.71	7.15
Wood products	2.79	2.71	2.63	2.52	2.57
Paper and printing	4.75	4.57	4.54	4.49	4.50
Chemical products ^{1/}	15.43	14.91	14.92	14.99	15.53
Nonmetallic mineral products	6.90	6.72	6.87	7.18	7.37
Primary metal	5.13	4.95	4.78	4.87	5.13
Metal products and machinery ^{2/}	31.16	33.11	32.07	31.57	30.27

	Share in overall US manufacturing imports from Mexico (percent)				
	1999	2000	2001	2002	2003
Food, beverage and tobacco products	2.80	2.57	2.57	2.93	3.16
Textile, apparel and leather products	10.21	9.26	8.80	8.52	7.97
Wood products	0.36	0.26	0.22	0.21	0.20
Paper and printing	0.64	0.60	0.61	0.67	0.75
Chemical products ^{1/}	4.45	4.39	4.05	3.98	4.17
Nonmetallic mineral products	1.54	1.39	1.41	1.48	1.52
Primary metal	2.36	2.12	1.98	2.08	2.04
Metal products and machinery ^{2/}	77.64	79.42	80.37	80.13	80.19

	Ratio of US imports to value added in Mexico (percent) ^{4/}				
	1999	2000	2001	2002	2003
Food, beverage and tobacco products	10.40	10.12	8.80	9.70	10.92
Textile, apparel and leather products	124.42	123.17	114.93	116.19	125.86
Wood products	13.59	10.64	8.44	8.56	8.85
Paper and printing	15.18	15.44	14.92	16.79	20.29
Chemical products ^{1/}	28.91	30.98	26.85	26.00	27.34
Nonmetallic mineral products	22.78	21.43	20.36	20.84	22.48
Primary metal	48.07	46.30	47.49	50.70	47.29
Metal products and machinery ^{2/}	230.36	246.44	250.00	257.74	288.47

^{1/} includes petroleum, coal, rubber and plastic products.

^{2/} includes computer, electrical, electronic and transport equipment and parts.

^{3/} is computed with 1993 constant pesos GDP.

^{4/} is computed using nominal GDP.

Notes: Data on Mexico's manufacturing GDP comes from INEGI. Data on U.S. imports of Mexico's manufacturing products is from the U.S. Census Bureau. Data on U.S. imports by product was aggregated to match Mexico's classification of manufacturing divisions. Miscellaneous products included in the category of "other manufactures" were dropped from the analysis.

Table 6. Mexico's and China's shares in U.S. manufacturing imports
(Percentages)

		1999	2000	2001	2002	2003
Total						
	Mexico	10.90	11.49	12.04	11.85	11.02
	China	9.08	9.64	10.53	12.57	14.73
Food, beverage and tobacco						
	Mexico	8.85	9.76	9.24	9.95	9.49
	China	0.91	0.93	0.95	1.07	1.17
Textile, apparel and leather products						
	Mexico	12.95	12.82	12.17	11.65	10.21
	China	12.57	12.76	13.48	14.94	17.16
Wood products						
	Mexico	3.43	3.43	3.01	2.63	2.06
	China	7.86	9.66	11.08	12.34	12.31
Paper and printing						
	Mexico	2.86	2.68	2.90	3.32	3.41
	China	2.95	3.30	3.60	4.73	5.88
Chemicals ^{1/}						
	Mexico	2.88	2.49	2.18	2.37	2.19
	China	2.63	2.37	2.44	3.21	3.26
Nonmetallic mineral products						
	Mexico	5.58	5.38	5.85	5.52	5.29
	China	7.53	8.08	9.32	9.68	9.85
Primary metals						
	Mexico	8.08	7.91	8.40	9.33	9.18
	China	6.97	7.49	8.92	10.97	12.97
Metal products and machinery ^{2/}						
	Mexico	13.29	14.35	15.63	15.25	14.47
	China	5.51	6.33	7.00	9.15	11.64

^{1/} includes petroleum, coal, rubber and plastic products.

^{2/} includes computer, electrical, electronic and transport equipment and parts.

Notes: Data by product category comes from the U.S. Census Bureau. The figures were aggregated to match Mexico's classification of manufacturing divisions. Miscellaneous products included in the category of "other manufactures" were dropped from the analysis.

Table 7. Cointegration tests for U.S. imports from Mexico and U.S. production by manufacturing division

	Industry							
	Food, beverage and tobacco products	Textile, apparel and leather	Wood products	Paper and printing	Chemical products	Nonmetallic mineral products	Primary metal	Metal products and machinery
Engle and Granger (1987)	-4.9420 ***	-2.3712	-2.7505	-2.3143	-5.7032 ***	-2.2223	-4.6547 ***	-1.8079
Campos, Ericsson and Hendry (1996)	-2.6968	-3.2630	-2.4229	-5.2436 ***	-3.1617	-2.6708	-5.6258 ***	-0.6694
Johansen (1991, 1995)								
Test statistic	16.3802	25.6696 ***	17.8040 *	21.0790 **	29.9451 ***	19.9261 **	27.4458 ***	15.8917
Long-run elasticity	-----	3.2428 *	0.5044	1.1724 *	5.6989 ***	1.7949 ***	1.0971 ***	-----
(s.e.)	-----	(0.3068)	(0.3986)	(0.3789)	(0.8569)	(0.2509)	(0.1700)	-----
Adjustment Coefficients								
Mexico	-----	-0.1065 ***	-0.2634 ***	-0.4429 ***	-0.4620 ***	-0.3474 ***	-0.9945 ***	-----
(s.e.)	-----	(0.0277)	(0.0631)	(0.0980)	(0.0826)	(0.0836)	(0.1968)	-----
US	-----	0.0244	-0.0333 *	-0.0014	0.0086	-0.0195	-0.1293 **	-----
(s.e.)	-----	(0.0137)	(0.0167)	(0.0147)	(0.0060)	(0.0332)	(0.0514)	-----
Gregory and Hansen (1996)								
<i>ADF</i> *	-5.0310 **	-5.1673 **	-4.6985 *	-4.7905 *	-8.2315 ***	-5.4778 ***	-5.3774 **	-7.6418 ***
Estimated breakpoint	Jul-1999	Sep-1999	Dec-2000	Jul-2000	Dec-1997	Jun-1999	Oct-1998	Apr-2000
Z_t *	-5.1140 **	-6.4512 ***	-4.5925	-4.8154 *	-8.4005 ***	-5.6857 ***	-4.7971 *	-7.6814 ***
Estimated breakpoint	Jul-1999	Aug-2000	Dec-2000	Jul-2000	Jan-1998	Aug-1999	Aug-2002	Apr-2000
Z_a *	-42.1132 *	-36.6267	-31.5169	-38.5683	-81.8790 ***	-47.7584 **	-36.4091	-70.6512 ***
Estimated breakpoint	Jul-1999	Mar-1998	Jan-2001	Jul-2000	Jan-1998	Aug-1999	Apr-2000	Jan-2000

Notes: The sample goes from January, 1996 to February, 2004. For the cointegration tests, one, two and three asterisks denote rejection of the null hypothesis at 10%, 5% and 1%, respectively. For coefficient estimates, one, two and three asterisks denote rejection of the null that the coefficient equals zero at 10%, 5% and 1% significance level, respectively. With the exception of Johansen's (1991, 1995) tests, in all cases the number of lags included in the test regressions were chosen with a downward testing procedure, starting with a maximum of 12 lags and eliminating the largest lags until the last lag was significant at the 5% level using *t* (or *F*) tests. The results were similar to those obtained when the lag order was chosen with Akaike's Information Criterion. For Johansen's (1991, 1995) tests, the Akaike and Schwartz Information Criteria were used to choose both the maximum lag order and the model specification. Long run elasticities and adjustment coefficients estimated with Johansen's (1991, 1995) procedure are only reported for the cases when the test rejects the null of no cointegration at least at a 10% significance level. Engle and Granger's (1987) and Campos, Ericsson and Hendry's (1996) test statistics were compared to critical values computed using MacKinnon's (1991) response surface regressions. With the exception of Gregory and Hansen's (1996) tests, all tests allow for a deterministic trend in the cointegrating relationship. Gregory and Hansen's (1996) tests use their model 4, which allows for both a level shift and a slope change at the time of the structural change, as alternative hypothesis. The results are very similar to those obtained if we used model 3, which allows for a deterministic trend and a level shift, but not a change in the slope, in the cointegrating vector.

Table 8. Industry-specific FM-OLS cointegrating vector estimates and stability tests

	Food, beverage and tobacco products	Textile, apparel and leather products	Wood products	Paper and printing	Chemical products	Nonmetallic mineral products	Primary metal	Metal products and machinery
Cointegrating vector estimates (FM-OLS)								
Constant (s.e.)	15.0152 *** (2.8408)	4.8109 *** (1.2483)	13.7592 *** (1.7576)	9.3553 *** (2.5098)	-4.9968 (3.9697)	10.5682 *** (1.1888)	11.7266 *** (1.1365)	15.9998 *** (0.5088)
Slope (s.e.)	0.8654 (0.6160)	3.2457 *** (0.2666)	0.7714 ** (0.3797)	1.7470 *** (0.5415)	5.3149 *** (0.8656)	1.6680 *** (0.2574)	1.5352 *** (0.2443)	1.3135 *** (0.1126)
Trend (s.e.)	0.0049 *** (0.0004)	0.0190 *** (0.0012)	-0.0068 *** (0.0006)	0.0080 *** (0.0008)	0.0003 (0.0011)	0.0053 *** (0.0003)	0.0057 *** (0.0007)	0.0013 * (0.0007)
Hansen (1992) stability tests								
L_c	0.4188	0.1351	0.3165	0.1649	0.2824	0.2601	0.5895 *	0.2072
$MeanF$	2.9710	2.5407	5.1164 *	2.6315	4.4481	3.5549	7.5590 **	5.9696 *
$SupF$	6.8942	8.3004	7.4825	5.0156	16.4452 **	7.5974	12.8387	12.2998

Notes: The sample is January, 1996 to February, 2004. For the coefficient estimates, one, two and three asterisks denote rejection of the null that the coefficient equals zero at 10%, 5% and 1% significance level, respectively. For the stability tests, one, two and three asterisks denote rejection of the null hypothesis at 10%, 5% and 1%, respectively. See main text for details on the construction of the stability tests.

Table 9. Andrews and Kim's (2003) cointegration breakdown tests
(*p* values)

Breakpoint			Breakpoint		
	Jan-2002	Jan-2003		Jan-2002	Jan-2003
Food, beverage and tobacco			Textile, apparel and leather		
P_a	0.1277	0.7606	P_a	0.0426	0.1268
P_b	1.0000	0.9296	P_b	0.2340	0.1690
P_c	1.0000	0.9577	P_c	0.2553	0.1549
R_a	0.0000	0.4085	R_a	0.0426	0.1972
R_b	0.3830	0.5775	R_b	0.5532	0.2535
R_c	0.8298	0.8169	R_c	0.6170	0.2394
Wood products			Paper and printing		
P_a	0.4894	0.2254	P_a	0.2979	0.0423
P_b	0.5745	0.2394	P_b	0.4468	0.0423
P_c	0.3404	0.2817	P_c	0.2340	0.0986
R_a	0.3830	0.1972	R_a	0.2766	0.0000
R_b	0.5106	0.1972	R_b	0.4043	0.0141
R_c	0.4681	0.2958	R_c	0.4681	0.1972
Chemical products			Nonmetallic mineral products		
P_a	0.1064	0.4648	P_a	0.0000	0.0000
P_b	0.6809	0.6479	P_b	0.0426	0.0000
P_c	0.7234	0.7465	P_c	0.0000	0.0423
R_a	0.0638	0.2676	R_a	0.0000	0.0000
R_b	0.3830	0.3662	R_b	0.0000	0.0000
R_c	0.5745	0.4648	R_c	0.0000	0.0000
Primary metal			Metal products and machinery		
P_a	0.0000	0.2535	P_a	0.0000	0.0000
P_b	0.0000	0.2958	P_b	0.0426	0.0000
P_c	0.0000	0.3239	P_c	0.0000	0.0000
R_a	0.4681	0.4930	R_a	0.0000	0.0000
R_b	0.6596	0.4225	R_b	0.0213	0.0000
R_c	0.2340	0.3803	R_c	0.0426	0.0000

Notes: The sample is January, 1996 to February, 2004. All tests are based on OLS residuals. See main text for details.

Figure 1
Mexico and U.S. manufacturing production indexes
(Logarithms)

Sources: Banco de México and Board of Governors of the Federal Reserve System.

Figure 2

Mexico and U.S. manufacturing production indexes
(Percentage annual growth rates)

Sources: Banco de México and Board of Governors of the Federal Reserve System.

Figure 3

Mexico and China's share in U.S. non-oil imports

Source: U.S. Census Bureau. Figures were seasonally adjusted by the authors using the X12-ARIMA program.

Figure 4

Number of active *maquiladora* plants in Mexico

Source: INEGI.

Figure 5

Coherence between Mexico and U.S. manufacturing production indexes

Figure 6

**Phase lead of Mexico's manufacturing production index with respect to U.S.
manufacturing production index**

(Quarters by which Mexico's cycle for each frequency leads U.S. cycle)

Figure 7

Business-cycle components of Mexico and U.S. manufacturing production indexes

(Log difference with respect to the trend)

Figure 8

Trend components of Mexico and U.S. log manufacturing production indexes

Figure 9
FM-OLS CUSUM Test

Figure 10
F Statistic Sequence

Figure 11
Estimates of β

Figure 12

**Observed vs. forecasted Mexico manufacturing production index
(Logarithm)**

Figure 13

Notes: Imports were aggregated from product-level data on imports from the U.S. Census Bureau, deflated with specific producer price indexes by product category. Production indexes were aggregated to be consistent with Mexico's classification from data published by the Board of Governors of the Federal Reserve System.